UNIVERSITAT DE BARCELONA

Essays on the Effect of International Migration on Local Labor Markets: Evidence from Europe

Stefano Fusaro

PhD in Economics

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Chapter 1

Introduction

In the two decades before the outbreak of the COVID-19 pandemic, the phenomenon of international migration experienced a rapid and consistent growth. By the end of 2020, the figures indicate that the worldwide stock of people living in a country different from the one in which they were born reached 281 million individuals (United Nations, 2020). Just to have an idea on the size of the global immigrant population, if all of them lived in only one country, this would be the third most populous one, exceeded only by China, India and the United States. In terms of the dynamics of international migration, the figures indicate that between 2000 and 2010 around 48 million people migrated globally. A further increase occurred between 2010 and 2020, with around 60 million international migrants around the world. The vast majority of these cross borders movements are due to family reunifications, or to job-related reasons. Alongside, humanitarian crises, mainly caused by the disruption of wars of by climate change, also played a crucial role. To this extent, the United Nations indicate that in the two decades between 2000 and 2020, the number of refugees or asylum seekers grew by about 17 million units. In terms of stocks, the estimates report that the number of refugees and asylum seekers in 2020 was around 34 million, nearly doubled since the year 2000.

As for the geographical distribution of the immigrant population, Europe is by far the continent hosting the largest number of worldwide foreign-born individuals, with almost 87 million people. It is then followed by America, with about 59 million and North Africa and Asia, hosting 50 million migrants each. Moreover, about the 63% of all migrants originates from middle-income countries, while only 13% of them from a low-income one. This, however, is not surprising, particularly if considering that migration is a fairly costly operation and is therefore not affordable to everyone.

Another peculiar feature of the global immigrant population is its gender composition. The available data indicate that in 2020 about half of all international migrants were women. Interestingly, the trend is confirmed even considering refugees or asylum seekers. This is a fundamental aspect to take into account, as immigrant women often play a crucial role in shaping the socio-political context of the countries of destination, and induce a shift in their cultural norms[1]. This is particularly valid in the case of Italy and the other EU countries analyzed in the different chapters of this Ph.D. thesis - namely Greece, Ireland, Portugal and Spain - that see how their female population is still often disadvantaged with regard to

[1]For instance by substituting native women in the household production services, with a positive effect on their labor supply, both at the intensive and extensive margins.

economic opportunities.

In terms of the age distribution of the worldwide immigrant population, despite the presence of a significant cross-country variation, overall an important share of it is composed by working-age individuals. This is fundamental because it implies that immigrants represent a significant component of the labor force of the destination country. Because of this reason, the issue of immigration has recently assumed a central position in the public and political debate of the host countries, with an upsurge of a widespread anti-immigration sentiment.

In this context, the analysis of the economic consequences of immigration has become increasingly important and interesting, also among scholars. To this extent, despite the fact that some analyses were carried out already in the early 1980s (Grossman, 1982; Borjas, 1983) and 1990s (Card, 1990; Altonji and Card, 1991; Hunt, 1992), it is only recently that the "Economics of Immigration" has assumed a central stage in the academic debate (see Borjas, 2014; Card and Peri, 2016).

The existing studies focus on different dimensions of the potential impact of immigration, such as economic growth (Boubtane et al., 2016; Borjas, 2019), public finances (Preston, 2014; Bettin and Sacchi, 2020), housing prices (Saiz, 2003, 2007), native education (Hunt, 2017; Brunello et al., 2020), innovation (Hunt and Gauthier-Loiselle, 2010; Bratti and Conti, 2018), trade (Bratti et al., 2014), crime (Mastrobuoni and Pinotti, 2015; Fasani, 2018) and electoral outcomes (Mayda et al., 2018). However, most of the interest has been on the impact of immigration on the labour market outcomes of the native population, like employment and wages (Angrist and Kugler, 2003; Dustmann et al., 2008; Borjas, 2017; Borjas and Monras, 2017; Llull, 2018*a*; Clemens and Hunt, 2019; Peri and Yasenov, 2019; Monras, 2020). To this extent, the empirical analysis of the impact of immigration allows to identify an extremely policy-relevant parameter: the (wage or employment) elasticity of labor supply.

Despite the large amount of studies carried out (see Dustmann et al., 2016), there is an on going debate among scholars on what is the actual impact (if any) that immigrant exert on the labor markets outcomes of natives, and the consensus is apparently far from being reached[2]. This involves different aspects of the economics of immigration. One of the main questions is relative to which is the more appropriate empirical methodology to use in order to identify the "true" impact of immigration. To simplify, all the existing literature can be reduced to two main approaches. The first can be defined as the "spatial approach"[3]. In this context, labor markets are defined based on geographical units and to identify the effect of interest, the existing studies typically exploit the uneven spatial distribution of immigrants, by comparing locations with a low immigrant penetration with others with a higher one. The main limitation of this approach is that local labor markets are not closed economies, and therefore natives may respond to the labor supply shifts induced by immigrants by moving to regions offering better economic opportunities and less affected by immigration shocks. If this is the case, then the estimation of the effect of immigration obtained with this approach might be contaminated by the native relocation responses and therefore not be reliable.

The second is the "national approach", pioneered by Borjas (2003). This setting exploits the different intensity in the immigration shock across education-experience groups, typically defined as skill-cells. In other words, labor markets are not defined based on geography, but

[2]However, the meta-analyses of Longhi et al. (2010) and Aubry et al. (2021) indicate that both the wage and employment responses to immigration are basically null.

[3]Notable examples of previous studies using this approach are Card (2009*a*) and Dustmann et al. (2013).

rather based on the combination between education and work experience (which are usually observed in the data). However, this approach also presents some important drawbacks. The first is the so-called immigrant skill-downgrading. In other words, particularly in the first years upon arrival, immigrants tend to be employed in occupations for which they are over-educated. This relies on the less-than-perfect substitutability of human capital that individuals often need to face when migrating abroad. If this occurs, then when assigning immigrants to a specific skill-cell just based on their (observed) education and experience, the risk of misallocate them is highly likely. An analogous problem may surge if some natives are over-educated. Moreover, the national skill-cells approach just identifies the own-group effects - that is, the impact that immigrants exert on equally educated and experienced natives - but not the total effect of immigration, which, however, is the most relevant from a policy point of view (see Dustmann et al., 2016).

Both approaches previously indicated present some important statistical difficulties, that, if not thoroughly addressed, may threaten the identification of the effect of interest. The first, and most relevant one, is that immigrants are almost never randomly distributed across space, but rather they typically settle in regions offering better economic opportunities. This is problematic as it generates a spurious correlation between local economic conditions and immigrant shocks. More precisely, on the one hand, immigrants affect the local wage structure by inducing a shift in the labor demand. On the other, however, the local wage dynamics also affect the spatial distribution of immigrants, as these are attracted to regions that are thriving. The combination of these two aspects creates a reverse causality bias that threatens the identification of the effects of interest.

There are different ways to address this problem. The first is to use a so-called "natural experiment", that consists in an influx of immigrants that depends on factors that are (arguably) exogenous with respect to the economic conditions of the labor markets of destination. Notable examples exploited by the previous literature are the flow of *Marielitos* from Cuba to Miami occurred in 1980 (Card, 1990; Borjas, 2017; Peri and Yasenov, 2019); the flow of Jewish *Émigrés* to Israel occurred in the 1990s after the collapse of the Soviet Union (Friedberg, 2001; Borjas and Monras, 2017; Clemens and Hunt, 2019); the flow of French *Pieds Noirs* repatriated to France after the Algerian independence of 1962 (Hunt, 1992; Borjas and Monras, 2017; Clemens and Hunt, 2019); the flow of people from the former Yugoslavia to Europe occurred after the outbreak of the Balkan war (Angrist and Kugler, 2003; Borjas and Monras, 2017; Clemens and Hunt, 2019); and finally the flow of Portuguese *Retornados* from Angola and Mozambique in the 1970s (Carrington and De Lima, 1996; Mäkelä, 2017). Another alternative is to exploit a policy change that exogenously allocate immigrants (or, more often, refugees) across the different regions of the country of destination (Foged and Peri, 2016; Gamalerio, 2018; Tumen, 2018). However, such events are fairly rare and, when they occur, it is sometimes difficult to have the information required to run a properly defined empirical analysis that exploits them.

The second way to address the reverse causality bias is to apply an instrumental variable approach. In the migration literature, the most common instrument used is the so-called "shift-share" variable. The rationale behind this type of variable relies upon the fact that, besides economic opportunities, another important aspect that immigrants take into account when deciding where to locate is the presence of "older" immigrants coming from the same countries of origin (Bartel, 1989; Munshi, 2003). Indeed, this can generate a sort of network effect that somehow attenuates the psychological and pecuniary costs of migrating. From the empirical point of view, this phenomenon implies that the historical settlements of immigrants are a good predictor of the current ones and therefore are a valid instrument for them. This variable has been used for the first time in the migration literature by Altonji and Card (1991), successively refined by Card (2001), and extensively used since then. Despite its extreme popularity[4], this variable has recently received a considerable amount of criticism (Adão et al., 2019; Jaeger et al., 2019; Borusyak et al., 2020; Goldsmith-Pinkham et al., 2020). More precisely, in presence of persistent labor demand shocks that affect both the native labor market outcome and immigrants' location decision, the historical settlement of the immigrant population may not be fully exogenous to the economic conditions of the local labor markets. This can be further magnified in presence of a countries of origin mix of the immigrant population particularly stable over time. If this occur, then the exclusion restriction would likely be violated, and therefore the instrument not be valid. There are different ways to address these issues. The first one is to choose a sufficiently lagged base-year that predates consistent immigrant shocks, as done, for instance, by Monras (2020).

The second is instead to use alternative instruments. On the one hand, is possible to use an instrument based on the distance between the region of residence in the destination country and the origin country (e.g. Angrist and Kugler, 2003; Ottaviano and Peri, 2005, 2006), or the gateways through which immigrants are assumed to enter the country of destination (Mocetti and Porello, 2010)[5]. Another interesting option is to build an instrument based on the push-factors of the countries of origin, as done, for instance, by Boustan (2010) and Heepyung (2019). Finally, it is also possible to combine the last two approaches and build a distance and push-factors-based instrument, as in Sparber and Zavodny (2020). The empirical exercises in the thesis use the IV approach with the aforementioned type of instruments to identify the impact of immigration.

As for the existing literature, it is important to underline that it mostly provides evidence on countries characterized by some important peculiarities (see Table 1 in Dustmann et al., 2016). First, in these countries immigration is not a recent phenomenon, but they typically have a long tradition of immigration. This is crucial because it affects the way in which (local) labor markets absorb the supply shifts induced by immigrants. Second, they are characterized by an extent of wage rigidity and employment protection fairly different to that in the countries under analysis in this thesis, where unions also have a more central role in the wage bargaining mechanism. Third, the skill distribution of the immigrant population is fairly different to those in the countries under analysis. Specifically, two of the countries mostly analyzed in the existing literature, namely the U.K. and the U.S., alongside low-skilled immigrants, typically attract also high-skilled ones (Manacorda et al., 2012; Peri, 2016). This is an important aspect to take into account, because the way in which the labor markets of destination absorb the inflows of immigrants also depends on their skills. In addition, high-skilled immigrants may exert a positive impact on innovation and productivity, generate knowledge spillovers and boost the creation of new firms, with all the (typically positive) consequences that this entails for the receiving economies (Moretti, 2012).

The present thesis fills this gap by providing evidences for the set of European countries previously mentioned, that are characterized by some common peculiarities that render the chapters of this thesis particularly interesting. First, in contrast with the the countries

[4] For more details, see Jaeger et al. (2019).

[5] To this extent, the higher is the distance between origin and destination countries, the lower is expected to be the share of immigrants from each specific origin. In addition, the distance arguably exogenous with respect to the local economic condition of the country of destination.

typically analyzed in the previous literature, they only relatively recently changed their role from immigrant-sending to immigrant-receiving countries. Indeed, in all cases the first consistent migratory inflows only started few decades ago and considerably increased since then. Related to this, another interesting peculiarity is that in all the countries under analysis the immigrant population is fairly heterogeneous, particularly in terms of its countries of origin mix. This aspect is not negligible, as there seems to be a relationship between the ethnic composition of the immigrant population and its economic performance in the countries of destination (Alesina and La Ferrara, 2005; Ottaviano and Peri, 2006). Moreover, in the countries under analysis the skill distribution of the immigrants population is typically skewed towards the lower tail of the skill-spectrum (with the exception of Ireland, that also attracts high-skilled individuals).

Second, another important specificity of the countries under analysis in the fairly high degree of employment protection and downward wage rigidity, as well as the strong role played by unions in the wage bargaining process, which. in addition, is usually highly centralized (D'Amuri and Peri, 2014).

Third, alongside the increase in the migratory inflows, these were also the countries mostly affected the the recent economic downturn caused the the Global Financial Crisis and by the subsequent Sovereign debt Crisis. To this extent, they were defined in a derogatory way as the *PIIGS* countries[6].

Fourth, these countries are characterized by large and persistent regional disparities that affect different dimensions of economic activity, such as innovation, productivity, human capital accumulation, employment rates (Fonseca and Fratesi, 2017), as well as the spatial distribution of the immigrant population.

Fifth, another important specificity of the countries under analysis are the peculiar social and cultural norms that, among other things, imply a relatively low female participation rate and a still significant gender wage gap.

Finally, all countries have recently experienced a resurgence of the anti-immigration rhetoric, that runs parallel to the outbreak of an anti-European sentiment, and that is reflected by the creation of typically right-wing political parties[7] whose main objective is to reduce as much as possible the inflows of immigrants.

Given the institutional and economic specificities of the countries under analysis, the general hypothesis of the thesis is that immigrant inflows do not exert a negative impact on both wages and employment of the native population, as instead often claimed by some political actors. This is because immigrants, due to their peculiar work-related characteristics, are not close substitute to native workers, not even to those with comparable skills. This is crucial, because it implies that immigrants do not represent an economic threat for the native-born workers. To this extent, the native education responses to immigration are expected to be negative, implying that, in the countries under analysis, natives do not need to invest in human capital in order to avoid immigrants' competition once in the labor market.

As for its structure, besides introduction and conclusions, the thesis is composed by three empirical chapters that investigate the impact of immigration on three different outcomes of the native population, namely employment, wages and human capital accumulation. In the three chapters of this thesis, I implement a spatial approach. The rationale behind this

[6]Which is an acronym derived from the first letters of these countries names (i.e. Portugal, Ireland, Italy, Greece and Spain).

[7]These are Golden Dawn in Greece, the National Party in Ireland, Lega in Italy, Vox in Spain and Chega in Portugal.

choice depend on three main reasons. The first is that this allows to identify the overall effect of immigration, which, as just indicated, is the more relevant effect, particularly in terms of the policy implications that it entails. The second is that in this way it is possible to exploit empirically the uneven geographical distribution of the immigrant populations in the countries under analysis as well as the large spatial disparities that involve different aspects of economic activity, such as the employment structure, the education distribution of the population and the sectoral composition. Finally, the third is that in the countries under analysis the relocation response to immigration shocks of the native population is basically negligible or, in some cases, even positive (that is, some natives tend to cluster in regions characterized by the higher presence of immigrants).

Also, in all chapters, I address the non-random immigrant settlements by means of the canonical shift-share instrument, which combines the spatial variation in the historical settlement of immigrants by countries of origin with the national levels stocks of the immigrant population by countries of origin in the years under analysis. To address the criticisms that this type of instruments has recently received, in the three chapters of the thesis I perform several empirical exercises to test their validity. Complementarily, I also use the alternative instruments to check the robustness of the results. Specifically, the results of the second and third chapter are robust to a version of the distance-to-gateways instrument proposed by Mocetti and Porello (2010). Similarly, in the fourth chapter I use a version of the push-factors and distance-based instrument built on Sparber and Zavodny (2020).

More in detail, the second and third chapter analyze the impact of immigration on the native labor market outcomes in Italy in the period following the recent economic crisis that severely hit many European economies. The interest of studying the Italian context is attributable to several aspects. First, Italy has only recently started to be an immigrant-receiving country. To this extent, Italy is nowadays one of the most popular destination countries for migrants coming from both Africa and Central and Eastern Europe. Second, the Italian labor market is characterized by a strong degree of wage rigidity and employment protection. This implies that the labor supply shift induced by immigrants may be absorbed through changes in the employment levels, more than on the wage structure.

Specifically, the second chapter of the thesis, *Immigration and Native Employment. Evidence from Italian Provinces in the Aftermath of the Great Recession*, exploits the variability in the incidence of recent immigration inflows and the change in native employment in the Italian provinces to shed light on the impact of immigration on employment in rigid local labor markets. The study focuses on the period that followed the financial and sovereign debt crises, which strongly hit the labor markets of the Italian provinces. The results reveal a negligible overall impact of immigration on provincial employment which, however, hides differentiated impacts for different groups of natives. Employment responses to immigration shocks vary greatly depending on the skills and gender of the natives. After three rounds of revision with three referees, this chapter has been published as article in the peer-reviewed journal *Papers in Regional Science*[8].

If the second chapter analyzes the impact of immigration on "quantities" (i.e. employment), the third one focuses instead on the effects on "prices" (i.e. wages), which is the other important element that immigrant inflows can influence. In this regard, this chapter, titled *On the Heterogeneous Impact of Immigrants on the Distribution of Native Wages. Evidence from Recent Immigrants in Italian Provinces*, provides new evidence on the extent to which

[8]More details are given at the beginning of the chapter.

immigrants affect the wage structure of the local labor markets of the host countries by proposing a methodology that combines the assessment of the impact of immigration along the native wage distribution, with a two-steps procedure that controls for changes in the composition of the native workforce, and for the endogenous allocation of immigrants across local labor markets. As previously mentioned, the analysis is carried out for Italy, which is a peculiar country as is characterized by relatively rigid product and labor markets and by well-known regional disparities, during a period of time dominated by the coexistence of the economic downturn and by the substantial increase of the migratory inflows of low-skilled individuals. The results contradict the simplistic belief that immigrants are indistinctly responsible for the decrease in native wages, and highlight instead two interesting facts. First, in line with the existing literature, foreign-born workers do not affect significantly the native average wages. Second, in terms of the impact along the native wage distribution, the effect that immigrants exert is non-negative, not even for those natives located in the lower part of the wage distribution (which, in principle, are more similar in terms of job characteristics to immigrants, and therefore are expected experience the larger wage losses). If anything, the estimates identify instead a positive impact in the upper part of the wage distribution, which is particularly valid in the case of native women residing in the Northern provinces. However, in the more demanding specifications (that is, those with province-specific trends) this positive effect is only marginally significant. An early version of this chapter has been presented in the workshop on *International and Internal Migration: Challenges and Opportunities in Europe* held at the GSSI in L'Aquila (Italy) on January 2020.

The fourth chapter of the thesis conducts instead a complementary analysis. In other words, this chapter, titled *Immigration, Local Specialization in Low-skilled Sectors and Native Education. Evidence from some EU Countries*, investigates the long-run native education responses to immigration in a set of European countries - namely, Portugal, Ireland, Greece and Spain - over the period 1980-2010. The analysis in this chapter sheds lights on a dimension of the impact of immigration that has been surprisingly scarcely explored in the previous literature, although it may have interesting consequences. Indeed, to measure the native education responses to immigration may provide evidence on the extent of substitutability or rather complementarity between immigrants and natives. The empirical analysis of this chapter is divided into two parts. The first, assesses the direct and indirect effects of immigration on native schooling, as well as the overall effect that encompasses the other two. Overall, the results indicate that in the period and countries considered, the presence of immigrants is associated with a reduction in the probability of natives to acquire at least upper-secondary education. The second part also considers the local employment structure and analyzes how the links between immigration and sectoral composition affect the decision of natives to invest in human capital. More in detail, I consider the specialization in low-skilled type of sectors (that is, those that require a low amount of human capital). All in all, the results indicate that the negative native education responses previously identified are stronger in regions specialized in low-skilled sectors.

Finally, the fifth chapter of the thesis summarizes the overall conclusions of the thesis and presents its main drawbacks, as well as the perspectives for future research.

Chapter 2

Immigration and Native Employment. Evidence from Italian Provinces in the Aftermath of the Great Recession

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2.1 Introduction

Although immigration from poor to developed countries is a longstanding phenomenon (e.g. Peri, 2016), the previously smooth pattern in the flow of immigrants has experienced a sharp rise in the most recent years. This has been particularly so in the case of Europe, that gained some 22 million international migrants between 2000 and 2017 (from 56.3 million in 2000 to almost 77.9 million in 2017). Mostly due to their geographical location, the increase in the immigrant population during this period has been exceptionally intense in some Southern European countries such as Italy, where the share of immigrants in the total population increased from around 2.4% in 2000 to 8.3% in 2017. The sharp increase in immigration, similar to that in other European countries that have only recently attracted large numbers of immigrants (e.g. Spain), explains that by 2017 the proportion of immigrants in the Italian population was close to that of the countries with greater immigration tradition, such as Germany and France[2].

This unexpected and unprecedented influx of immigrants has resulted in an increase in foreign-born workers in the Italian labor force, fueling the argument that immigration has a pernicious impact on native employment. In this regard, theoretical arguments in

[1] See https://rsaiconnect.onlinelibrary.wiley.com/doi/10.1111/pirs.12580.

[2] See the Online Supplemental Material (OSM) for more details on the recent evolution of the share of immigrant in the population in the aforementioned countries and in Italy. The OSM of this chapter is available at: https://www.dropbox.com/home/OSM_Thesis?preview=OSM_CH2.pdf.

the literature predict that this impact depends mainly on three aspects. First, whether and to which extent immigrants are complementary or substitute in production to native workers. Immigration is assumed to exert a positive effect on native's employment in the former case (Foged and Peri, 2016), whereas natives may instead experience job losses in the latter (Borjas, 2003). Second, the skill composition of the immigrant population would play a crucial role. If immigrant inflows alter the skill distribution of the workforce (because, for example, immigrants are mainly low-skilled, as is the case of Italy), the adjustment to restore the pre-immigration equilibrium will imply changes not only in wages, but also in the employment structure (Dustmann et al., 2005). Last, the employment effect of immigration would depend on the amount of rigidities in the labor and product markets, that is, on the institutions of the host country. In brief, wage adjustments to immigration shocks would be more intense in economies with flexible institutions whereas native job losses would be more frequent in countries with restrictive institutions (Angrist and Kugler, 2003).

The evidence on the employment impact of immigration is inconclusive. Some studies for the U.S. estimate a detrimental impact of immigration on native employment (Altonji and Card, 1991; Anastasopoulos et al., 2018), while others provide evidence in favor of a negligible and even positive impact (Card, 2005; Ottaviano et al., 2013; Basso and Peri, 2015). In the case of Europe, studies have focused on countries with a long tradition of immigration - Hunt (1992) and Edo (2015) for France; Dustmann et al. (2005) and Lemos and Portes (2014) for the U.K.; D'Amuri et al. (2010), Glitz (2012) and Dustmann et al. (2017) for Germany; Foged and Peri (2016) for Denmark; Basten and Siegenthaler (2018) for Switzerland -, whereas evidence for countries in which immigration is a more recent phenomenon is less abundant- e.g. González and Ortega (2011) for Spain[3]. Overall, it can be said that the evidence from immigration in Europe is also ambiguous, although results in recent studies that exploit variability among European countries point to a more intense employment response in economies with less flexible institutions (Angrist and Kugler, 2003; D'Amuri and Peri, 2014; Moreno-Galbis and Tritah, 2016). In the specific case of Italy, the literature on the employment effect of immigration is surprisingly scarce, particularly when it comes to the impact of the most recent immigration episodes[4]. Venturini (1999) analysed the impact of illegal immigrant workers in Italy on legal employment between 1980 and 1995, concluding that immigrants working without a regular contract crowd-out legal workers in the agricultural sector. On the other hand, she identifies the presence of a complementarity in production between legal workers and illegal foreign workers in the non-tradable sector. Along these lines, Venturini and Villosio (2006) concluded that the presence of regular immigrants in Italy in the period from 1993 to 1997 did not affect the probability of native workers to change their status from employed to unemployed and vice-versa. For a period closer to the one in this study, Labanca (2020) analysed the employment impact of the unexpected migration flows subsequent to the Arab Spring (from 2009 to 2012). His findings suggest that immigrants tended to displace native Italian workers in the short-run, particularly in sectors like mining, hotel and restaurant, and wholesale trade. The effect is

[3] See Table 1 of Dustmann et al. (2016) for detailed information on the countries for which evidence is available.

[4] However, it is important to mention that other recent studies have analysed the effect of immigration in Italy on different magnitudes. For example, De Arcangelis et al. (2015) indicate that an increase in the provincial foreign-born population has a positive impact on the value added of the manufacturing sector as compared to the value added of the services sector. Brunello et al. (2020) find instead that the presence of low-skilled immigrants has induced a human capital polarization in the Italian provinces. Finally, Bratti and Conti (2018) find no evidence of an impact of low-skilled immigrants on innovation, while Michelangeli et al. (2019) conclude that ethnic minorities affect positively productivity of Italian provinces.

instead positive in construction and educational services[5].

Therefore, according to theoretical predictions and existing evidence, the impact of recent immigration episodes in Italy may have been less straightforward than some actors preach and, at the same time, may have affected different population groups differently, depending on their work characteristics and the elasticity of their labor supply. This study tackles this issue by analysing the impact that the recent migratory waves have exerted on the native employment of the Italian provinces, in a context of economic recession and quite rigid labor market institutions.

The Italian case is of particular interest for several reasons. First, mostly because of its position in the centre of the Mediterranean Sea, Italy has become one of the most popular destinations for African migrants since the beginning of the 21st century, while it has also attracted citizens of Central and Eastern European countries since their accession to the European Union in 2004 and 2007 (Hanson and McIntosh, 2016; Labanca, 2020). Interestingly, part of this period of intense immigration in Italy overlaps with the severe downturn of the Italian economy caused by the global financial crisis and the European sovereign debt crisis. According to OECD (2017), Italy began to recover from the long and deep recession caused by these crises only in 2017. The figures indicate that the Italian real GDP per capita fell by about ten percentage points during the recessionary period. In this regard, the extant literature indicates that immigrants usually have (i) a lower reservation wage; and (ii) different social norms (e.g. Edo, 2015). This makes them more likely to be hired, as they are less costly compared to natives. The presence of a recession may further increase this mechanism as firms and capital owners often need to face budget cuts. Therefore, the potentially negative employment effect of immigration may be further exacerbated by the economic downturn. In this scenario of deep recession, one of the sectors more harshly affected was the labor market. To this extent, it is important to note that the Italian labor market has been characterised by a high degree of employment protection and downward wage rigidity[6]. In contrast to countries like the U.S. and U.K., wage flexibility in Italy is constrained by the system of wage bargaining, which is highly centralized at the national level (e.g. D'Amuri et al., 2010). This makes wages less sensitive to labor supply (or demand) shocks, provoking adjustments via changes in employment. This feature clearly plays a crucial role in the extent to which local labour markets are able to absorb an immigration-induced supply shock. Hence, it is motivating exploring how local employment reacted to sizeable immigration flows in a context of economic recession in a country characterized by a far from flexible labor market.

Second, the distribution of immigrants in the Italian geography is far from uniform. In fact, spatial disparities in the immigrant population resembles the ones frequently reported for other socio-economic variables in Italy (González, 2011). This is consistent with the attraction of immigrants to places that offer greater economic opportunities. Immigrants were about 10% of the population in the northern and central parts of the country in 2017,

[5]It should be noted that Labanca (2020) focused only on illegal immigrants from Egypt, Libya, Tunisia and Yemen, and was interested in the specific sectorial effects. In contrast, our study considers immigrants regardless of their country of origin and pays attention to the heterogeneity of the effect depending on the gender and skills of the native population.

[6]Italy is among the countries with the highest values of the OECD strictness of the employment protection index (http://www.oecd.org/employment/emp/oecdindicatorsofemploymentprotection.htm) and the ILO employment protection legislation index (https://www.ilo.org/dyn/eplex). Similarly, the value of the index of flexibility of wage determination produced by the World Economic Forum (https://tcdata360.worldbank.org) reveals that the degree of centralization of wage bargaining in Italy is comparable to that in France, Germany, and Spain, but far above that in Denmark, U.K. and the U.S.

whereas they were just about 4% in the South. In terms of the total number of immigrants in Italy that year, 58% of them concentrated in the northern regions, about a quarter in the central regions, while only 16% in the Southern ones. It is worth noting that there were also sizeable differences in these figures between provinces within regions. For example, in Lombardia, a region in the North, the proportion of immigrants in the population was 5.1% in the province of Sondrio and 13.9% in that of Milano, whereas in Puglia, in the South of Italy, the rate was 2.2% in Taranto and twice this figure in Foggia (4.5%)[7]. The great spatial variation in economic performance and resilience to recessions (Faggian et al., 2018), jointly with the heterogeneous distribution of immigrants that characterizes the Italian provinces, provide an excellent framework to evaluate the reaction of native employment to a migration-induced shock to local labor markets during a period of severe economic downturn.

Finally, the issue of immigration has reached a central position in the political debate in Italy today (Mayda et al., 2018)[8]. There is a fairly widespread belief that the labor market outcomes of the Italian-born have worsened as a result of increased labor competition brought by immigrants (Mayda, 2006). In fact, the perception that immigration hinders the employment of natives has gained momentum, amplifying anti-immigration messages. To this extent, if in the past the Italian public opinion was split into two opposing factions, partisans and opponents (Gavosto et al., 1999), more recently the latter have somewhat "taken over" the former[9].

Against this background, this study tests whether the recent influx of immigrants to local labor markets in Italy has really had a negative impact on the employment of natives. By using data drawn from the Italian Labor Force Survey (LFS) and the Demographic Portal of the Italian Office for Statistics (DP-ISTAT) during the period 2009-2017, we apply the so-called "spatial correlation approach" (Borjas, 2014) to estimate the effect of a change in the immigrant population on the change in native employment in the Italian provinces. In the first place, this effect is estimated for the overall native population. Then, differences between groups of natives are explored. To be clear, the study hypothesizes that the complementary-substitutability relationship between immigrant and Italian-born workers varies for high- and low-educated natives and by gender. In all cases, in order to identify the causal link that connects immigrants to native employment, we control for potential labor demand shocks and compensating adjustments through internal migration that could confound the estimate of the impact of immigration. The empirical model also controls for unobserved province effects and province-specific trends in native employment. While the former seeks to account for the large heterogeneity of labor markets between the Italian provinces in general, and their employment levels in particular, the latter aims to capture differences among provinces in the path followed by employment after the crisis. In addition, an IV estimator is implemented to account for the likely endogeneity caused by the non-random sorting of immigrants into local labor markets. To isolate the supply-driven shocks associated to immigration in each provincial labor market, the empirical analysis uses the so-called "shift-share" instrument, that combines information about the pre-sample settlements of immigrants in the Italian provinces and the evolution in the number of immigrants

[7]See Figure SM1 in the OSM for additional information on the distribution of immigrants in the Italian provinces.

[8]In the political elections of March 2018, one of the parties that won most public support was the Northern League, well-known for its anti-immigration rhetoric.

[9]As an example, the 2017 Special Eurobarometer (number 469) reveals that the Italian population greatly overestimates the proportion of immigrants in the country's population. Similarly, the evidence derived from the Standard Eurobarometer surveys indicates that immigration and unemployment are among the problems that most concern the Italian population.

by country of origin in the whole of Italy.

In contrast with the idea that immigrants indiscriminately "take away jobs from natives", the evidence in this study points to an overall negligible impact of immigration on native employment in the representative Italian province. However, when considering the effect on specific groups of natives the results reveal a positive impact on high-educated and a negligible one on low-educated individuals. When using occupations instead of formal education to distinguish native workers with different labor market skills, the results identify a positive, albeit marginally significant, effect on skilled manual native workers, whereas a negligible impact is observed for workers in occupations that require lower skills and for white collar workers. These results partly deviate from the canonical theoretical model of immigration (Boeri and Van Ours, 2008) according to which immigrants - that in Italy are mostly low-skilled (see Bratti and Conti, 2018) - act as complementary with high-skilled natives but compete for jobs with low-skilled ones. The evidence in this study suggests that in the Italian provinces, during the period under analysis, immigrants would have been less substitutes with low-skilled natives than estimated by Ottaviano and Peri (2012) for the U.S. and by Romiti (2011) for Italy in the period from 1987 to 2004. Interestingly, the results that distinguish by gender indicate that the employment of native males was not significantly affected by the immigration shocks of the 2009-2017 period. This is so regardless of their skills. In contrast, the employment of native females in the representative Italian province would have been stimulated by immigration, particularly in the case of women with high education and working in skilled manual occupations.

The rest of the chapter is structured as follows. Section 2.2 outlines the empirical model and discusses the identification strategy, while Section 2.3 introduces the dataset and provides preliminary descriptive evidence on the relationship between immigration and native employment. The main results are presented and discussed in Section 2.4, distinguishing between the overall effect of immigration and the specific effects for groups of natives formed according to their skills and gender. Finally, Section 2.5 concludes.

2.2 Empirical Model

2.2.1 The Spatial Correlation Approach

Our interest in this study is the estimation of the short-run impact of immigration on the native employment of the representative (average) Italian province in the period under analysis. To this aim, we follow the so-called "Spatial Correlation Approach", pioneered by Altonji and Card (1991), which exploits the fact that different places generally experience non-homogeneous immigrants' inflows (in terms of the number of people entering each particular labor market)[10]. The uneven spatial distribution of foreign-born individuals in Italy represents an interesting source of variation that can be exploited to estimate the impact of the recent immigration flows on native employment. As in Card (1990); Hunt (1992); Basso and Peri (2015); Foged and Peri (2016); Borjas (2017), we estimate the effect of interest from a specification where the employment of the total native population in a province is assumed to depend on the total number of immigrants in that province. Unlike the "National Skill-Cell Approach", the one based on regional variations allows to identify the "overall" impact

[10] By contrast, the "National Skill-Cell Approach" initiated by Borjas (2003) relates the labor market outcome of interest in a group of natives with similar education and work experience (skill-cell) to the number of immigrants within the same skill-cell in the country as a whole.

of immigration rather than the "relative" effect that immigrants exert on the most similar natives, i.e. belonging to the same education-experience group (Dustmann et al., 2016). In addition, in contrast to the approach based on the skill-cells, the spatial strategy does not impose the assumption that immigrants and natives are homogeneous in terms of their observable levels of education and experience. Indeed, the evidence indicates that immigrants experienced the so-called skill-downgrading (Dustmann and Preston, 2012), leading to an incorrect classification of immigrants in the education-experience groups that, in turn, biases the estimated impact of immigration.

Based on Card and Peri (2016), the baseline specification used to estimate the overall impact of immigration on native employment in the Italian provinces is[11]:

$$\frac{\Delta(N_{p,t})}{L_{p,t-1}} = \psi_t + \mu_p + \beta \frac{\Delta(M_{p,t})}{L_{p,t-1}} + \nu_{p,t} \tag{2.1}$$

where $\Delta(N_{p,t}) = (N_{p,t} - N_{p,t-1})$ and $\Delta(M_{p,t}) = (M_{p,t} - M_{p,t-1})$ are the changes in, respectively, native employment and the number of working-age foreign-born individuals in province p between years t and $t-1$. $L_{p,t-1}$, the working-age population of province p in $t-1$, accounts for the size of the province labor market in $t-1$. Therefore, the outcome variable is the yearly change in native employment of province p relative to the size of its labor market in $t-1$, whereas the regressor proxies for the relative flow of immigrants in each province and year. The specification includes time dummies, ψ_t, to account for country-wide year-specific shocks in employment, and province fixed-effects, μ_p, that control for province-specific trends[12]. Specific trends induced by provincial differences in the impact of the crisis on employment during the period analysed is a potential source of heterogeneity that must be taken into account to properly identify the impact of immigration on native's employment (Wooldridge, 2002)[13]. Finally, $\nu_{p,t}$ is an *i.i.d.* random term with zero mean and variance σ^2_ν.

The coefficient of interest, β, measures the short-run response of relative native employment in the representative Italian province associated to an increase of immigrants in the province of one percent point of its working-age population. As mentioned above, the aggregate spatial approach internalizes the possible spillover effects between different education-experience groups and therefore identifies the overall effect of immigration on native employment of the representative Italian province over the period analyzed.

2.2.2 Identification of the Effect of Immigration

In the specification in (2.1), the relative change in native employment is regressed on the relative change in province immigration to get rid of the unobserved time-invariant differences between local labor markets that may confound the estimate of the impact of immigrant inflows. The specification in relative changes, therefore, controls for the correlation between the two variables of interest that may be due to permanent or persistent local economic conditions driving both the foreign-born population and the level of local employment. The analysis also accounts for province-specific trends in native employment

[11] Borjas (2003) suggested an alternative specification to analyze the impact of immigration on native labor market outcomes, that has been used by several subsequent studies. However, we have not followed this approach to minimize the risk and consequences of spurious correlation between the variables of interest in this study (as pointed out by Card and Peri, 2016).

[12] It should be noted that the province fixed-effects, μ_p, result from the differentiation of the specification in levels that includes province-specific trends.

[13] We thank an anonymous reviewer for raising this point.

(province fixed-effects in the specification in the changes of the variables) which is another source of province heterogeneity that could confound the estimate of the effect of immigration. But besides the unobserved local economic conditions, the empirical model must account for an important feature that characterizes the performance of local labor markets and, therefore, affects both natives and immigrants, namely the evolution of the industry in which they are employed (e.g. Acemoglu and Autor, 2011; Basso and Peri, 2015). The period analysed was characterized by the turbulences caused by the global financial crisis and the European sovereign debt crisis on the Italian economy. It is sensible thinking that the particular reaction of the Italian local economies in general and their labor markets in particular depended heavily on the productivity changes of the industries in which they are specialized. If so, the lack of control of the productivity changes that different industries experienced in the period analysed will lead to confounding the estimation of the effect of immigration. Therefore, in order to identify the specific impact induced by immigration flows on the change in native employment, we include in equation (1) the "quantity version" of the so-called Bartik shock (Bartik, 1991), defined as[14]:

$$B_{p,t} = \sum_{j} \frac{E_{j,p,t_o}}{E_{p,t_o}} \cdot \left[\frac{(E_{j,t}^{IT} - E_{j,t-1}^{IT})}{E_{j,t-1}^{IT}} \right]$$

and $E_{j,p,t_o}/E_{p,t_o}$ is the employment share of industry j in province p in the initial year t_0, $E_{j,t}^{IT}$ is the employment of industry j in Italy in year t, and thus the second term in the left-hand side is the annual growth of employment in industry j in the country. The Bartik shock captures changes in province labor demand that are sector-driven, and could hinder the identification of the effect of immigration.

A well-known criticism of the spatial correlation approach is that local labor markets are not closed economies (e.g. Borjas, 1999, 2006). This means that there may be compensatory flows if some natives move to other locations as a reaction to the changes in wages and employment opportunities induced by immigrant inflows. Under this scenario, an analysis conducted at the local level could indicate a weak (or even absent) correlation between immigrants and native labor market outcomes, not because foreign-born individuals are not actually harmful for the employment perspectives of the natives, but because internal migration diffuses the effect of the immigration shock to other local labor markets. To counter the concern of compensatory flows, studies using the spatial correlation approach have claimed that immigrants do not induce significant migratory responses by natives (e.g. Card and DiNardo, 2000; Peri and Sparber, 2011). In the specific case of Italy, Venturini and Villosio (2006) and Mocetti and Porello (2010) found that immigration has a trivial impact on overall native internal mobility, albeit there could be some compensatory responses of the low-educated and highly-educated natives that would alter the skill composition of the regional labor markets[15]. Although we believe that the annual changes considered in this study do not leave much room for labor market adjustments through compensatory population flows, we add to the baseline specification in (2.1) a control of internal migration, namely the net migration rate:

[14]This is one of the most widely used methods to capture the productivity-induced changes in labor demand in a local economy (e.g. Baum-Snow and Ferreira, 2015).

[15]In contrast with evidence from other countries (e.g. Aydede, 2017, for Canada), this is in line with the existing analyses that point to limited interregional migration in the European economies as a response to immigration shocks (e.g. Zimmermann, 2009; Glitz, 2012; Lewis and Peri, 2015).

$$IM_{p,t} = \left[\frac{(I_{p,t} - O_{p,t})}{L_{p,t-1}}\right] \cdot 1000$$

where $I_{p,t}$ is the number of people immigrating into province p at year t, $O_{p,t}$ the number of people emigrating out of the same province in the same year, and $L_{p,t-1}$ is as defined above[16].

As a result, the extended specification is as follows:

$$\frac{\Delta(N_{p,t})}{L_{p,t-1}} = \psi_t + \mu_p + \beta\frac{\Delta(M_{p,t})}{L_{p,t-1}} + \gamma B_{p,t} + \delta IM_{p,t} + \nu_{p,t} \tag{2.2}$$

As in the case of the baseline specification, in equation (2.2), the coefficient of interest is β which captures the short-run impact of immigration on native employment, net of productivity-induced changes in local labor demand, internal compensatory flows, province unobserved heterogeneity, and province-specific trends.

However, it is well known that the identification of the causal effect of immigration based on the specification in equation (2.2) faces another problem, namely that immigrants' location decisions are not randomly taken. In brief, we will observe a spurious positive correlation between the change in native employment and the inflow of immigrants if the latter tended to settle in provinces with positive, or less negative, demand-driven shocks during the period analysed. A common way to solve this bias in the estimate of the causal effect of immigration is using an instrumental variable approach. Following the path set by Altonji and Card (1991) and Card (2001), we use an instrument that proxies the labor supply-driven shocks of the immigrants' inflow. The main rationale behind this instrument is that immigrants tend to settle in locations characterized by the presence of other individuals coming from the same country of origin (e.g. Bartel, 1989). The number of foreigners from a country in province p at year t is assumed to be connected with the past number of immigrants from this country in the province but unrelated to the current shocks that affect the local economy. A shift-share type of instrument for $\Delta(m_{p,t}) = \Delta(M_{p,t})/L_{p,t-1}$, widely used in the existing literature (e.g. Card, 2001; Barone and Mocetti, 2011; Basso and Peri, 2015), is computed as:

$$\Delta(\widehat{m}_{p,t}) = \frac{\Delta(\widehat{M}_{p,t})}{\widehat{L}_{p,t-1}} = \frac{\widehat{M}_{p,t} - \widehat{M}_{p,t-1}}{\widehat{M}_{p,t-1} + ITb_{p,t-1}} \quad \text{where} \quad \widehat{M}_{p,t} = \sum_o \frac{M_{o,p,t_0}}{M_{o,t_0}} \cdot M_{o,t}$$

and $ITb_{p,t-1}$ refers to the Italian-born population in working-age in province p at year $t-1$[17]. The subscript o denotes the immigrants' countries of origin and t_0 a baseline year that must be distant enough from the years in which the change in native employment is measured to guarantee the unrelatedness to current shocks. The validity of the instrument requires that, conditional to the controls and unobserved province heterogeneity considered in equation (2.2), the distribution of immigrants by country of origin in the Italian provinces in the baseline year does not correlate with province-specific demand changes in native

[16]It should be noted that $IM_{p,t}$ accounts for the annual change in the province population due to internal (inter-province) migration decisions relative to the working-age population of the province in $t-1$.

[17]All migrants from each country of origin, instead of those of working-age, are used to compute the instrument. This favors compliance with the exclusion restriction.

employment in the period analysed. As in Bratti and Conti (2018) the baseline year is set to 1995, which is well before the onset of the financial and sovereign debt crises that strongly affected the Italian economy in the period under analysis. This year also predates the accession to the European Union of the member states from central and eastern Europe that spurred a substantial inflow of immigrants from these countries to Italy, as well as the migratory shock that followed the Arab spring. These two facts work in favour of the validity of the instrument, since it is sensible arguing that province-specific labor market shocks that affected the distribution of immigrants in 1995 do not strongly correlate with changes observed in employment from 2009 on (see Goldsmith-Pinkham et al., 2020). Even in the case of strong persistence in the province-specific shocks, their effect on employment changes about a decade and a half later should be largely captured by the controls (particularly the Bartik variable) and the elements associated to unobserved heterogeneity of the province included in equation (2.2). On the other hand, the validity of the instruments also requires that the overall inflow of immigrants from each country of origin to Italy does not correlate with shocks exerting an impact on employment changes in the province. Considering the prevalence of immigrants from the above-mentioned origins, push factors associated to the internal situation of the places of birth probably had more influence on migration decisions that pull factors motivated by the economic performance of the Italian provinces.

In any case, some evidence will be provided in section 2.4 to mitigate concerns about the exogeneity of the instrument.

2.3 Data and Descriptive Analysis

In this section we provide information on the data sources used to construct the variables introduced in the previous section. It also presents the results of a descriptive analysis that sheds some preliminary light on the relationship between the flow of immigrants and the evolution of native employment in the Italian provinces over the period 2009-2017.

2.3.1 Data Source

Population censuses are the data sources commonly used in the existing literature on the economics of immigration. However, such information is not available for Italy during the period under analysis. Therefore, alternative sources of data are considered in this study. Information on the stock of foreign-born individuals with no Italian citizenship and that for the native population used to compute the immigration regressor is taken from the DP-ISTAT[18]. In both cases, data refers to the resident population at the beginning of each year. Regarding the data on native employment, we draw the information from the microdata files of the Italian LFS, carried out on a quarterly basis by ISTAT[19]. More precisely, we use the homogeneous cross-sectional quarterly data available as of the first quarter of 2009. The LFS is representative of the main magnitudes of the Italian labor market (e.g. employment

[18]See `http://demo.istat.it/index_e.html`. Suitable data on all foreign-born individuals, either non-citizens or naturalized, is not available for the Italian provinces. As pointed out by an anonymous reviewer, this may rise concerns due to the endogeneity of immigrants' naturalization. However, it should be taken into account that our empirical exercise focuses on annual changes of the variables of interest. In this case, as emphasized by Angrist and Kugler (2003), the group of non-nationals largely overlaps the group of recently-arrived foreign born.

[19]Other studies of the impact of immigration on the European labor markets have also used data from the LFS, including Angrist and Kugler (2003), Dustmann et al. (2005), D'Amuri and Peri (2014) and Edo (2015).

status, type of job, job search, wages, etc.), disaggregated by gender, age, citizenship and geographical scope (up to the provincial level). In particular, we use the LFS files for the period 2009 to 2017 to compute the yearly changes in the number of native employees in each Italian province. As the information on the resident population is relative to the first of January of every year, in order to maximize homogeneity in the dataset, we use the data for the first quarter of the LFS for each of the years under analysis.

It is worth mentioning that, since the objective of the paper is to assess the effect of immigration on native employment, we consider only the working-age population, for both natives and immigrants. In Italy, the minimum legal working-age is 15 years, so we include in the analysis individuals from 15 to 64 years of age. On the other hand, it should be said that the main results in the paper are obtained for the Italian provinces (NUTS 3 regions in Italy), which is the territorial breakdown closer to the concept of local labor markets for which the required data for the analysis can be computed. Due to some changes in the configuration of the set of provinces in the period analysed, we had to make some adjustments that led us to work with the same group of 102 provinces for the entire period[20].

With respect to the information required to compute the instrument, the share component uses the data collected by the Italian Ministry of Interior on the number of resident permits issued to foreign-born individuals by country of origin in each province in 1995[21]. In turn, the shift component is computed using annual data on the stock of immigrants in Italy by country of origin, available in the DP-ISTAT. Figures on all immigrants instead of those in working-age are used to compute the instrument as this minimises the risk of violating the exclusion restriction[22].

2.3.2 Descriptive Analysis

Before presenting the results of the estimation of the coefficients of the empirical model sketched in section 2.2, in the rest of this section we present the descriptive statistics of the variables involved in the analysis as well as preliminary evidence on the relationship between the changes in native employment and immigration in the Italian provinces over the period analysed.

The descriptive statistics of the variables are reported in Table 2A1 of the Appendix[23]. On average during the period, native employment decreased each year by 0.45 percent points of the working-age population of the representative (average) Italian province. This is consistent with the impact of the recession on the Italian labor market. However, the value of the standard deviation confirms that this figure varied greatly between provinces and years. Information in Table 2A1 also reveals interesting differences between groups of workers. On average, the change in employment was positive for the highly-educated native workers and negative for those endowed with less education. When distinguishing by groups of natives based on occupations, the figures are somewhat consistent with the ones for the groups based on the level of education. On the other hand, the distinction of the change in employment by gender suggests the existence of interesting differences for female and male natives, that will be worth taking into account in the next section.

[20] See the OSM for details.

[21] These data were gently provided by Prof. Massimiliano Bratti (University of Milan).

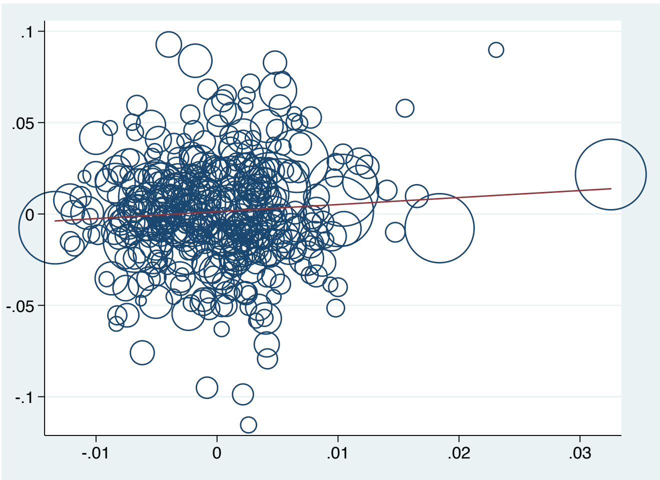
[22] We thank an anonymous reviewer for pointing this out.

[23] Details on the geographic distribution of the change in native employment and immigration are provided in Figures SM2 and SM3 of the OSM.

Regarding the immigration indicator, on average over the 2009-2017 period, the representative Italian province increased its immigrant population by 0.25 percent points of its working-age population per year. It should be stressed that this figure is consistent with that reported for the inflow of immigrants in other countries (e.g. Peri, 2016)[24]. Interestingly, the standard deviation and the maximum and minimum values confirm a great geographical variability in the immigrant inflows in the analysed period. Finally, it should be noted that the mean of the Bartik variable is consistent with an average negative shock on local labor demand over the period, while the low value of the internal migration rate agrees with the limited internal mobility in Italy, excepting in the case of some particular provinces in specific years.

The degree of association between the change in native employment and the change in immigrants relative to the working-age population in each province can be derived from Figure 2.1. The correlation between the two variables is positive and significant, which suggests that the influx of immigrants did not worsen the prospects of native employment in the average Italian province in the aftermath of the Great Recession but, on the contrary, there could have been some complementarity between immigration and native employment. However, this correlation should not be interpreted as evidence of a causal effect due to the arguments put forward in section 2.2. The estimation of such causal effect is the aim of the next section.

Figure 2.1: Scatterplot of change in native employment and change in immigrants in the Italian provinces (2009-2017).



Note: *Variables are expressed in changes over the entire period and are cleaned from the time average. The size of the circle is proportional to the initial working-age population in the province.*

2.4 Results

This section summarizes the results of the estimation of the effect of immigration on native employment from the specifications sketched in section 2.2 using the data for the

[24]It is worth noting that, as shown in Table SM1 of the OSM, immigration inflows in Italy were not of the same intensity in all the years over the period analysed. They were more intense until 2011 and in 2013 and 2014, with a net outflow in 2012 after the worsening of the economic situation in Italy due to the sovereign debt crisis, and a stagnation in the proportion of immigrants since 2014. In any case, it should be noted that this temporal pattern was not shared by all provinces.

Italian provinces described above. The impact on the overall native population is discussed first. Next, we explore differences in the effect of immigration on different groups of native workers, defined based on their skills and gender. Weighted regressions, using as weights the total number of working-age individuals in the province at the beginning of the period, are used in all cases, while standard errors are clustered at the province level.

2.4.1 Overall Impact of Immigration

The ordinary least squares estimation (OLS) of the impact of immigration on native employment from the baseline specification in equation (2.1) is reported in the first column of Table 2.1.

Table 2.1: Overall impact of immigration on native employment.

	(1) OLS	(2) OLS	(3) 2SLS
$\Delta(m_{p,t})$	0.307***	0.268**	0.305
	(0.092)	(0.129)	(0.227)
$B_{p,t}$	-	0.074	0.073
		(0.168)	(0.156)
$IM_{p,t}$	-	0.017	0.020
		(0.091)	(0.088)
Year FE	YES	YES	YES
Prov FE	NO	YES	YES
First stage F-stat	-	-	128.4
R^2	0.080	0.110	0.110
Observations	816	816	816

***Note**: The dependent variable is the change in native employment as share of the initial working-age population, while the main independent one is the change in immigrant population as share of the initial working-age population. All regressions are weighted by the total number of working-age individuals in the province at the beginning of the period. The first-stage F statistic is above the 10% maximal IV size critical value of the Stock and Yogo (2005) weak identification test. Standard errors, in parentheses, are clustered at the province level.*
*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

This estimation suggests a significant positive effect of immigration. To be clear, for the representative Italian province, an increase of immigrants by one percentage point of the local working-age population was associated to an increase in native employment of around 0.31 percentage points of the working-age population. However, the baseline specification does not control for changes in local labor demand, compensatory internal migration flows, and province-specific trends in employment that may confound the estimate of the impact of immigration. To move towards the identification of the effect of interest, we estimate the coefficients of the extended specification in equation (2.2). The OLS results are summarized in the second column of Table 2.1. It is observed that the inclusion of the control variables slightly reduces the estimate of the coefficient of interest (from about 0.31 to 0.27), affecting also its significance (the estimated effect is significant at 5% in the extended specification).

This positive response of native employment was also derived from a group of EU countries that includes Italy for the period 1998-2004 by Moreno-Galbis and Tritah (2016). In addition to the complementarity between immigrants and natives, these authors argued that immigrants could be exerting a positive externality since they are more profitable workers (the productivity-wage gap is wider for immigrants than for natives). Increase in profits would lead firms to open more vacancies, which would improve the employment prospects of the natives. However, this type of adjustment is more likely in host countries with flexible institutions, which is not the case of Italy in the period analysed. In fact, the positive estimate of the effect of immigration on native employment could be due to the bias of the OLS estimator if the location decisions of immigrants in the period analyzed were not random, and immigrants moved to those provinces less affected by the recession. The inclusion of the Bartik variable aims to control at least part of this problem, particularly with regard to industry-specific shocks that affected labor demand in each province depending on its sectoral composition. Still, to address this potential source of bias we use the immigrants shift-share instrument to implement an IV estimator for the extended specification.

Before presenting the results obtained with this estimator, it is worthwhile to show some evidence supporting the validity of the instrument. As mentioned in section 2.2, it can be argued that highly persistent province-specific shocks induce correlation between the provincial distribution of immigrants in 1995 (year used to compute the share component of the predicted amount of immigrants in each province-year) and shocks to changes in provincial employment during the period 2009 to 2017. In that case, the instrument and changes in native employment will be spuriously correlated, leading to violation of the exclusion restriction. To rule this concern out, in the spirit of Mitaritonna et al. (2017), we first computed the correlation between the instrument in the period 2009-2017 and the change in native employment between 2009 and 2010 (i.e. the first two years under analysis). The results confirm that there is no significant association between the outcome variable at the beginning of the period under analysis and the instrument measured in the following years[25]. A more accurate test would require to relate the pre-2010 values of the change in native employment to the post-2010 values of the instrument. Unfortunately, lack of data for native employment in the Italian provinces before 2009 prevents us to carry out such type of test. As an alternative, we can test the validity of the instrument computed for the second part of the period only (2013-2017) relative to changes in native employment for the first part (2009-2013). A significant correlation between them would point to strong serial correlation between earlier demand shocks and later values of the instrument (changes in predicted share of immigrants), casting doubt on its validity. The result clearly suggests that this is not the case, since the correlation between the change in employment in the first half of the period and the instrument in the second is not statistically different from zero[26]. Therefore, these results, jointly with the arguments provided in section 2.2, support the validity of the instrument based on the immigration enclaves in 1995.

The results of the first-stage regression are summarised in Table 2A2 of the Appendix. They show that the instrument correlates strongly with the change in the immigrants indicator. Accordingly, the value of the first-stage F-statistic clearly leads to rejecting the null hypothesis of weak instrument (based on the critical values computed by Stock and Yogo, 2005). The second-stage results of the IV estimator are reported in the third column of Table 2.1. It can be observed that after controlling for the endogeneity of immigration shocks, the estimated effect on native employment is still positive although not statistically different from zero. It is important to notice that although the IV estimate of the effect of immigration is slightly higher than that obtained by the OLS estimator, it is estimated with less precision, leading to a non-significant effect. Therefore, this result points to a non-significant effect of the recent inflow of immigrants on native employment in the average Italian province, being consistent with the evidence reported in D'Amuri et al. (2010) for

[25] The value of the parameter estimated in the simple regression between the two variables is -0.021, with s.e.=0.040.

[26] The value of the parameter estimated in the simple regression between the two variables is 0.011, with s.e.=0.026.

Germany and D'Amuri and Peri (2014) for a set of 15 EU countries.

This conclusion about the overall impact of immigration on native employment in the Italian provinces in the period analysed is robust to a set of alternative specifications and samples, as reported in Table 2A3[27]. In particular, column (1) shows that the estimated effect is not driven by the use of province population weights, since similar results are obtained with the unweighted IV estimator. In turn, results in column (2) correspond to the specification that adds two lags of the change of the immigrants indicator. Following Jaeger et al. (2019), in this way we aim to capture the dynamic response of employment to immigration, due to adjustments to past immigration shocks. In brief, these authors argue that the instrument will be correlated with ongoing responses to previous shocks if the provincial distribution of the inflow of immigrants remains stable during the analysed period. As a result, the IV estimator based on the shift-share instrument will not identify the short-run effect of immigration on employment but a mixture of the short- and long-run effects. Following the method proposed by Jaeger et al. (2019), we use the first two lags of the shift-share instrument to account for the endogeneity of the corresponding lags of the immigration regressor. Results in column (2) confirm that the estimated short-run effect of immigration is not statistically different from zero, this also being the case of the dynamic response of employment to the immigration shocks[28].

We also check the robustness of the estimated short-run impact of immigration to the exclusion of internal migration. This is important inasmuch as it can be argued that this variable is a bad control given that, as long as it is determined by immigration shocks, it is an outcome variable rather than a valid control. The results in column (3) confirm that the estimation of the impact of immigration is not affected by the exclusion of internal migration. As an alternative to this control, in line with Altonji and Card (1991) and Dustmann et al. (2005), we included a set of variables that capture changes in the composition of the provinces population. These are changes in native (ΔA^{NAT}) and immigrant (ΔA^{IMM}) average age, and the (log of the) ratios of high- to low- ($\Delta\Pi^{HL}$) and intermediate- to low- ($\Delta\Pi^{IL}$) educated natives. These controls aim to capture differences in the propensity to migrate of different population groups that affect the composition of the population. As shown in column (4), the estimated effect of immigration in this alternative specification is somewhat lower and remains not statistically significant. The main conclusion on this effect is also not affected by the inclusion of the proportion of workers in each skill group. Although results in column (5) confirm that the shares of skilled manual (E^{skm}) and white collar (E^{white}) workers affect positively changes in native employment, their inclusion as controls does not modify the conclusion about the impact of immigration. Furthermore, its impact is also not affected by the inclusion of an additional control that aims to account for the effect of agglomeration economies. To be clear, the results in column (6) correspond to the estimated coefficients in a specification that includes the lagged annual growth of population density (i.e. $\Delta\vartheta_{p,t-1}$) interacted with year dummies to allow for the effect of agglomeration to vary across years[29].

[27]We thank three anonymous reviewers for suggesting several of the robustness checks.

[28]The result of the first-stage F-statistic in the dynamic specification clearly rejects that the instrument and its lags are jointly weak. As indicated in Jaeger et al. (2019), evidence on weak instruments is obtained when there are no substantial changes over the period analysed in the composition by country of origin of national inflows, meaning that the instruments will be highly correlated. In this regard, it is worthwhile noting that there is no significant serial correlation neither in the instrument nor in the immigration regressor used in this study. On the other hand, it should be mentioned that similar results were obtained with a different (reasonable) number of lags.

[29]We thank an anonymous reviewer for suggesting the inclusion of the interaction between a measure of

Finally, we check the robustness of the estimated impact of immigration to the exclusion from the analysis of the largest provinces (column 7) and immigrants from EU 15 countries others than Italy (column 8). In the first case, removing provinces with more than 2 million inhabitants (Milano, Rome and Naples) leads to an increase in the estimated effect of immigration. However, it is not statistically significant as there is also a decrease in the precision with which the parameter is estimated. Therefore, there is no evidence that the estimated effect reported in column (3) is driven by the evolution of employment and immigration in the most populated Italian provinces. Similarly, results in column (8) confirm that the estimate of the impact of immigration on native employment is robust to the exclusion of the group of immigrants from countries of the EU 15.

Overall, these robustness checks confirm that there was a positive although not significant effect of immigration on native employment in the Italian provinces in the period from 2009 to 2017. In other words, empirical evidence does not support the assertion of a general negative impact of immigration on native employment in the Italian provinces in the aftermath of the Great Recession.

2.4.2 Heterogeneity in the Impact of Immigration

So far, we have considered all native-born workers in a single group, as if they were homogeneous workers and were similarly affected by immigration shocks, regardless of their job characteristics. Nevertheless, both theoretical arguments and empirical evidence seem to contradict this hypothesis (Kerr and Kerr, 2011; Borjas, 2014; Dustmann et al., 2016). In brief, immigrants can act as complementary for a part of the native population, specifically the highly educated (Chassamboulli and Palivos, 2013; Dustmann et al., 2017) and as substitute for natives with low levels of education (Altonji and Card, 1991; Dustmann et al., 2017). If so, the negligible overall estimated effect could be masking significant opposite effects for native workers with different skills, which cancel out in the aggregate.

Therefore, following the advice of Dustmann et al. (2016), we investigate the effect of immigration shocks on the employment of several native groups. Specifically, to assess the impact of immigration on natives with different skills, we classify them into two groups based on their level of education: one formed by native workers with primary and secondary education and another composed of natives with a university degree and higher stages of tertiary education. According to the low level of education of immigrants in Italy during the period analysed (e.g. Fullin and Reyneri, 2011; Bratti and Conti, 2018) and the imperfect transferability of the education that they acquired in their countries of origin, i.e. skill-downgrading (e.g. Dustmann and Preston, 2012), the former group of natives would have been more exposed to immigrants' competition. Also, the elasticity of labor supply is likely to differ between the two groups, leading to different employment responses to the immigration shocks (Dustmann et al., 2016). Finally, natives with different levels of education could be subject to downward wage rigidities with different intensity (e.g. depending on the type of contract - Edo, 2015). Therefore, the extended specification in equation (2.2) is estimated by IV for the samples defined by these two categories of workers to identify the specific effect of immigration on the employment of native workers of high and low education[30].

agglomeration and year dummies as a robustness check. The results in the next subsection are obtained without the agglomeration controls since their inclusion does not modify the estimated effect of immigration and due to our concern about the endogeneity of population density, the treatment of which is beyond the scope of the current study. Undoubtedly, this is an issue that deserves further attention in a specific study.

[30]The results of this section using the OLS estimator are reported in Tables SM4 and SM5 of the OSM.

The results are summarized in the first two columns in Table 2.2. They suggest a positive impact on the highly educated natives and a negligible one for those with a low endowment of education. To make it clear, an increase in immigrants of one percentage point of the local working-age population in an Italian province would have caused, on average, an increase in the employment of natives with high education of 0.37 percentage points of the working-age population. Surprisingly, the estimated impact on low-educated workers is very close to zero (i.e. -0.066) and not statistically significant. This result, although somehow counterintuitive[31], is consistent with a situation in which immigrants are employed in occupations different from the ones undertaken by natives (even if similarly skilled, see Ottaviano and Peri, 2012), that are typically manual intensive (Peri and Sparber, 2009; Foged and Peri, 2016). Therefore, in this scenario immigrants (i) do not directly compete with natives and (ii) induce natives to upgrade their jobs, moving to more communication intensive tasks, for which they have a comparative advantage *vis-à-vis* immigrants (Peri and Sparber, 2009; Giuntella, 2012).

Table 2.2: Impact of immigration on native employment by skills.

	by Education		by Occupation		
	(1) Highly Educated	(2) Low Educated	(3) White Collars	(4) Skilled Manual	(5) Blue Collars
$\Delta(m_{p,t})$	0.371**	-0.066	0.095	0.239	0.070
	(0.170)	(0.142)	(0.149)	(0.147)	(0.105)
$B_{p,t}$	-0.237**	0.310**	-0.145*	0.249*	-0.022
	(0.101)	(0.153)	(0.087)	(0.151)	(0.090)
$IM_{p,t}$	-0.038	0.058	0.031	-0.050	0.041
	(0.051)	(0.100)	(0.045)	(0.089)	(0.039)
Year & Prov FE	YES	YES	YES	YES	YES
Centered R^2	0.077	0.098	0.195	0.122	0.039
Observations	816	816	816	816	816

***Note**: IV estimates using as instrument the change in the shift-share variable based on the residence permits issued in 1995. Each column refers to a different sample of the native population as indicated at the top of the column. All regressions are weighted by the total number of working-age individuals in the province at the beginning of the period. In all cases the value of the first-stage F-statistic is 128.4, well above the 10% maximal IV size critical value of the Stock and Yogo (2005) weak identification test. Standard errors, in parentheses, are clustered at the province level.* *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

However, the interpretation of these results should take into account that "over-education" is a characteristic of labor markets in different countries in southern Europe, including Italy (e.g. Flisi et al., 2014). In short, the considerable proportion of native workers in the Italian provinces employed in occupations that required less education than they had can somehow affect the conclusions derived for the groups of workers with different levels of educational attainment. To overcome this drawback, under the usual assumption that occupations differ in term of the required skills, we complement the analysis with the results of the grouping of the native population according to occupations. In particular, we classify native workers into three categories of occupations, from more to less skilled: white collar, skilled manual, and blue collar. The results are shown in the last block of columns in Table 2.2. Although the estimated effect of immigration is not significant in any of the skill groups, the point estimate for skilled manual natives is much higher than that of the other two groups, being

[31] The canonical theoretical model of immigration predicts a negative effect of low-skilled immigrants on the employment prospects of their natives counterparts (Boeri and Van Ours, 2008).

almost significant (at the 10% level). Therefore, these results suggest that, in the aftermath of the Great Recession, immigration may have stimulated, on average, the employment of skilled manual natives, while having no effect on white collar and low skilled manual workers. However, there could have been a wide dispersion in impact even for skilled workers which, as shown below, can be explained by different responses of female and male natives.

As a final stage in the analysis, we explore heterogeneous responses in native employment by gender. The reason is that some studies have suggested that the impact that immigrants exert when entering the host country's labor market might affect in a different way male and female natives (Barone and Mocetti, 2011; Farré et al., 2011; Forlani et al., 2015). This could be particularly important in the case of the Italian labor market, characterized by striking gender disparities. For example, the employment rate of the working-age population in Italy in 2017 was 67.1% for males but only 48.9% for females. However, the gender gap narrowed in the case of workers with tertiary education. In this case the employment rates were 83.1% and 74.7%, respectively[32]. In this context, the complementarity/substitutability mechanisms may have worked differently for native male and female workers. For example, Barone and Mocetti (2011) argued that the high presence of immigrants providing household services is associated with an increase of the hours worked by the high-skilled native females. A gender heterogeneous response could indeed be behind the low precision with which the effect of immigration is estimated for the overall population of natives and, particularly, for the group of skilled manual workers.

For these reasons, in Table 2.3 we report not only the estimated effect of immigration on the employment of female and male natives, but also that obtained by distinguishing between female and male workers of different levels of education and occupations. First, it is observed that the estimate of the overall effect for female natives is positive and significant, whereas it is not statistically significant and, indeed, very close to zero in the case of their male counterparts. Second, the distinction by skills reveals interesting differences in the reaction of female and male employment to immigration shocks. The results confirm complementarity with the highly educated natives. More precisely, they suggest a positive and significant effect for the high-educated females of around 0.20. For males with high education the estimated effect is somewhat smaller in magnitude (0.18), being estimated with much less precision. In fact, for this group of native workers the estimated effect is not significant from a statistical point of view. Interestingly, the results do not support the claims that immigrants hinder employment of the low-educated Italians. The effect estimated for the native males with low education is negative, which is consistent with certain degree of substitutability between low-skilled immigrants and their native males counterparts. However, the coefficient for this group of workers is not statistically significant. The effect of immigration is also not significant for low-educated native females, although in this case the point estimate is positive.

The positive impact on employment for females (both high- and low-skilled) may be explained by the fact that immigrants, particularly females (whose share over the total immigrant population in Italy is higher than that of males in the period under analysis[33]) tended to substitute native females in the household production services (Cortés and Tessada, 2011; Farré et al., 2011) therefore allowing the latter to increase their labor force participation. Consistent with the evidence in Forlani et al. (2015) from a group of de-

[32]Data from the LFS available in the Eurostat Database, https://ec.europa.eu/eurostat/data/database.

[33]For more details, refer to Table SM1 of the OSM.

veloped countries, the evidence from the Italian provinces during the period analysed in this study supports a positive impact of immigration on the employment of native females, particularly for highly-skilled women.

Table 2.3: Impact of immigration on native employment by gender and skills.

		by Education		by Occupation		
	(1) All Workers	(2) Highly Educated	(3) Low Educated	(4) White Collars	(5) Skilled Manual	(6) Blue Collars
Panel A: Women						
$\Delta(m_{p,t})$	0.302**	0.197***	0.106	0.023	0.220**	0.055
	(0.123)	(0.069)	(0.091)	(0.097)	(0.093)	(0.045)
$B_{p,t}$	0.096	-0.078	0.173*	-0.029	0.151	-0.067
	(0.103)	(0.075)	(0.103)	(0.066)	(0.101)	(0.058)
$IM_{p,t}$	0.012	-0.020	0.032	0.042	-0.020	-0.008
	(0.053)	(0.033)	(0.049)	(0.036)	(0.051)	(0.018)
Year & Prov FE	YES	YES	YES	YES	YES	YES
Centered R^2	0.082	0.073	0.068	0.324	0.163	0.044
Observations	816	816	816	816	816	816
Panel B: Men						
$\Delta(m_{p,t})$	0.002	0.175	-0.172	0.073	0.018	0.017
	(0.140)	(0.116)	(0.106)	(0.070)	(0.128)	(0.089)
$B_{p,t}$	-0.022	-0.160***	0.137	-0.116*	0.098	0.045
	(0.104)	(0.057)	(0.105)	(0.061)	(0.116)	(0.072)
$IM_{p,t}$	0.008	-0.018	0.026	-0.010	-0.030	0.051
	(0.064)	(0.027)	(0.074)	(0.028)	(0.053)	(0.034)
Year & Prov FE	YES	YES	YES	YES	YES	YES
Centered R^2	0.090	0.052	0.093	0.084	0.067	0.042
Observations	816	816	816	816	816	816

Note: *IV estimates using as instrument the shift-share variable based on the residence permits issued in 1995. Each column refers to a different sample of the native population as indicated at the top of the column. All regressions are weighted by the total number of working-age individuals in the province at the beginning of the period. In all cases the value of the first-stage F-statistic is 128.4, well above the 10% maximal IV size critical value of the Stock and Yogo (2005) weak identification test. Standard errors, in parentheses, are clustered at the province level.* *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

The evidence from the occupational groups that distinguish between male and female natives (columns 4 to 6 of Table 2.3) confirms the lack of a substitution effect of immigration on native employment, regardless of the skills of the latter. In fact, the results point to the complementarity between immigration and the employment of native females. To be clear, the point estimate of the impact of immigration on native male employment is quite modest (and statistically insignificant) in all occupation groups. By contrast, the impact on females in skilled manual occupations would have been significant and sizeable.

Summing up, the evidence from Italian provinces in the period that followed the financial and sovereign debt crises confirms that native employment reacted differently to immigration shocks depending on skills and gender. On the one hand, there would have been a positive response in the employment of highly-skilled natives, particularly in the case of females. On the other, recent immigrant inflows to the Italian provinces would have not hinder significantly the employment of low-skilled natives.

2.5 Conclusions

This study has provided evidence on the short-run effect of recent immigration shocks on native employment in the Italian provinces. The results are particularly interesting because they have been obtained for a country where immigration is a relatively recent phenomenon, the inflow of immigrants in recent years has been particularly intense, and their geographical distribution has been far from uniform. Besides, in contrast to most previous studies, this one has considered a period of economic recession that strongly hit the labor markets of Italian provinces and, in particular, the employment prospects of their native populations. The overlapping of large immigration inflows and job losses would have contributed to fuelling a passionate anti-migratory rhetoric. Interestingly, the study has estimated differentiated employment responses for separate groups of natives to account for heterogenous impacts of the immigration shocks depending on characteristics of the natives and their labor supply elasticities. All of this for an economy with intense downward wage rigidity, a feature that has been shown to favour employment adjustments to immigration shocks.

Once local labor demand shocks, internal compensatory flows, province-specific trends in employment and, especially, the non-random spatial distribution of immigrants are controlled for, the results point to a negligible adjustment of provincial native employment to the immigration shocks over the period analysed. However, the study shows that this overall impact of immigration on employment hides interesting heterogeneous responses of different groups of natives. To be clear, the evidence from the Italian provinces since the onset of the Great Recession confirms that the employment response varied according to the skills of the natives. While immigration inflows would have stimulated employment of the highly-educated natives, the impact on the low-educated would have been negligible. Therefore, although most immigrants that arrived in Italy in the period under analysed were low-educated and that they probably experienced to some extent the skill-downgrading, it does not seem that immigrants substitute native of similar skills. Labour market rigidities that make difficult and costly to fire workers to replace them with newcomers could partly explain this result.

Interestingly, the study has revealed that the distinction by gender is crucial. There are no signs that clearly indicate a pernicious effect of immigration inflows on the employment of native males in the aftermath of the Great Recession. In sharp contrast, the evidence points to a positive response of female employment, that would be particularly significant in the case of native women with high education and employed in skilled manual occupations. This result is consistent with immigrants substituting native females in housekeeping and child and elderly care, which leads them to increase participation in the labour market. In this regard, the results in this study are particularly important from a policy perspective due to the still substantial gender disparities in the participation rate that characterize the Italian labour market and the differences in female participation between the Italian local economies.

Finally, we must admit some shortcomings of this empirical exercise. For example, the analysis focused only on the partial (short-run) employment effect of immigration, although responses in the longer-term involving different mechanisms (impact on productivity, investments in education of the natives, innovation, etc.) can also be of great importance. On the other hand, it could be argued that the annual changes considered in this study prevent controlling for highly persistent dynamic effects even in the case of adopting the approach suggested by Jaeger et al. (2019). Last, it should be noted that the study have just considered the employment effect of immigration at the extensive margin, while responses at the intensive margin could also be relevant. In any case, based on arguments and evidence in the extant literature, we can speculate that these additional sources of influence of the immigration shocks in Italy in the aftermath of the Great Recession would have probably contributed to enhance the employment prospects of the Italian-born beyond the short-run effect estimated in this study. This, therefore, would contradict one of the most powerful arguments of rhetoric against immigration in Italy and in other European countries that have experienced similar immigration episodes in the recent past.

2.6 Appendix

Table 2A1: Descriptive statistics.

Variable	**Mean**	**S.D.**	**Min.**	**Max.**
Change in native employment				
Whole population				
All workers	-0.0045	0.0258	-0.1082	0.1068
High-educated	0.0026	0.0157	-0.0459	0.0663
Low-educated	-0.0070	0.0250	-0.1018	0.0735
White collars	0.0009	0.0147	-0.0540	0.0424
Skilled manual	-0.0050	0.0255	-0.1321	0.0937
Blue collars	-0.0002	0.0140	-0.0422	0.0616
Only women				
All Women	-0.0010	0.0158	-0.0732	0.0705
High-educated women	0.0017	0.0010	-0.0345	0.0373
Low-educated women	-0.0027	0.0151	-0.0741	0.0882
White collars women	0.0016	0.0010	-0.0314	0.0360
Skilled manual women	-0.0023	0.0158	-0.1136	0.0670
Blue collars women	-0.0003	0.0071	-0.0283	0.0342
Only men				
All Men	-0.0035	0.0179	-0.1057	0.0654
High-educated men	0.0009	0.0101	-0.0325	0.0597
Low-educated men	-0.0044	0.0179	-0.0829	0.0542
White collars men	-0.0007	0.0098	-0.0325	0.0363
Skilled manual men	-0.0028	0.0180	-0.0707	0.0491
Blue collars men	0.0001	0.0116	-0.0361	0.0583
Change in immigrant population	0.0025	0.0066	-0.0289	0.0393
Controls				
Bartik variable	-0.0024	0.0131	-0.0368	0.0731
Internal migration	-0.0765	3.1367	-11.2624	7.9068
Natives' average age	41.4648	1.0664	38.5320	43.8006
Immigrants' average age	36.8774	1.0769	33.8902	39.9180
Log high/low educated natives	-0.6109	0.3993	-1.9667	0.6351
Log interm./low educated natives	0.3312	0.2940	-1.1513	1.3583
Employment shares				
High-educated	0.1875	0.0451	0.0725	0.3424
Low-educated	0.8125	0.0451	0.6576	0.9275
White collars	0.1675	0.0402	0.0464	0.3736
Skilled manual	0.6696	0.0489	0.4375	0.8434
Blue collars	0.1629	0.0383	0.0488	0.3226

Note: *Mean, standard deviation, minimum and maximum of the variables used in the analysis, using the observations of all provinces and years in the sample.*

Table 2A2: First-stage estimates.

	(1)	(2)	(3)
$\Delta(\widehat{m}_{p,t})$	0.441***	0.663***	0.506***
	(0.064)	(0.059)	(0.070)
Controls	YES	YES	YES
Year FE	YES	YES	YES
Prov trends	NO	YES	YES
Weights	YES	YES	NO
R^2	0.770	0.890	0.834
F-statistic	47.3	128.4	51.7
Observations	816	816	816

***Note**: The dependent variable is the change in immigrant population as share of the initial working-age population, while the main independent variable is the change in the shift-share instrument based on residence permits issued in 1995. Controls include Bartik and internal migration variables. Column (1) reports the first-stage estimates without the inclusion of province FE, column (2) corresponds instead to column (3) of Table 1 and column (3) to column (4). The first-stages corresponding to the other columns of Table 1 are available upon request. Standard errors, in parentheses, are clustered at the province level.*
*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

Table 2A3: Robustness tests on the overall impact of immigration on native employment.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
$\Delta(m_{p,t})$	0.323	0.198	0.295	0.211	0.331	0.301	0.505	0.281
	(0.395)	(0.213)	(0.212)	(0.235)	(0.232)	(0.288)	(0.510)	(0.215)
$\Delta(m_{p,t-1})$	-	0.021	-	-	-	-	-	-
		(0.321)						
$\Delta(m_{p,t-2})$	-	-0.248	-	-	-	-	-	-
		(0.162)						
$B_{p,t}$	0.015	0.073	0.070	0.104	0.092	0.134	0.033	0.058
	(0.206)	(0.156)	(0.154)	(0.153)	(0.158)	(0.162)	(0.193)	(0.158)
$IM_{p,t}$	-0.121	0.020	-	-	0.006	-0.000	-0.062	0.034
	(0.089)	(0.093)			(0.088)	(0.083)	(0.097)	(0.095)
$\Delta(A^{NAT}_{p,t})$	-	-	-	-0.017	-	-	-	-
				(0.024)				
$\Delta(A^{IMM}_{p,t})$	-	-	-	0.006*	-	-	-	-
				(0.004)				
$\Delta(\Pi^{HL}_{p,t})$	-	-	-	0.014**	-	-	-	-
				(0.007)				
$\Delta(\Pi^{IL}_{p,t})$	-	-	-	-0.005	-	-	-	-
				(0.007)				
$E^{white}_{p,t}$	-	-	-	-	0.114**	-	-	-
					(0.053)			
$E^{skm}_{p,t}$	-	-	-	-	0.069*	-	-	-
					(0.040)			
Year FE	YES	YES	YES	YES	YES	YES	YES	YES
Prov FE	YES	YES	YES	YES	YES	YES	YES	YES
Year × Pop-Den FE	NO	NO	NO	NO	NO	YES	NO	NO
Weights	NO	YES	YES	YES	YES	YES	YES	YES
First stage F-stat	51.7	46.6	137.6	145.1	128.9	57.5	71.2	97.1
R^2	0.084	0.112	0.110	0.127	0.116	0.131	0.100	0.107
Observations	816	816	816	816	816	816	792	816

***Note**: The dependent variable is the change in native employment as share of the initial working-age population, while the main independent one is the change in immigrant population as share of the initial working-age population. All regressions are weighted by the total number of working-age individuals in the province at the beginning of the period. The first-stage F statistic is above the 10% maximal IV size critical value of the Stock and Yogo (2005) weak identification test. Standard errors, in parentheses, are clustered at the province level.*
*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

The Online Supplemental Material of this chapter is available at: https://www.dropbox.com/home/OSM_Thesis?preview=OSM_CH2.pdf

Chapter 3

On the Heterogeneous Impact of Immigrants on the Distribution of Native Wages. Evidence from Recent Immigrants in Italian Provinces

3.1 Introduction

In the last few decades, many developed economies have experienced the resurgence of the migratory phenomenon (United Nations, 2017). Despite the large amount of studies carried out, there still is no consensus on whether immigrants represent a benefit or a disadvantage for the labor market outcomes of the native population in the destination countries[1].

To this extent, a crucial aspect to understand is whether immigration equally affects all native workers, or rather the effect varies depending on their position in the skill distribution. Economic theory predicts that immigration may exert downward pressure on both wages and employment opportunities of native workers that are closer substitutes to immigrants themselves, only if the immigration-induced labor supply shift alters the pre-existing local skill distribution (Dustmann et al., 2008). In order to identify the potential "winners and losers" from immigration, earlier studies distinguish between the wage impact on skilled and unskilled workers. Skills are typically defined based on educational attainments (Altonji and Card, 1991; Dustmann et al., 2005), or a combination of experience and education (Borjas, 2003; Aydemir and Borjas, 2007; Ortega and Verdugo, 2014). The wage effects can differ depending on the extent of substitutability/complementarity between immigrants and natives skills. In this respect, there is evidence of the fact that immigrants, irrespective on their skills, tend to be employed in occupations different from the ones undertaken by natives (Ottaviano and Peri, 2012; Llull, 2018*a*). This occurs because immigrants have a comparative advantage in more simple and physical type of tasks (Peri and Sparber, 2009),

[1] For a comprehensive overview of the extant literature, see Kerr and Kerr (2011) and Dustmann et al. (2016), as well as the meta-analyses of Longhi et al. (2010) and Aubry et al. (2021).

due to their often limited knowledge of the language spoken in the destination countries (Lewis, 2013) and to their different social norms (Edo, 2015). In turn, this implies that immigrants tend to specialize in sectors that make a prevalent use of low-skilled labor, like construction or agriculture (González and Ortega, 2011). In reaction to this, natives that have a comparative advantage in communication-intensive type of tasks (Peri and Sparber, 2009), tend to specialize in sectors or occupations that require such skills (D'Amuri and Peri, 2014; Foged and Peri, 2016). This mechanism ultimately implies that natives experience a sort of task-upgrade and move to occupations that are complementary with respect to the ones undertaken by immigrants and that are characterized by both higher wages (Ortega and Verdugo, 2014; Cattaneo et al., 2015) and better working conditions *vis-à-vis* the foreign-born workers (Fullin and Reyneri, 2011; Giuntella, 2012).

Another element that plays a key role in the way in which local labor markets are able to absorb an immigration-induced supply shock is the extent of wage flexibility. Indeed, in an economy characterized by wage rigidities deriving from the presence of labor market institutions, such as minimum wage, unemployment benefit or unions (as is the case for Italy), immigration may depress the employment opportunities of the native workers, without any significant effect on the wage structure (Angrist and Kugler, 2003; Boeri and Van Ours, 2008; Blau and Kahn, 2015).

With these premises, the objective of the present chapter is to provide empirical evidence on the impact that recent immigration inflows to Italy exert on the wage structure of native workers, also allowing for heterogeneous responses over the distribution of native skills. Indeed, an important limitation of the skill-based approaches performed in the previous literature is that they allocate workers to different skill-groups depending on their observable characteristics. However, this procedure may be problematic for different reasons, such as (i) the tendency of some natives to be over-educated; and (ii) the phenomenon of "skill-downgrading" suffered by immigrants (Dustmann and Preston, 2012; Dustmann et al., 2016). To deal with these issues, this study approximates the skill of each individual with the position that he occupies along the wage distribution[2], that is, it considers as high-skilled those individuals located at the upper tail of the wage distribution and as low-skilled those located at the lower tail[3]. It then estimates the impact of immigration at different points of the wage distribution, similarly to what Dustmann et al. (2013) do for the U.K. Specifically, based on a theoretical model they indicate that the parameter capturing the wage impact of immigration "is proportional to the density of immigrants along the native wage distribution". To this extent, they show that immigration exerts downward pressure on the wages in the lower end of the wage distribution (i.e. below the 20^{th} percentile) and has a positive impact in the upper end (i.e. above the 40^{th} percentile).

In addition, the empirical analysis of the present chapter also accounts for changes (over time and across provinces) in the composition of the native workforce, which is something that he literature on the impact of immigration on local labor market only marginally takes into consideration. This is because, on the one hand, regions specialized in high-skilled industries tend to attract high-qualified workers, and are therefore characterized by higher average (nominal) wages (Combes et al., 2008). However on the other, the same regions may also attract low-skilled immigrants that complement high-skilled natives in some type of activities, such as household production services (Cortés and Tessada, 2011). At the same time, the extant literature indicates that immigrants tend to cluster in more expensive

[2] A somewhat similar procedure is also followed by Card (2009*a*,*b*).
[3] The same idea is used in Matano and Naticchioni (2016).

cities or regions where nominal wages are also higher (Accetturo et al., 2014; Albert and Monras, 2020). This implies that local wages may correlate with this composition effect, and failing to consider this phenomenon can be problematic when estimating the wage impact of immigration. Indeed, the estimate of the parameter would somehow internalize the potential changes (over time and across local labor markets) in the composition of native population and therefore confound the estimates of the effect of interest.

To the best of our knowledge, this is the first paper that combines the evaluation of the impact of immigration along the native wage distribution with a methodology that allows to account for changes in the composition of the native population.

By using data drawn from the Italian Labor Force Survey (LFS henceforth) provided by the Italian National Office for Statistics (ISTAT), we conduct the analysis for the Italian provinces[4] in the aftermath of the Great Recession (i.e. over the period 2009-2017). The rationale behind the choice of the country depends of the fact that, for different reasons, Italy represents an interesting case study. First, it has only recently changed its role from being an immigrant-sending to an immigrant-receiving country (Del Boca and Venturini, 2005). Although the first migratory waves started in the 1970s, the peaks were reached in the 1990s, with inflows from the Balkans and eastern Europe, and in the 2000s with inflows from the Arab Spring countries (Labanca, 2020) and with the subsequent refugee crisis (Hanson and McIntosh, 2016). This implies that, over the period under analysis, the presence of immigrants in the Italian territory has experienced a significant increase both in terms of stocks and as share of the total population. In addition, despite its relatively poor colonial history, Italy is characterized by a consistent variation in the countries of origin composition of its immigrant population, which is an almost unique case among the western economies (De Arcangelis et al., 2015). This aspect plays a crucial role in terms of the potential economic effects of immigration (Alesina and La Ferrara, 2005), as there seems to be a positive relation between the ethnic composition of the immigrant population and native wages (Ottaviano and Peri, 2006), or productivity (Kemeny and Cooke, 2018).

Second, Italy is also characterized by a peculiar institutional framework, sufficiently different from the one of the Northern European countries or the U.S.[5]. In particular, due to the relatively strong presence of labor market institutions that hinder downward wage adjustments, the Italian labor markets are sufficiently rigid[6].

Third, in terms of the immigrant population, Italy mostly host low-skilled individuals (Docquier et al., 2014; Bratti and Conti, 2018), which is something that is not valid for countries with a longer tradition of immigration, as the U.K. (Manacorda et al., 2012), or the U.S. (Peri, 2016). Fourth, another important peculiarity of the Italian economy is its well-known socio-economic divide between a more advanced and productive North and a still struggling South. Finally, in Italy the issue of immigration has recently reached the center of the public debate and the top of the political agenda (Mayda et al., 2018).

[4]NUTS 3 regions in Italy.

[5]Which are the countries mostly considered when estimating the economic impact of immigration (see Table 1 in Dustmann et al., 2016).

[6]Italy is indeed one of the European countries that shows the highest values of both the OECD strictness of the employment protection index (http://www.oecd.org/employment/emp/oecdindicatorsofemploymentprotection.htm) and the ILO employment protection legislation index (https://www.ilo.org/dyn/eplex). In addition, according to the index of flexibility of wage determination produced by the World Economic Forum (https://tcdata360.worldbank.org), the degree of centralization of the wage bargaining mechanism in Italy is in line with that of countries with relatively rigid wage structure, such as France, Germany, or Spain, and considerably above that of countries with more flexible labor markets, like Denmark, the U.K. or the U.S.

Against this background, the hypothesis to test is that the higher density of recent immigrants in Italy in the lowest end of the skill distribution, together with the constraints to downward wage adjustments due to the Italian institutional setting, should imply a non-negative effect on both average wages and on the whole distribution of native wages. Conversely, we might expect a positive effect in the upper part of the wage distribution, due to a complementarity effect between low-skilled immigrants and high-skilled natives.

In order to test the hypotheses, we implement a two-steps procedure. In a first step, we estimate a native provincial mincerian equation that allows us to obtain a *conditional* native wage distribution, that is a native wage distribution net of the effect of observable individual characteristics. In addition, we include a set of workplace characteristics that supposedly capture workers unobserved heterogeneity. This procedure is assumed to thoroughly control for changes in the composition of the native workforce. The second step builds instead on the analysis of Dustmann et al. (2013). Specifically, we apply a spatial correlation approach that allows us to estimate the impact of immigration along the conditional native wage distribution previously obtained. We additionally control for endogeneity by including a set of local controls, and by applying an IV methodology using an instrument based on the settlement patterns of older immigrants.

We initially conduct the analysis considering all native workers together. Then, we perform an heterogeneity analysis along two lines. In the first, we distinguish by gender, to verify whether the overall results are confirmed also when considering men and women separately. The rationale behind this choice is that the wage elasticity to immigrant shocks is assumed to be quite gender-specific. If, on the one hand, men typically supply their labor, irrespective of the wage dynamics, the same is not valid for women. Immigrant workers, especially if women, may indeed substitute native women in the household production services, with a positive impact on their participation rates and wages. This is particularly valid in the Italian context, whose peculiar cultural norms imply that females are often disadvantaged in terms of economic opportunities.

In the second, we distinguish between Northern and Southern provinces, to verify whether the wage response to immigration is similar in these two portions of the Italian territory, which is something that is not possible to argue ex-ante, given their consistent disparities in economic performance. Gender differences within Northern and Southern provinces are also considered.

Two main elements emerge from the empirical analysis. First, in line with a strand of the extant literature (e.g. Hunt, 1992; Pischke and Velling, 1997; Gavosto et al., 1999; Card, 2005; Staffolani and Valentini, 2010; González and Ortega, 2011), the results contradict the idea of an "inverse relation between relative wages and immigration-induced supply shifts" (Aydemir and Borjas, 2011). By contrast, they present a scenario in which immigrants do no exert any significant impact on the average wages of their native counterparts. Second, we estimate a positive wage response in the upper part of the native wage distribution, especially in the case of female workers residing in the Northern part of the country. However, it is important to highlight that, in the more demanding specification (that is, the one that includes a comprehensive list of local labor market controls and region-specific trends), this positive effect is only marginally significant

The rest of the paper is structured as follows. Section 3.2 presents the background of the analysis, in which the most important characteristics of the Italian context are discussed, both in terms of the phenomenon of immigration and of its institutional framework. Section 3.3 depicts the empirical strategy implemented. Section 3.4 describes the data used to construct the figures, and gives some preliminary descriptive statistics on them. A more detailed explanation of the main results is then given in section 3.5, and finally section 3.6 concludes.

3.2 The Recent Dynamics of Immigration and Local Labor Markets in Italy

Regarding the phenomenon of immigration, irrespective of the relatively short-time frame, in the period under analysis (2009-2017) the presence of immigrants in the Italian territory grew considerably both in absolute terms (i.e stock) and as share of the total population. Figure 3A1 of the Appendix provides a graphical visualization of this evolution. Despite a small reduction around 2012-2013, the Figure clearly indicates the presence of an increasing pattern: from an immigrant share over the total population of almost 8 percentage points in 2009 (which corresponds around 3 million individuals), to roughly 10.2 percent in 2017 (i.e. almost 4 million individuals). These numbers imply an average annual growth rate of the foreign-born population of around 0.3 percent, which is in line to that reported for countries with a longstanding tradition of immigration (Peri, 2016). In addition, Figure 3A2, reports the evolution over time of the gender composition of the immigrant population. Interestingly, the larger portion is represented by females, which account for - on average - the 53 percent of the total immigrant population in working-age[7].

Another important aspect to underline is the high diversification in terms of the countries of origin composition of the immigrant population. According to ISTAT, in the period under analysis Italy hosted people coming from - on average - 195 different countries. This represents a fairly peculiar case among the developed economies, that is even more remarkable if we consider the relatively reduced Italian colonial history (especially if compared with other European countries such as France, Portugal, Spain, U.K., or the Netherlands). In the period under analysis, the more represented countries are Rumania, Albania, Morocco, China and Ukraine. Together, they represent - on average - around the 10 percent of the total immigrant population residing in Italy.

In terms of the skill composition of the immigrant population, in the existing literature there is a relatively broad consensus on the fact that immigrants do not represent a random sample of the countries of origin's population (Llull, 2018*b*). What is less clear is where, in the country of origin's skill distribution, are located those individuals who decide to migrate abroad. There are some studies supporting the "negative selection" hypothesis (i.e. the low-skilled individuals are the ones most likely to move), especially for destination countries characterized by a high wage inequality (Borjas, 1987; Fernández-Huertas Moraga, 2011; Abramitzky et al., 2012). Some other identify instead the presence of an intermediate (Chiquiar and Hanson, 2005), or even positive selection mechanism (Grogger and Hanson, 2011). In Italy things are slightly different. Indeed, the fact that it has only recently become a major immigrant destination country as indicated before, does not depend on the quality of life it offers, but rather on its geographical position in the middle of the Mediterranean Sea. To this extent, the unprecedented migratory waves that Italy experienced in the last decades are mostly driven by push factors originating in the sending countries[8] (Hanson and

[7] A similar process of "feminization" of the immigrant population occurred in France (see Edo and Toubal, 2017; Borjas and Edo, 2021).

[8] Once in Italy, however, their location decisions are not completely random, but might depend on the

McIntosh, 2016; Labanca, 2020). Immigrants in Italy are indeed mostly unskilled (Bratti and Conti, 2018), and struggle to be economically and socially assimilated (Venturini and Villosio, 2008; Accetturo and Infante, 2010; Fullin and Reyneri, 2011; Dell'Aringa et al., 2015).

Regardless of the growing political relevance of the issue of immigration, the existing studies on the economic impact of foreign-born workers in the Italian context are surprisingly scarce. Overall, the evidences obtained present a scenario in which immigrants have a positive impact on the native labor outcomes. More precisely, Gavosto et al. (1999), analyzing the period from 1986 to 1995, find that immigration has a positive impact on the wages of manual workers and this positive relation is particularly strong in small firms located in the North of the country. They also indicate that the results are particularly driven by two aspects. First, the strong presence of trade unions that render wages somehow insensitive to changes in the labor supply. Second, the presence of complementarity in production between immigrants and natives. This fact is additionally fostered by Venturini and Villosio (2006) and Labanca (2020) that highlight how immigration does not have an impact on native employment. Along the same lines, Staffolani and Valentini (2010) indicate that, in the period from 1995 to 2004, immigration has a positive impact on the wages of both blue and white collars native workers, employed in the "regular sector" (i.e. with a legal contract). Similarly, De Arcangelis et al. (2015) analyze how the presence of immigrants affects the productive structure of the Italian provinces. They find that, in the period from 1995 to 2006, a twofold increase of the immigrants-to-population ratio is associated with an increase in the manufacturing value added relative to the service sector value added of around 12 to 21 percentage points.

All in all, both the second chapter of the thesis, and the existing studies on the Italian case agree on the fact that the extent to which local labor markets are able to absorb an immigration-induced labor supply shock also depends on the institutional framework of the destination country. More precisely, in case in which the labor market institutions (e.g. collective bargaining, minimum wage, unemployment benefits) reduce the wage flexibility, then an immigration-induced labor supply shock may negatively affect the native employment, but not the wages of similarly skilled natives (see Blau and Kahn, 2015). At this purpose, it is important to underline that, since the early 1990s, Italy is characterized by a particular wage-bargaining system. At a first level, there is a form of collective bargaining which is set at the national level. Its general objective is to help workers to maintain their purchasing power. To this extent, unions intervene by negotiating an increase in wages so to absorb the (expected) inflation rate. This collective bargaining is repeated every two to four years. Then, there is a second level of bargaining, which is decentralized at the sub-national level (or, in some cases, at the firm level). Its objective is to contract a form of incentives mechanism related to the general performance of the firm in which each worker is employed. In addition, in this second level unions are allowed to negotiate other elements of the working conditions, such as the number of hours worked (Matano and Naticchioni, 2017). Finally, it is also important to stress that in Italy workers can individually bargain their wage with their own employers. Of course, the only threshold is represented by the minimum wage which is set at the national level (Matano and Naticchioni, 2017)[9].

presence of "older" immigrants (in terms of their year of arrival) from the same country of origin. This is the so-called *network effect* which is relatively well documented in the literature (e.g. Bartel, 1989; Munshi, 2003).

[9]For a more detailed description of the Italian institutional framework, refer to Cappellari et al. (2012) and Matano et al. (2019).

In addition, Italy is characterized by consistent gender differences in both economic performances and opportunities. According to ISTAT, in the period under analysis the employment rate of the men workers was, on average, 72.3%, considerably higher than the value of 50.4% reported for females. Along the same lines, the female unemployment rate was around 11%, while that for men is slightly lower than 9%. Also, the non-participation rate for men is almost ten percentage points lower than that for women (15.5% vs 24.1%).

Another important peculiarity of the Italian economy is its North/South divide (also defined as the "Southern Question"). This issue is as old as Italy itself and is mainly due to an insufficient process of industrialization occurred in the Southern regions (González, 2011). Because of this, Italy shall be considered as a "dualistic economy", dominated by a relatively advanced North and a less-developed South. Specifically, the Northern and Southern regions of Italy are characterized by differences in (i) industrial mix (Lagravinese, 2015); (ii) labor market characteristics (Brunello et al., 2001); (iii) education (Abramo et al., 2016); (iv) female labor force participation (Marino and Nunziata, 2017); and (v) cultural norms (Jurado Guerrero and Naldini, 1996). In addition, the process of labor market flexibilization occurred in the early 2000s has increased exposition of Italian workers to economic downturns (Lagravinese, 2015). This has been particularly pronounced in the case of Southern regions (Destefanis and Fonseca, 2007), also because they are characterized by lower level of resilience, particularly with respect to the Northern ones (Faggian et al., 2018). Moreover, despite a reduction occurred in the 1980s (Cracolici et al., 2007), Southern regions are still experiencing a phenomenon of "brain drain", particularly due to the low demand of medium- and high-skilled jobs, with all the negative consequences that this entails (Lagravinese, 2015; Anelli et al., 2019).

To this extent, the lower economic performances of female workers are particularly marked in the southern regions. More in detail, between 2009 and 2017 the female employment rate was, on average, 33.7% in the Italian "Mezzogiorno", while almost the double in the North (61%). Less strong, but still evident, is instead the comparison for men, whose employment rate was 62% in the South and 78% in the North. A similar pattern is observed for the unemployment rate, which for females was 18.8% in the South, against 7.5% in the North, while for males was 14.7% in the South and 5.5% in the North, and for the non-participation rate (43.3% in the South vs 13.5% in the North for females, and 26.9% in the South vs 8.3% in the North for males).

3.3 Empirical Strategy

In the present paper, the attention is focused on the distributional impact of immigration, that is on the impact (if any) that immigrants exerts along the native wage distribution. The empirical background of our study is represented by Dustmann et al. (2013) that were the first to conduct an analysis on the effect of immigration along the native wage distribution. This approach allows to avoid to make assumptions on the skill distribution of both natives and immigrants based on their observable characteristics (such as education). In this way, it is possible to control for two different issues: (i) the tendency of some native workers to be employed in occupations for which they are over-educated, and (ii) the fact that immigrants tend to suffer what is called "skill-downgrading" once they enter in the labor force of the destination country (Dustmann and Preston, 2012; Dustmann et al., 2013, 2016). Indeed, it is possible to affirm that the position occupied by each worker in the wage distribution is a more reliable indicator of his productivity than his level of education (Matano and Naticchioni, 2017). In addition, this approach allows us to provide some insights on the potential heterogeneity in the native response to immigration of workers situated in different points of the wage distribution, that may be concealed when considering average wages.

However, an important dimension that Dustmann et al. (2013) fail to take into account (or do so only partially) is the potential spatial sorting of workers. Indeed, there are evidences of the fact that workers with similar skills tend to cluster in the same locations (e.g. Combes et al., 2008). In the particular case of this study, the factors that native workers take into account when deciding their location may correlate with the determinants that affect the spatial distribution of immigrants within the country. If so, the estimate of the impact of immigration on native wages could include this composition mechanism of natives, therefore confounding the estimates of the effect of interest (Combes et al., 2015, 2020).

In line with these arguments, we propose a strategy that captures this phenomenon by controlling for a comprehensive set of observable individual characteristics of natives that proxy for their skills and that may affect wages. Moreover, as indicated above, in addition to observed workers characteristics, the study accounts for job characteristics, which is an indirect way of controlling for unobserved skills (such as ability, motivation, etc.). To further address the endogeneity issue deriving from the non-random immigrants' settlement across labor markets, we combine the procedure that allows to control for changes in the workforce composition with an adaptation of the methodology designed by Chetverikov et al. (2016). They indicate that the more efficient way to estimate the effect of an aggregate magnitude (such as immigration) on the distribution of a micro-level one (e.g. wages) is to implement a two-steps procedure. To this extent, in our specific case, we initially estimate an individual mincerian quantile regression for the native workers of each Italian province. From it, we compute for each year the provincial wage net of the effects of observable individual and job characteristics in a set of quantiles. In this way we obtain the native wage distribution net of the effect of spatial sorting/composition effect in every Italian administrative province and year. Then, in a second step we test what is the impact that immigrants exert along the native conditional (log) wage distribution previously computed by means of an aggregate version of the "Pure Spatial Approach" as defined by Dustmann et al. (2016).

3.3.1 Two-steps Procedure

As previously indicated, the first step of our empirical strategy consists in estimating for each province and year the distribution of native wage conditional on observed individual and job characteristics of the natives. The idea is to filter out from the native wages the effect of their characteristics that may affect their location decisions and determine differences in their productivity. This allows to identify the changes in local wages dynamics that are only attributable to the labor supply-shift induced by immigrants, and not on changes in the composition of the native population. Specifically, for each province p, the wage equation for the u^{th} quantile is given by:

$$Q^u_{w_{p(i),t}|z_{p(i),t},\,\alpha^u_{p,t}} = \gamma^u_p z'_{p(i),t} + \alpha^u_{p,t} \quad \text{with} \quad u \in U \tag{3.1}$$

where $Q^u_{w_{p(i),t}|z_{p(i),t},\,\alpha^u_{p,t}}$ is the u^{th} quantile of $w_{p(i),t}$ that is the (log) hourly wage for individual i (native), in province p at time t (expressed in euros as of 2017), conditional on $z'_{p(i),t}$ and $\alpha^u_{p,t}$. The former is a vector of observed characteristics of individual i, in province p at time t, that includes experience, education, occupation and characteristics of the contract, with γ_p^u representing a vector of coefficients associated to each of them[10]. The latter indicates instead a set of area-year fixed-effects that, along the lines of Combes et al. (2008), can be interpreted as the u^{th} quantile of (log) wage in province p at time t net of the effect of observable micro-level covariates[11].

It is important to notice that γ_p^u is specific to province p but constant over the period analyzed. This means that we assume that the impact of each micro-level covariate is only allowed to vary across provinces. The rationale behind this idea is that, on the one hand, consistent changes in the individual wage determinants are unlikely to occur within provinces, especially considering the short time span analyzed (2009 to 2017). On the other however, this procedure allows us to capture differences across provinces that are more likely present in a context such as the Italian one characterized by longstanding differences between different regions, especially Northern and Southern ones.

In the second step, we use the conditional wage distribution previously computed to empirically assess what is the impact that immigrants exert along the native wage distribution. Our specification is the standard reduced-form equation founded on the theory and widely used in the existing literature (e.g. Borjas, 1999; Card, 2001; Dustmann et al., 2005; Lemos and Portes, 2014; Basso and Peri, 2015), that exploits the so-called "spatial correlations" between immigrants and native wages (Borjas, 2014)[12]. More formally:

$$\Delta\alpha_{p,t}^u = \beta_0 + \beta_1^u \Delta m_{p,t} + \beta_2^u \Delta X_{p,t} + \psi_t + \nu_{pt}^u \tag{3.2}$$

where $\Delta\alpha_{p,t}^u = (\alpha_{p,t}^u - \alpha_{p,t-1}^u)$ is the change between years t and $t-1$ in the u^{th} quantile of the native conditional (log) hourly wage in province p at time t, obtained in the first step. The main independent variable is $\Delta m_{p,t}$ and is defined as:

$$\Delta m_{p,t} = \frac{\Delta M_{p,t}}{L_{p,t-1}} = \frac{(M_{p,t} - M_{p,t-1})}{L_{p,t-1}}$$

In practice, we define the immigration-induced supply shift as the yearly change in the local immigrant stocks, standardized by the one-year lagged total population in working-age. This allows us to somehow reduce the endogeneity concern deriving by the spurious correlation between local wages and the working-age population both measured in the same year t (Card and Peri, 2016; Dustmann et al., 2017)[13]. Finally, $X_{p,t}$ is a vector of province-

[10]Some studies also control for unobservable individual heterogeneity through the inclusion of individual fixed-effects. Unfortunately, the repeated cross-sectional nature of the dataset used prevent us to do so. On the other hand however, the richness of the information available, that provide information on both the individual level and on the characteristics of workplace, allow us to include variables that, at least partially, address this issue (a similar assumption is made in Combes et al., 2015, 2020). More details are given in section 3.4 and in Table 3A2 of the Appendix.

[11]It is important to stress that this also allows to controls for changes over time and across provinces in the composition of the native population.

[12]As in many aspects of the economics of immigration, there is an on-going debate on the validity of such approach. To this extent, Borjas (2003, 2006) argues that the parameter estimated through the "area" approach might be offset by the native internal migration. On the other hand, Peri and Sparber (2011) and Lewis and Peri (2015) provide evidences in favor of the validity of the spatial approach. As for the present study, especially because of the relatively limited internal migration of natives in Italy (and in general in Europe, as indicated by Basso et al., 2019), we believe that the spatial correlation approach is the more appropriate to exploit the relationship of interest.

[13]On the contrary, Borjas and Monras (2017) indicate that the more appropriate way to capture the immigration-induced supply shock is to use the *post-shock* local population. We address this in the robustness

specific control variables. These are aimed at capturing some factors of the local economies that may influence the wage dynamics or to affect the way in which local labor markets absorb the supply-shift induced by immigrants. Specifically, we include the "price version" of the Bartik variable (from Bartik, 1991) that is defined as:

$$B_{p,t} = \sum_{j} \frac{E_{j,p,t_0}}{E_{p,t_0}} \cdot \Delta \ln w_{j,t}$$

where the first component is the share of employment by industries j, in each province p and in the base-year t_0, which is augmented by the industry-specific wage growth at the national level in the subsequent years. This variable is assumed to control for industry-specific labor demand shocks that, if not controlled for, could hinder the identification of the effect of interest[14].

In addition, a strand of the existing literature indicates that the local population often reacts to immigration shocks by moving to regions less affected by the inflows of foreign-born workers (Borjas, 2006; Monras, 2020)[15]. Therefore, to account for the potential migratory response of the local population, we include as additional control the net migration rate, defined as:

$$IM_{p,t} = \left[\frac{(I_{p,t} - O_{p,t})}{L_{p,t-1}}\right] \cdot 1000$$

where the variable is given by the difference between the inflows $(I_{p,t})$ and the outflows $(O_{p,t})$ of people in each province-year cell, standardized by the yearly-lagged local population in working-age $(L_{p,t-1})$.

In equation (3.2) the coefficient of interest is β_1^u which is assumed to capture the percent change in the u^{th} quantile of the native wage associated to an increase of immigrants corresponding to one percent of the total population in working-age (Basso and Peri, 2015). It is important to stress that this coefficient is allowed to vary at the different quantiles of the native conditional (log) wage distribution. The sign and size of the effect will depend as well on the substitutability or complementarity of skills, and this may be different for male and females and between the North and South of Italy (depending for instance on sectoral specialization). This heterogeneity in the effect of interest depending on both the gender and geographical distinction is exploited in section 3.5.3.

3.3.2 Identification Issues

In the first step, we control for changes in the skill composition in the native workforce across and within provinces that can confound the estimates of the impact of immigrants on native wages. Still, as extensively indicated in the existing literature (see, for instance, Friedberg and Hunt, 1995) and summarized in Lemos and Portes (2014), the estimates of

checks.

[14] In the literature on the wage effects of immigration, a similar control variable is used, for instance, in Basso and Peri (2015).

[15] At this purpose, however, we want to highlight that there is a controversy in the existing literature, as there are some studies that indicate that the native migratory responses to immigrant inflows are irrelevant (e.g. Card and DiNardo, 2000; Card, 2001; Peri and Sparber, 2011).

β_1^u "would be biased in the presence of a non-zero correlation between the error term and the migration rate". There are different reasons for this to occur.

The first, and most important, is the so-called "Omitted Variable Bias" (OVB), which is present in the case in which variables that affect both local wages and the migration rate are omitted (Angrist and Pischke, 2009). On the one hand, we address this issue by including in our specification a set of local controls, namely the Bartik variable and the net migration rate, that capture both labor demand shocks and the potential internal movements of individuals between provinces that could ultimately offset the estimates on the wage impact of immigration (Dustmann et al., 2005, 2013). In addition, we also control for province-specific trends that are assumed to capture differences across provinces in wage dynamics that, if not addressed, may somehow confound the estimates of the effect of interest. Moreover, their inclusion is assumed to foster the identification of the causal link between immigration and native wages. On the other, we include a set of time and province fixed-effects that capture year- and province-specific unobserved heterogeneity. The former are captured by the variable ψ_t in equation (3.2), while the latter by means of the specification in first-differences.

Second, another source of bias may arise in presence of measurement error in the construction of the variables under analysis, especially for the variable capturing the presence of immigrants in the different areas (Aydemir and Borjas, 2011). In our case however, we do believe that the data source that we use to drawn information on the immigrant population is particularly reliable and represent an accurate measure of the immigrants stock across provinces (as indicated by Bettin and Sacchi, 2020).

Third, the last source of bias stems from the fact that (i) as already indicated, individuals migration decisions are not randomly taken but are the consequence of a rational process; (ii) once the decision to migrate abroad has been taken, also the choice of the country of destination is not random, but depends on the economic opportunities that all the potential countries of destination offer; and (iii) once in the country of destination, the settlement decision is the result of a rational process, too. These three aspects imply that the native wage and the migration variable are likely to be jointly determined, leading to a reverse causality bias in the empirical specification used. This is particularly accentuated in the case of an area approach, such as the one in equation (3.2) (Borjas, 2003). Conditioning to native skills may mitigate the reverse causality bias that, however, remains an important problem to take into account.

To address these issues (i.e. measurement error and reverse causality), we implement an instrumental variable procedure using as instrument the canonical shift-share variable, popularized in the migration literature by Altonji and Card (1991) and Card (2001), and widely used thereafter. This type of variable combines the spatial variation in the historical settlement of the immigrant population by countries of origins with the time-series variation in the immigration trends from each sending country, and is defined as follows:

$$\Delta\widehat{m}_{p,t} = \frac{(\widehat{M}_{p,t} - \widehat{M}_{p,t-1})}{\widehat{M}_{p,t-1} + ITb_{p,t-1}}$$

where $\widehat{M}_{p,t}$ is given by the sum across all countries of origin o of the share of immigrants from origin o, residing in province p in a base year t_0 (i.e. $M_{o,p,t_0}/M_{o,t_0}$), augmented by the national-level stock of immigrants from the same countries of origin in the years under analysis $(M_{o,t})$[16]. $ITb_{p,t-1}$ indicates instead the one-year lagged local Italian-born

[16]In other words, $\widehat{M}_{p,t} = \sum_o \left(\frac{M_{o,p,t_0}}{M_{o,t_0}}\right) \cdot M_{o,t}$

population.

The rationale behind this instrument is that past immigrants' settlement is (i) a good predictor of the immigrants' location decisions, since foreign-born individuals tend to settle in areas characterized by the consistent presence of people coming from the same countries of origin (Bartel, 1989; Munshi, 2003); and (ii) unlikely to be correlated with the current conditions (i.e. at time t) of the local labor markets (i.e. area p that in our case are the Italian administrative provinces), conditional to the set of controls included in equation (3.2). Besides, the share component of the instrument is built using information on the number of residence permits issued at the province level, which is presumably exogenous with respect to the local economic conditions.

Although one can never be sure to have completely eliminated these problems, we do believe that the control for differences in the skill composition of the native labor force in the first step, the inclusion of the control variables, the province-specific trends, as well as the time and province fixed-effects, combined with the instrumental variable procedure, have largely addressed the issues and allowed us to isolate the variation in native wages that is mainly due to the presence of immigrants.

3.3.3 Instrument Validity

Despite being extremely popular (also because of the usually strong first-stage) the shift-share type of instruments have recently being object of severe criticisms (Borjas et al., 2012; Adão et al., 2019; Jaeger et al., 2019; Borusyak et al., 2020; Goldsmith-Pinkham et al., 2020). The main limitation is that, in presence of persistent labor demand and supply shocks that affect both the outcome variable and the immigrants settlement patterns, the exclusion restriction of the instrument is likely to be violated. In order for this not to occur, two conditions must be satisfied. The first is that the share component, computed in a base year t_0, must not be correlated with local labor markets conditions of the period under analysis. This is fundamental because, when using this type of instruments, most of the identification comes from the share component (Goldsmith-Pinkham et al., 2020). Related to this condition, in the present paper we use 1995 as base-year (as done in Bratti and Conti, 2018). This is arguably sufficiently distant in time from three events that have likely shaped the geographical distribution of immigrants across Italian provinces. The first is the E.U. enlargement occurred in 2004 (the largest enlargement so far, involving Czech Republic, Estonia, Cyprus, Latvia, Lithuania, Hungary, Malta, Poland, Slovakia and Slovenia), 2007 (involving Romania and Bulgaria), and 2013 (when Croatia entered the E.U.)[17]. The second is the Arab Spring occurred in the period 2009 to 2012, which led to considerable immigrant inflows towards the Italian peninsula (Labanca, 2020). The third is the economic downturn caused by the outbreak of the financial and sovereign debt crises that affect severely the Italian economy and some regions more than others. The choice of a base-year that predates the beginning of the economic crisis guarantees that the economic performances of the local labor markets that shaped the spatial distribution of immigrants in 1995 are not related with the province-specific shocks occurred from 2009 onwards. In addition, as already indicated, we use information on the residence permits issued, which typically depends on the political environment, more than on the local economic dynamics. In any case, to empirically verify this assumption, in Panel A of Table 3A4 we compute the correlation between the share component of the instrument and the change in employment

[17] See https://ec.europa.eu/neighbourhood-enlargement/policy/from-6-to-27-members_en.

and wages computed on the first and second parts of the period under analysis. The absence of significance in the parameter estimates corroborates the validity of the instrument used.

The second condition is instead the absence of serial correlation of the immigrant stock by countries of origin. To this extent, Jaeger et al. (2019) indicate that this concern is less relevant in the the case of the European economies than for the U.S., which instead exhibit a country-of-origin mix of their immigrant population fairly stable over time. In any case, to eliminate any doubt, we compute the autocorrelation coefficient for the stock of immigrants by countries of origin in the period 2009-2017. Its estimate is around 0.67, which is considerably lower than that estimated by Jaeger et al. (2019) for the U.S. (which varies between 0.9-0.99), and in line with that for Spain estimated by Özgüzel (2021).

Finally, in line with Mitaritonna et al. (2017), in Panel B of Table 3A4 we perform some further tests to assess the validity of the instrument. Specifically, we first compute the correlation between the change in the instrument over the period 2009-2017 and the change in native employment and average wages between the first two years under analysis, namely 2009 and 2010. The parameter estimates, reported in columns (1) and (3), are not statistically significant. We additionally estimate the correlation between the change in the instrument computed on the second part of the period under analysis (2013-2017) and the change in employment and wages computed on the first part of the period under analysis (2009-2013). We report the estimates in columns (2) and (4). Again, all coefficients are not statistically different from zero. All in all, these results are reassuring in terms of the validity of the instrument used.

3.4 Data and Descriptive Analysis

To construct the figures used in the empirical analysis we use two main data sources. In particular, information on wages and on individual characteristics of the native population used in the first-step are drawn from the microdata files of the Italian LFS for the period 2009-2017[18] that is conducted by ISTAT on a quarterly basis[19]. It is important to underline that these are repeated cross-sections, since the lack of an individual identifier prevent us from tracking individuals over the different waves of the LFS. The choice of the period under analysis is attributable to two factors. On the one hand, it allows us to capture at the same time (i) the aftermath of the crisis[20]; and (ii) the migratory crisis of 2011, induced by the outbreak of the Arab Spring, and the following refugee crisis (Hanson and McIntosh, 2016). On the other, information on earnings in the LFS is only available from 2009 onwards, precluding us the possibility of going back in time. Information on the working-age population (i.e. between 15 and 64 years of age) for both native and immigrants, and on their average ages are instead drawn from the Demographic Portal of ISTAT (DP-ISTAT)[21]. In line with the existing literature on migration, we use the definition of immigrants provided by ISTAT

[18]We have used the files relative to the first quarter of every year. The reason behind this choice is that the information on the resident population (for both immigrants and natives) is relative to the 1^{st} of January of each year. Therefore using the first quarter of the LFS allow us to have a more homogeneous dataset.

[19]Specifically, we have used the "files for research purposes" version of the LFS. Fore more details, see https://www.istat.it/en/analysis-and-products/microdata-files.

[20]By "crisis" we mean the sovereign debt crisis occurred in Europe after the global financial crisis originated by the collapse of Lehman Brothers in 2008. It is important to highlight that by 2009 many Western economies were already in a recessionary phase. However, the first real signs of economic downturn started in Italy in 2010-2011 and culminated with the resignation of Berlusconi's government and the appointment of a "technical" government led by Mario Monti.

[21]See http://demo.istat.it/index_e.html.

and consider as immigrants those individuals residing in Italy without the Italian citizenship (Borjas, 2003; D'Amuri and Peri, 2014; Edo, 2015; Bettin and Sacchi, 2020).

As for the wage data, the microdata files of the LFS used report, among other things, information on both the after-tax monthly earnings[22] and on the number of hours worked in the week previous to the interview. In the empirical specification, our dependent variable is the (log) hourly wage of the natives that we compute in the following way. First, we obtain a measure of hours worked per month using information on the hours worked per week, from which we deduct the number of overtime hours[23]. Subsequently, we obtain a measure of hourly wage by dividing the after-tax monthly wage by the number of hours worked in the month. Throughout the analysis, the wage measures are computed in euros as of 2017, using the Consumer Price Index (CPI) provided by ISTAT[24].

The data on both immigrant and total populations required to build the migration regressor are drawn from the DP-ISTAT. Specifically, we have used the information on the working-age individuals residing in every administrative province (i.e. NUTS 3 regions) at the beginning of each year under analysis.

Finally, the information required to build the share-component of the instrument are derived from the number of residence permits issued by the Italian Ministry of the Interior to all foreign-born individuals in every province in 1995, distinguishing by countries of origin. The shift-component is instead built based of the stock of foreign-born individuals by countries of origin provided by the DP-ISTAT[25].

All the figures are computed as yearly changes. This has allowed us to obtain a dataset composed by 8 time periods for the 102 Italian administrative provinces. Despite the fact that these are not functional labor markets, it is important to highlight that they are geographical unit closer to the definition of local labor markets for which information on the variables in this study is available. In this regard, it is worth pointing out that the analysis has been conducted using much more spatial units than those in Dustmann et al. (2013) (that use 17 regions in the U.K.).

Table 3A1 presents some descriptive statistics of the main variables used in the analysis. Different important features emerge. First of all, in the period under analysis the average Italian province received an immigrant inflow of about 0.25 percent, figure that is quite similar to that reported for more longstanding countries of immigration (see, to this purpose, Dustmann et al., 2013)[26]. Alongside, the mean of the Bartik variable is -0.018, implying that, in the period under analysis, a representative local economy has experienced a negative labor demand shock. As for the net migration rate, its mean is almost zero, highlighting a low tendency of migrating internally for the Italian population. Finally, immigrants are, on average, younger than natives (their average age is almost 37 years, against about 41 for natives). As for the educational composition of the province population, it is possible to notice that high-educated individuals are about the half of low-educated ones, while intermediate-educated natives are instead more than low-educated ones.

[22]Such information is reported for dependent workers only (i.e. employees).Therefore, self-employed individuals are excluded from the analysis, which is a standard procedure in the extant literature (e.g. Borjas, 2003; Ottaviano and Peri, 2012; Edo and Toubal, 2017).

[23]Basically, we multiply the number of hours worked per week, net of the overtime hours, by 4.35, which is the the number of weeks in an "average" month.

[24]For more details, see `https://www.istat.it/it/archivio/14413`.

[25]We are extremely grateful to Prof. Massimiliano Bratti (University of Milano) to kindly supply these data.

[26]More precisely, Dustmann et al. (2013) estimate for the period 1997-2005 a yearly increase of the immigrants-to-natives ratio of around 0.3 percent in the U.K.

3.5 Results

In this section, we present the results of the impact of immigration on native wages. We start presenting the specifications that report the estimates of the impact of immigration on the native average hourly wages. To test the validity of our results, we present a set of robustness checks. We then move to the analysis of the impact of immigration along the native wage distribution. In the first place, we consider all the Italian provinces, then we exploit the well known North/South divide and present results separately for these two macro regions. In all these cases, we provide the results relative to all native workers together, and then distinguishing by gender.

Before presenting the results of the wage impact of immigration, it is worth highlighting that the instrument used in the IV strategy presents a fairly strong first-stage, with typically high F-statistics (see Table 3A3). This is reassuring in terms of the validity of the instrument because, even in presence of a small deviation of the exogeneity condition, the resulting bias in the estimates of the effect of interest is likely to be not large.

3.5.1 Impact of Immigration on Native Average Wages

We report the results of the estimates of the effect on average wages in Table 3.1[27]. Specifically, the first column is relative to the baseline specification, where we only control for for time-invariant unobserved heterogeneity. Overall, point estimates are positive but not statistically significant, indicating absence of correlation between immigrants and native average wages. This result, however, may be driven by the omission of relevant variables. There might indeed be province-year specific (observable) factors that affect the dynamics of native wages, and, if not controlled for, may bias the estimates of the effect of interest. To address this issue, in column (2) we control for labor demand shocks, by means of the Bartik variable and for native internal migration by including the net migration rate, as defined previously. Their inclusion, however, does not alter significantly the estimates of interest and the correlation between immigrant shocks and native average wages remains not significant.

So far we have considered immigrants as if they were exogenous to the economic conditions of the local labor markets of destination. However, it is well known that this is actually not the case, as immigrants typically settle in regions that are thriving (Borjas and Monras, 2017; Albert and Monras, 2020; Borjas and Edo, 2021). To this extent, in column (3) we estimate a 2SLS model where we instrument the change in immigrant population with the canonical shift-share variable as defined previously. Again, the point estimates are usually positive but not statistically significant. Finally, in columns (4) and (5) we further control for province-specific trends (province dummies in the specification in first-differences) to address the fact that the wage dynamics may differ between provinces. The estimates however remain similar.

In order to better disentangle this results, that is to verify whether the absence of a significant wage response to immigration also occurs for both males and females, we have replicated the empirical strategy for these two groups of natives separately. Indeed, despite the fact that the vast majority of existing studies only focus on men (Borjas, 2014; Edo,

[27] For the sake of space, Table 3.1 only reports the estimates of the coefficient of interest. The coefficients associated to the control variables are instead available in Table SM1 of the OSM of this chapter, which is available at `https://www.dropbox.com/s/ueopox7oh1s5fmh/OSM_CH3.pdf?dl=0`.

2015)[28], we do believe that distinguishing by gender may be interesting, especially in a context like the Italian one. Italy is in fact characterized by the combination of a relatively low amount of public facilities providing child- or elderly-care, a reduced flexibility in the contracts, and the presence of deep-rooted cultural norms. Because of these reasons, the female labor force participation in Italy is considerably lower than that of males or than that of other European countries (Marino and Nunziata, 2017). In such a scenario, we expect that an inflow of immigrants may substitute native women in the household service occupations, in which foreign-born workers (especially females) tend to specialize (Barone and Mocetti, 2011), and induce an increase in their labor force participation. The extant literature indicates that this mechanism may operate both at the intensive (Farré et al., 2011; Forlani et al., 2015) and extensive margins (Barone and Mocetti, 2011; Cortés and Tessada, 2011). The results relative to the gender distinction are reported in the corresponding rows in Table 3.1. However, estimates indicate that, for both native men and women, the wage response to immigration is not significantly different from zero.

All in all, in line with the previous literature (e.g. Aubry et al., 2021), these results imply that immigrants did not exerted any significant impact on the average wages of their native counterparts in Italy over the period analyzed.

Table 3.1: Effect of immigration on average wages. All provinces.

	(1) OLS	(2) OLS	(3) 2SLS	(4) OLS	(5) 2SLS
All Workers	0.164	0.124	0.197	0.055	-0.128
	(0.403)	(0.416)	(0.456)	(0.406)	(0.578)
Men	0.310	0.273	0.189	0.238	-0.095
	(0.499)	(0.521)	(0.483)	(0.489)	(0.602)
Women	0.032	-0.005	0.268	-0.112	-0.082
	(0.393)	(0.401)	(0.518)	(0.416)	(0.634)
Controls	NO	YES	YES	YES	YES
Province trends	NO	NO	NO	YES	YES
First-stage F-statistics	-	-	163.3	-	75

Note: *The table reports the estimates of the impact of immigration on average wages. The dependent variable is the conditional average hourly wage, while the main independent one is the change in immigrant population, standardized by the yearly-lagged total population in working-age. Controls include the Bartik variable and the net migration rate. In all specifications the number of observation is 816 and the first-stage F-statistics are always well above the 10% maximal IV size critical value of the Stock and Yogo (2005) weak ID test. All regressions include year fixed-effects and are weighted by the working-age population of each province-year cell at the beginning of the period. Standard errors, in parentheses, are clustered at the province level. All wage measures are expressed in euros as of 2017, using the 2017 Consumer Price Index provided by ISTAT.*
*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Finally, to be sure that this conclusion is not driven by the choices of the empirical specification made in section 3.3, we also perform a set of robustness tests, whose 2SLS results are presented in Table 3A5. Specifically, in column (1), we perform the unweighted regressions. This allows us to verify whether the results of Table 3.1 are driven by the weighting scheme chosen[29]. In column (2) we use a different measure of the immigrant supply shock, namely the change in the immigrants-to-natives ratio (used, for instance, in Dustmann et al., 2013). In column (3) we substitute the control variables with those used in Dustmann et al. (2013), which capture changes in the composition of the local population. These are the change in: (i) the log of the ratio between high- and low-educated natives;

[28] A notable exception is the recent study of Borjas and Edo (2021).
[29] In addition, we also address the concerns on the use of regression weights (Solon et al., 2015).

(ii) the log of the ratio between intermediate- and low-educated natives[30]; (iii) immigrants' average age; and (iv) natives' average age. In column (4), to control for the local employment structure, we add as controls the changes in the employment shares in white collars, skilled-manual, and blue collars (this last is the reference category). In column (5), we follow the guidelines of Borjas and Edo (2021) and add the change in the (log) native workforce, that controls for the size of the local labor market[31]. Lastly, in column (6) we drop from the analysis immigrants coming from EU 15 countries, and in column (7) we perform the analysis excluding the larger provinces (that is, those with more that 2 million inhabitants, namely Milano, Rome and Naples). Reassuringly, even in these cases the absence of any significant impact of immigrants on native wages is confirmed.

3.5.2 Impact of Immigration along the Native Wage Distribution

If, on the one hand, the analysis on the average wage may be interesting because it provides some insights on the overall effect of immigration, on the other it may hide the presence of heterogeneity in the native response. Instead, an analysis on the wage distribution is more accurate as it allows to better identify the potential "winners" and "losers" from immigration (if any). Because of these reasons, in this section we present the results relative to the impact of immigration along the native wage distribution.

To this extent, Table 3.2 reports the results of the estimates for the deciles of the conditional wage distribution relative to all provinces. We initially report the estimates of the effect of interest for the whole native population and obtained without the inclusion of province-specific trends (column 1). As can be seen, the coefficients associated to the first two deciles are negative, but not precisely estimated. Then, as long as we move along the wage distribution, point estimates increase in magnitude and, starting from the median, become statistically significant (with the exception of the coefficient associated to the 7^{th} decile). According to these estimates, a unitary percentage increase in the change of immigrants population leads to an increase in native wages that varies from 0.55 percent at the median, up to 1.16 percent at the 9^{th} decile. However, when we additionally control for province-specific trends (column 2), the precision of the estimates is considerably reduced. Specifically, the only significant coefficient is the one associated to the 8^{th} decile and implies that a one percentage increase in the change of immigrants population is associated to an increase in native wages by about 1 percent.

As in the case of average wages, it is of particular interest to verify whether the results for the full native sample are confirmed when considering men and women separately. To this extent, we replicate the general empirical analysis but distinguishing by gender. The corresponding estimates are reported in columns (3) to (6) of Table 3.2. Specifically, in the case of men, when we do not control for province-specific trends (column 3), the results identify a negative though not significant impact in the lower tail of the wage distribution and a positive significant effect in the upper part. These findings are somehow confirmed when we additionally control for province-trends (column 4), although now the positive effect estimated in the upper tail of the wage distribution is only significant at 10% level. A similar pattern is also found in the case of women. More in detail, in column (5) we observe a positive wage response, particularly from the median of the wage distribution onwards

[30]These first two measures allow to control for human capital externalities within each province.

[31]We want to stress that this variable is very likely endogenous, as natives respond to wage dynamics by settling to regions that offer better economic opportunities. However, the treatment of the endogeneity of the native workforce is beyond the scope of this analysis and is left to future research.

(with the exception of the 9^{th} decile). Point estimates indicate that a unitary percentage increase in the change of immigrants population is associated with an increase in native women wages that varies from 0.88 percent at the median, up to 1.53 percent at the 8^{th} decile. Again however, the inclusion of the province trends reduces quite substantially the precision of the estimates (column 6). More in detail, the coefficients of interest are only significant at 10% level and vary between 0.88 (6^{th} decile) to 1.20 (8^{th} decile)[32].

Table 3.2: Effect of immigration along the conditional wage distribution. All provinces.

	All Workers		Men		Women	
Quantile	(1)	(2)	(3)	(4)	(5)	(6)
1^{st} decile	-0.476	-0.927	-0.436	-0.787	-0.540	-1.021
	(0.615)	(0.834)	(0.777)	(0.867)	(0.802)	(0.944)
2^{nd} decile	-0.180	-0.420	0.071	-0.238	-0.128	-0.492
	(0.549)	(0.537)	(0.600)	(0.585)	(0.602)	(0.591)
3^{rd} decile	0.341	0.140	0.223	0.064	0.102	-0.228
	(0.286)	(0.424)	(0.368)	(0.487)	(0.463)	(0.548)
4^{th} decile	0.445	0.208	0.054	-0.143	0.602	0.316
	(0.337)	(0.405)	(0.321)	(0.405)	(0.438)	(0.525)
5^{th} decile	0.545^{*}	0.316	0.340	0.199	0.882^{**}	0.527
	(0.322)	(0.379)	(0.333)	(0.367)	(0.440)	(0.521)
6^{th} decile	0.725^{***}	0.516	0.564^{*}	0.443	1.054^{**}	0.884^{*}
	(0.282)	(0.406)	(0.341)	(0.508)	(0.434)	(0.517)
7^{th} decile	0.712	0.461	1.069^{**}	0.966^{*}	1.285^{**}	0.993^{*}
	(0.486)	(0.492)	(0.509)	(0.565)	(0.628)	(0.590)
8^{th} decile	1.128^{***}	0.990^{**}	1.033^{**}	0.918^{*}	1.526^{**}	1.198^{*}
	(0.382)	(0.456)	(0.481)	(0.512)	(0.615)	(0.632)
9^{th} decile	1.156^{***}	1.019	0.778	0.744	0.471	0.428
	(0.358)	(0.642)	(0.476)	(0.648)	(0.697)	(0.825)
Province trends	NO	YES	NO	YES	NO	YES

Note: *The table reports the 2SLS estimates of the impact of immigration along the native wage distribution, for all the Italian provinces. The dependent variables are the different deciles of the conditional hourly wage, while the main independent one is the change in immigrant population, standardized by the yearly-lagged total population in working-age. All regressions include as controls the Bartik variable and the net migration rate, as well as year and province fixed-effects, and are weighted by the working-age population of each province-year cell at the beginning of the period. In all cases the number of observation is 816. The first-stage F-statistics equals 163.3 in the specifications without province trends, and 75 in those with province trends. In both cases, they are well above the 10% maximal IV size critical value of the Stock and Yogo (2005) weak ID test. Standard errors, in parentheses, are clustered at the province level.*
*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

Taken together, these results have two interesting implications. The first is that immigration does not exert a significant downward pressure on native wages, not even in the case of the least skilled natives (that is, those located in the lower tail of the conditional wage distribution), who are typically expected to experience wage losses in reaction to the inflow of immigrants in the labor market. Second, if anything, we estimate a positive effect, that in any case could be only marginally significant, in the upper part of the conditional wage distribution. This might confirm the presence of a complementarity in production between immigrants and high-skilled natives in the Italian provinces and in the period analyzed,

[32] The drop in the precision in the estimates associated to the 1^{st} and 9^{th} decile may be due to the fact that the wage data provided by ISTAT are bottom and top coded, to guarantee the anonymity of the respondents. We somehow address this issue by trimming in the 1^{st} and 99^{th} percentile from the native wage distribution in every province and year, as done in Dustmann et al. (2013).

3.5.3 Spatial Heterogeneity: The North/South Divide

Another source of heterogeneity in the native response to immigration may depend on the different industrial composition of distinct local labor markets. Specifically, the extent to which an immigration-induced supply shift is absorbed may considerably vary between regions characterized by more dynamic labor markets and more technology-intensive industries, from others characterized by poorer labor market conditions and more labor-intensive industries.

Given the previously described North/South divide of Italy, taking into account this distinction may be of extreme interest. Indeed, the extent to which Northern and Southern regions react to an immigration-induced supply shock may be different, and a general analysis as the one presented in the previous section may conceal such heterogeneity. Along these lines then, in this section we perform the empirical analysis previously described allowing for a different impact of immigration on the native wage distribution between Northern and Southern provinces. Furthermore, we additionally distinguish between the impact for male and female native workers in both Northern and Southern provinces.

We initially briefly present the results relative to the average wages, that are reported in Table 3.3. In line with the results of Table 3.1, both in the case of Northern and Southern provinces, the estimated effects are typically not statistically significant. This is valid for the full native sample, as well as for men and women separately. Interestingly, despite the absence of statistical significance in both set of provinces, the point estimates are negative in the case of men, while positive in the case of women.

Table 3.3: Effect of immigration on average wages. Spatial heterogeneity.

	(1) OLS	(2) OLS	(3) 2SLS	(4) OLS	(5) 2SLS
			Northern Provinces		
All Workers	0.107	0.052	0.040	0.070	-0.136
	(0.417)	(0.427)	(0.538)	(0.425)	(0.637)
Men	0.261	0.215	-0.182	0.260	-0.313
	(0.525)	(0.541)	(0.577)	(0.524)	(0.658)
Women	-0.023	-0.075	0.419	-0.090	0.198
	(0.399)	(0.411)	(0.673)	(0.440)	(0.733)
First-stage F-statistics	-	-	81.2	-	54.3
			Southern Provinces		
All Workers	-0.020	-0.104	0.286	0.217	-0.180
	(0.886)	(0.889)	(1.748)	(1.180)	(1.964)
Men	0.289	0.217	-1.187	0.486	-1.565
	(1.257)	(1.281)	(2.347)	(1.570)	(2.243)
Women	-0.115	-0.193	2.411	0.130	1.808
	(1.254)	(1.257)	(2.309)	(1.598)	(2.411)
First-stage F-statistics	-	-	28.4	-	34.7
Controls	NO	YES	YES	YES	YES
Province trends	NO	NO	NO	YES	YES

***Note**: The table reports the estimates of the impact of immigration on average wages. The dependent variable is the conditional average hourly wage, while the main independent one is the change in immigrant population, standardized by the yearly-lagged total population in working-age. Controls include the Bartik variable and the net migration rate. In all specifications the number of observation is 816 and the first-stage F-statistics are always well above the 10% maximal IV size critical value of the Stock and Yogo (2005) weak ID test. All regressions include year fixed-effects and are weighted by the working-age population of each province-year cell at the beginning of the period. Standard errors, in parentheses, are clustered at the province level.*
*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

The impact of immigration along the native wage distribution is instead reported in Tables 3.4 (Northern provinces) and 3.5 (Southern provinces). Starting from the former, when considering the full native sample in the specification that does not control for province-specific trends (column 1), we observe that in the lower tail of the conditional wage distribution the estimated effects are not statistically significant. Conversely, from the 6^{th} decile onwards we identify a positive wage response that hovers between 0.9 to 1 percent. However, once again, after controlling for province trends, both the magnitude and the precision of the estimates are reduced, and only the coefficient associated to the 8^{th} decile remains statistically significant (at 10% level). As for the gender distinction, in the case of native men (columns 3 and 4) the wage impact of immigration is insignificant all along the wage distribution, both in the estimations with or without province trends. Things are slightly different in the case of women, Specifically, when we do not control for province-specific trends (column 5), the wage response is positive although clearly not significantly different from zero in the the lower tail of the wage distribution. However, in the upper part immigrant inflows exert a positive significant impact. Point estimates imply that a unitary percentage increase in the immigrant population is associated with an increase in female wages of about 1.3 to 1.5 percent from the 6^{th} to the 8^{th} deciles. Once controlling for province specific trends, the positive effects identified in the upper part of the wage distribution for native females in the Northern provinces still hold (column 6), although both the magnitude and the significance of the parameter estimates are somehow reduced.

Table 3.4: Effect of immigration along the conditional wage distribution. Northern provinces.

	All Workers		Men		Women	
Quantile	(1)	(2)	(3)	(4)	(5)	(6)
1^{st} decile	-0.583	-0.808	-1.119	-1.293	0.368	0.031
	(0.680)	(0.957)	(0.815)	(0.997)	(0.953)	(1.129)
2^{nd} decile	-0.154	-0.244	-0.585	-0.623	0.380	0.091
	(0.625)	(0.632)	(0.640)	(0.644)	(0.769)	(0.742)
3^{rd} decile	0.348	0.259	-0.121	-0.161	0.336	0.152
	(0.351)	(0.513)	(0.440)	(0.571)	(0.563)	(0.688)
4^{th} decile	0.389	0.299	-0.022	-0.046	0.626	0.463
	(0.410)	(0.506)	(0.440)	(0.519)	(0.528)	(0.642)
5^{th} decile	0.618	0.505	0.407	0.406	1.138^{**}	0.859
	(0.426)	(0.495)	(0.477)	(0.493)	(0.572)	(0.669)
6^{th} decile	0.962^{**}	0.846	0.659	0.653	1.291^{**}	1.190^{*}
	(0.397)	(0.528)	(0.463)	(0.630)	(0.520)	(0.640)
7^{th} decile	0.765	0.618	1.034	1.036	1.527^{*}	1.349^{*}
	(0.604)	(0.607)	(0.650)	(0.684)	(0.796)	(0.744)
8^{th} decile	1.056^{**}	1.061^{*}	0.816	0.825	1.343^{*}	1.204
	(0.484)	(0.564)	(0.593)	(0.626)	(0.715)	(0.744)
9^{th} decile	0.921^{*}	1.021	0.551	0.734	-0.051	0.068
	(0.493)	(0.756)	(0.691)	(0.781)	(0.820)	(0.945)
Province trends	NO	YES	NO	YES	NO	YES

***Note**: The table reports the 2SLS estimates of the impact of immigration along the native wage distribution, for the group of Italian Northern provinces. The dependent variables are the different deciles of the conditional hourly wage, while the main independent one is the change in immigrant population, standardized by the yearly-lagged total population in working-age. All regressions include as controls the Bartik variable and the net migration rate, as well as year and province fixed-effects, and are weighted by the working-age population of each province-year cell at the beginning of the period. In all cases the number of observation is 816. assume the values of 81.2 in the specifications without province trends, and 54.3 in those with province trends. In both cases, they are well above the 10% maximal IV size critical value of the Stock and Yogo (2005) weak ID test. Standard errors, in parentheses, are clustered at the province level.*

*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

In the case of the Southern provinces the wage responses are instead estimated with lower precision, implying that the estimates of the effect of interest are not statistically significant. Specifically, in the case of the whole native population, when we do not control for province-specific trends (column 1 of Table 3.5), the only positive coefficient is the one associated to the 6^{th} decile and indicates that immigration shocks induce an increase in wage of around 3.3 percent. However, when we add province trends (column 2) no coefficients remain significant. As for Northern provinces, in the case of the subsample of native men the estimates of the effect of immigration are always not statistically significant, irrespective of the inclusion or not of province-specific trends. Finally, in the case of women, in the specification without country trends, we estimate a positive effect in the lower part of the wage distribution (first two deciles) that imply strong wage gains that vary between 4.4 to 7 percent (column 5). This is consistent with the presence of downward wage rigidity and to the upgrade of natives to more skilled occupations (i.e. low skilled job losses replaced by more skilled job hires), as indicated by Peri and Sparber (2009) for the U.S. or D'Amuri and Peri (2014), Cattaneo et al. (2015) and Foged and Peri (2016) in the case of Europe. Conversely, when including province trends, all coefficients become not anymore statistically significant.

Table 3.5: Effect of immigration along the conditional wage distribution. Southern provinces.

	All Workers		Men		Women	
Quantile	(1)	(2)	(3)	(4)	(5)	(6)
1^{st} decile	0.244	-0.125	-3.836	-4.202	6.978**	6.079
	(2.252)	(2.656)	(3.156)	(3.105)	(3.451)	(3.713)
2^{nd} decile	1.157	0.763	-2.718	-2.837	4.439**	3.444
	(1.871)	(1.985)	(2.427)	(2.300)	(2.248)	(2.745)
3^{rd} decile	1.435	0.945	-1.134	-1.459	2.871	2.339
	(1.310)	(1.793)	(1.885)	(2.019)	(2.585)	(2.846)
4^{th} decile	1.228	0.820	0.656	0.513	1.974	1.310
	(1.318)	(1.822)	(1.906)	(2.041)	(2.099)	(2.400)
5^{th} decile	2.042	1.595	2.124	1.599	3.514	2.775
	(1.403)	(1.843)	(1.961)	(1.971)	(2.523)	(2.632)
6^{th} decile	3.253**	2.747	2.124	1.859	3.460	2.953
	(1.613)	(2.008)	(1.898)	(2.174)	(2.724)	(2.558)
7^{th} decile	2.077	1.522	1.752	1.440	3.964	3.396
	(1.793)	(2.010)	(2.162)	(2.229)	(3.216)	(2.751)
8^{th} decile	1.650	1.474	0.564	0.295	1.596	1.242
	(1.728)	(1.966)	(2.370)	(2.367)	(2.353)	(2.514)
9^{th} decile	0.795	1.033	0.265	0.675	-1.874	-2.002
	(1.998)	(2.235)	(3.153)	(2.947)	(2.715)	(2.926)
Province trends	NO	YES	NO	YES	NO	YES

***Note**: The table reports the 2SLS estimates of the impact of immigration along the native wage distribution, for a representative Italian Southern province. The dependent variables are the different deciles of the conditional hourly wage, while the main independent one is the change in immigrant population, standardized by the yearly-lagged total population in working-age. All regressions include as controls the Bartik variable and the net migration rate, as well as year and province fixed-effects, and are weighted by the working-age population of each province-year cell at the beginning of the period. In all cases the number of observation is 816. The first-stage F-statistics assume the values of 28.4 in the specifications without province trends, and 34.7 in those with province trends. In both cases, they are well above the 10% maximal IV size critical value of the Stock and Yogo (2005) weak ID test. Standard errors, in parentheses, are clustered at the province level.*
*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

To summarize, the results of this section have some interesting implications. First, the overall wage response to immigration is non-negative, as instead often claimed by some political actors. Moreover, this is valid also for native workers located in the lower part of the wage distribution, even for those residing in the Southern provinces. This results is somewhat surprising, as one might instead expect that immigration induces wage drops for low-skilled natives. Second, we estimate positive wage effects in the upper part of the wage distribution. This appear to be mostly driven by native female workers residing in the Northern provinces, since the positive effects estimated for females in the Southern provinces are not statistically significant. This might be due to two main reasons. On the one hand, the labor supply of women is more elastic than that of men (that is, women are typically more responsive to labor supply shock that men). On the other, high-skilled women are more positively affected by immigrant inflows (which, as indicated in Figure 3A2, are mostly composed by women) because they can be replaced by low-skilled immigrants in household production services, with a positive effect on their wages.

3.6 Conclusions

The resurgence of the migratory phenomenon that occurred in the last few years in many Western economies has caused a revival of the discussion of the socio-economic consequences of immigration. With this chapter we contribute to the literature by assessing the distributional impact of immigration in the Italian context, which, for different reasons, represents an interesting case of study. Indeed, its recent shift from immigrant-sending to immigrant-receiving country, its peculiar institutional framework characterized by rigid product and labor markets, make this study particularly appealing, especially in a period (i.e. 2009-2017) characterized by the combination of the further increase of the migratory inflows and by the outbreak of the economic crisis.

The hypothesis that we wanted to test is that the fact that Italy mostly attracts low-skilled workers, combined with the rigidity in the downward wage adjustments due to its institutional setting, would imply a non-negative effect on both average wages and along the whole wage distribution. In addition, the skill complementarity between immigrants and native workers, may suggest to expect a positive effect in the upper end of the wage distribution.

In order to test the hypotheses, we have implemented a two-steps methodology that allows us to account for both changes in the composition of the native workforce and the non-random settlement of immigrants across provinces. Our findings appear to survive to a set of robustness checks and are in line with both the initial hypotheses and the existing literature on Italy (Gavosto et al., 1999; Staffolani and Valentini, 2010). Specifically, we identify a positive impact of immigration along the native conditional wage distribution, in particularly from the middle of the distribution onwards. Point estimates indicate that a one percent increase in the immigrant population leads to an increase in native wages that vary between 0.55 to 1.16 percent. Moreover, in order to better disentangle the origins of this positive effect, we have performed a heterogeneity analysis along two lines. First, we have distinguished the native population by gender. Then, we have exploited the well-known north-south divide that characterizes Italy since its very foundations. Interestingly, these analyses suggest that the general results could be mostly driven by native female workers with medium and high skills, and residing in the North of the country.

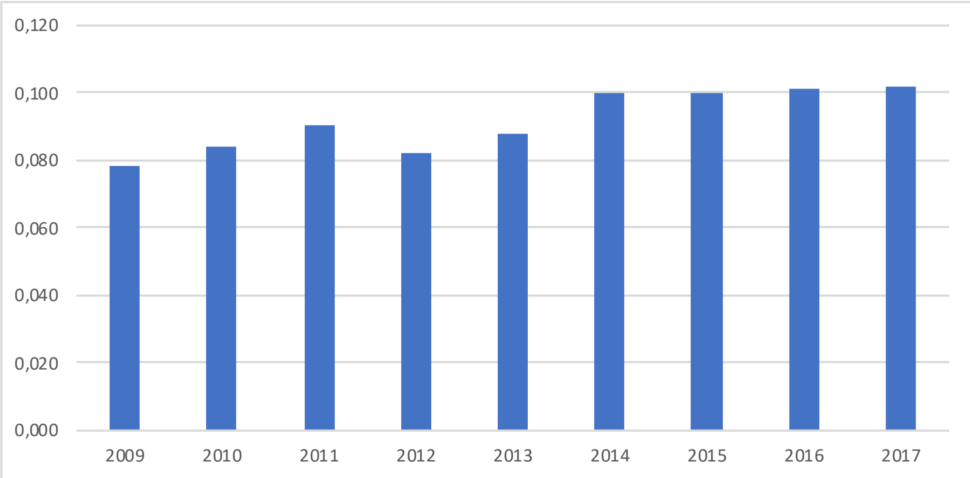
All in all, this positive relationship seems to be due to the presence of a complementarity in production between natives and immigrants, as predicted for the Italian case by Romiti (2011). As in the case of rural migrants in China (Combes et al., 2015), this effect is stronger for high-skilled natives, particularly females (in line with Barone and Mocetti, 2011). In the case of low-skilled natives the impact that immigrants exert seems to be insignificant. This is probably due to the fact that low-skilled natives and immigrants are not closer substitute and to the presence of strong wage rigidity determined by the Italian institutional framework. A negative aspect that emerge is that, if anything, immigrants seem to contribute to an increase in the wage inequality of natives due to the higher positive impact at the top of the distribution.

We believe that these findings have different implications. First, the analyses of the wage impact of immigration should allow for heterogeneous impacts for natives with different skills. This also allows to capture potential differences in the response by gender or region that may derive from different elasticities of substitution between skills. Second, it is also necessary to consider that natives do not distribute uniformly across regions and neglecting this spatial sorting mechanism biases the estimates of the wage impact of immigration. Third, against extended opinion, this evidence suggests that recent immigration into Italy has not depressed the wages of natives, not even of those with the lowest skills, potentially more affected by competition. Fourth, the impact varies by gender and between North and South. Based on the arguments presented in Dustmann et al. (2013), this can be explained by a highest complementarity between the high-skills of female natives and the low skills of recent immigrants to Italy. This is particularly so in the more developed Northern provinces.

To finally conclude, we do believe that is important to acknowledge the drawbacks of this study. First, the relatively limited availability of data only allows us to estimate the partial (i.e. short-run) wage effects of immigration, but not the long-run effects. Second, the repeated cross-sectional structure of the microdata used did not allow us to control for workers unobserved characteristics. Third, the IV estimates show large standard errors, particularly in the spatial heterogeneity analysis. A potential solution could be to use alternative or complementary instrument that could increase the amount of information used in the IV strategy, reducing in this way the standard errors of the estimated effects. In future lines of research, it could be extremely interesting to address all these questions.

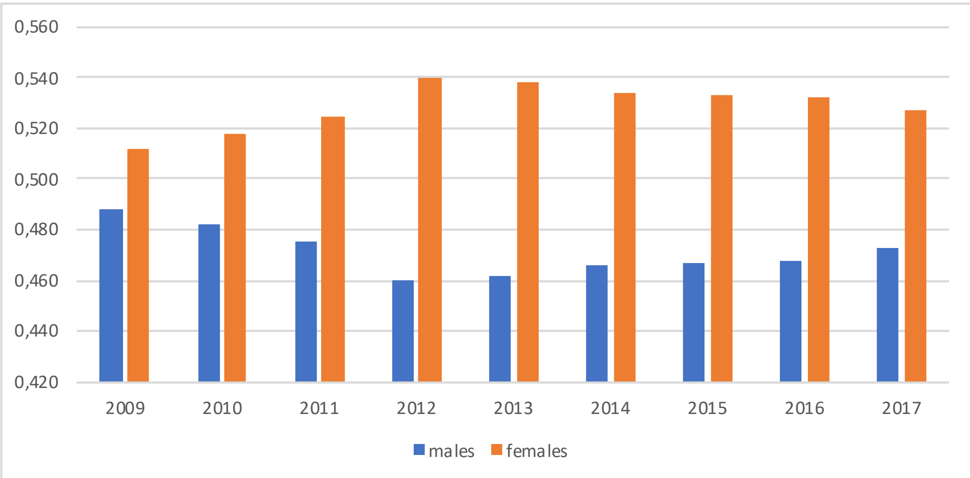
3.7 Appendix

Figure 3A1: Immigrant share over the total population (working-age individuals).



Note: *The figure reports the evolution over the period under analysis of the immigrant share over the total working age population.* ***Source:*** *DP-ISTAT.*

Figure 3A2: Gender composition of the immigrant population.



Note: *The figure reports the evolution over the period under analysis of the gender composition of the immigrant population* ***Source:*** *DP-ISTAT.*

Table 3A1: Descriptive statistics.

Variable	Mean	S.D.	Min.	Max.
Native average wages				
All provinces	2.2448	0.0745	2.0199	2.5654
Northern provinces	2.2548	0.0676	2.0544	2.5654
Southern provinces	2.2265	0.0827	2.0198	2.5249
Immigrant measures				
Change immigrant population	0.0025	0.0066	-0.0289	0.0393
Change immigrant-native ratio	0.0033	0.0069	-0.0304	0.0371
Controls				
Bartik variable (wage)	-0.0185	0.0348	-0.0632	0.1000
Internal migration	-0.0765	3.1367	-11.2624	7.9068
Natives' average age	41.4648	1.0664	38.5320	43.8006
Immigrants' average age	36.8774	1.0769	33.8902	39.9180
Log high/low educated natives	-0.6109	0.3993	-1.9667	0.6351
Log interm./low educated natives	0.3312	0.2940	-1.1513	1.3583
Employment shares				
White collars	0.1675	0.0402	0.0464	0.3736
Skilled manual	0.6696	0.0489	0.4375	0.8434
Blue collars	0.1629	0.0383	0.0488	0.3226

Note: *Mean, standard deviation, minimum and maximum of the variables used in the analysis, using the observations of all provinces and years in the sample.* ***Source***: *LFS 2009-2017 and ISTAT.*

Table 3A2: Overview of the variables used in the first-step.

Variable	Definition	Categories
Experience:		
Age	Age of each individual	-
Age squared	Squared age of each individual	-
Education	Level of education	primary, secondary and tertiary
Occupation	Type of occupation	white collar, skilled manual and blue collar
Contract information:		
Length	Length of the contract	temporary vs permanent
Type	Type of the contract	full time vs part time
Gender	Sex of each individual	Males vs Females

The table reports the description of the variables used in the first-step of the empirical strategy as described in the text. ***Source***: *LFS 2009-2017 and ISTAT.*

Table 3A3: First-stage estimates

	(1)	(2)	(3)
$\Delta(\widehat{m}_{p,t})$	0.576***	0.590***	0.649***
	(0.041)	(0.046)	(0.075)
Controls	NO	YES	YES
Province trends	NO	NO	YES
R^2	0.844	0.849	0.888
F-statistic	192.1	163.3	75
Observations	816	816	816

Note: *The dependent variable is the change in immigrant population as share of the initial working-age population, while the main independent variable is the change in the shift-share instrument based on residence permits issued in 1995.. Column (1) reports the first-stage estimates relative to the baseline specification, in column (2) we add as controls the Bartik and internal migration variables, and finally in column (3) we further include province-specific trends. All regressions include year fixed-effects and are weighted by the total number of working-age individuals in the province at the beginning of the period. Standard errors, in parentheses, are clustered at the province level.*
*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

Table 3A4: Instrument validity

	Panel A			
	Change employment		Change wages	
	2009-2010 (1)	2009-2013 (2)	2009-2010 (3)	2009-2013 (4)
Share component instrument	-0.034	0.045	-0.029	0.030
	(0.029)	(0.032)	(0.020)	(0.027)
	Panel B			
Change instrument 2009-2017	-0.021	0.024	-0.012	0.011
	(0.049)	(0.044)	(0.019)	(0.029)
Change instrument 2013-2017	-0.019	0.011	-0.011	0.009
	(0.036)	(0.032)	(0.014)	(0.020)

Table 3A5: Robustness checks.

	(1) unw. reg.	(2) other imm. shock	(3) other controls	(4) emp. shares	(5) Δ (log) natives	(6) No EU15 imm.	(7) No large prov
All Workers	0.822	-0.100	-0.156	-0.113	-0.138	-0.117	0.899
	(0.810)	(0.702)	(0.588)	(0.546)	(0.795)	(0.526)	(0.906)
Only Men	1.021	-0.041	-0.239	-0.076	-0.054	-0.087	1.637
	(0.930)	(0.730)	(0.610)	(0.592)	(0.822)	(0.549)	(1.021)
Only Women	0.614	-0.074	0.007	-0.073	-0.104	-0.075	0.099
	(0.988)	(0.743)	(0.655)	(0.618)	(0.858)	(0.576)	(1.153)
F-statistics	103.7	87.6	83.5	75.1	327.1	61.9	131.9

Note: *The table reports the estimates of the different robustness tests as indicated at the top of each column. All regressions include as controls the Bartik variable and the net migration rate, year and province fixed-effects and province trends, and are weighted by the total number of working-age individuals in the province at the beginning of the period. The number of observations is 816 in columns (1) to (6) and 792 in column (7). Standard errors, in parentheses, are clustered at the province level.*
*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

The Online Supplemental Material of this chapter is available at: https://www.dropbox.com/s/ueopox7oh1s5fmh/OSM_CH3.pdf?dl=0

Chapter 4

Immigration, Local Specialization in Low-skilled Activities and Native Education. Evidence from some EU Countries

4.1 Introduction

Over the past few decades, the phenomenon of immigration experienced an unprecedented increase. According to the United Nations, the number of immigrants in the world increased from around 173 million people in 2000 to 258 million in 2017. In terms of their geographical distribution, Asia hosts 80 million migrants and is followed very closely by Europe with 78 million. Relatively far behind is North America, hosting 58 million individuals (United Nations, 2017). In a globalized world, immigration is an important socio-economic phenomenon and represents a challenge not only for immigrants themselves, but also for the host economies. The understanding of its consequences has become increasingly crucial, both from a social and also from a purely economic perspective, as it may trigger economic growth (Boubtane et al., 2016)[1] and induce welfare gains to the receiving economies (Clemens, 2011).

A large amount of studies have taken advantage of this large increase in the migratory inflows and analyzed the impact of immigration in the host countries, focusing in particular on native labor market outcomes, such as wages and employment (e.g. Card, 1990; Borjas, 2003; Dustmann et al., 2013; Monras, 2020; Hunt, 1992; Dustmann et al., 2005; Borjas and Monras, 2017; Glitz, 2012), internal migration (Borjas, 2006; Card and DiNardo, 2000; Card, 2001; Monras, 2021; Peri and Sparber, 2011), or other magnitudes, such as innovation (Hunt and Gauthier-Loiselle, 2010; Bratti and Conti, 2018), or housing prices (González and Ortega, 2013; Saiz, 2003, 2007)[2]. Less attention has been instead devoted to the potential

[1] However, related to this last point, the recent work of Borjas (2019) indicates that in the case U.S. the link between immigration and economic growth is far from evident. What is more clear is the strong contribution of immigrants to aggregate output. In the same spirit, Charles (2021) shows that for a group of European countries the relationship between immigration and economic growth is far from clear cut.

[2] For an extensive review on the large literature on the labor market impact of immigration, see Kerr and Kerr (2011) and Edo (2019).

mechanisms of adjustment that host countries may enact to reduce the potentially negative impact of immigration that the canonical theoretical models would predict[3]. Education is one of these. To this extent, immigration is thought to have a two-fold effect on the education of the native population. First, what may be defined as *direct effect*, that is, the impact that immigrants students exert on the school performances of their native counterparts. In this respect, there is a relatively widespread belief that the presence of immigrants tends to reduce the resources available for the natives and may create negative externalities that worsen the quality of education (also because immigrants generally have a poorer knowledge of the language spoken in the host country). This has been proven to be the case in both the U.S. (Betts, 1998; Chin et al., 2013) and also in some European countries, such as Italy (Frattini and Meschi, 2019). As a consequence, natives may decide (or be forced) to drop-out from school or, in general, to invest less in education.

On the other hand, however, an inflow of immigrants, particularly if low-skilled, may alter the skill distribution of the host country by increasing the supply of low-skilled workers (Dustmann et al., 2008; Dustmann and Glitz, 2011; Lewis and Peri, 2015). In the short-run, this may reduce both wages and employment prospects of the low-skilled natives and, in presence of complementarity in production between low- and high-skilled workers, increase those of the high-skilled. As a consequence, the return to skills increases and therefore natives are induced to spend more years at school in order to acquire those abilities that are "complements (rather than substitutes) of immigrant skills" (Lewis and Peri, 2015). In other words, migration may affect the average levels of education in the host country and thus generates educational externalities and new incentives for human capital investments for the native population. This may be defined as *indirect effect* of immigration on education. To this purpose, Figure 4.1 shows the relationship between the share of natives with upper-secondary education and immigration in the countries under analysis in this study (i.e. Greece, Ireland, Portugal and Spain) over the period 1981-2011. The line is indeed positively sloped, indicating the presence of a positive correlation between the two magnitudes.

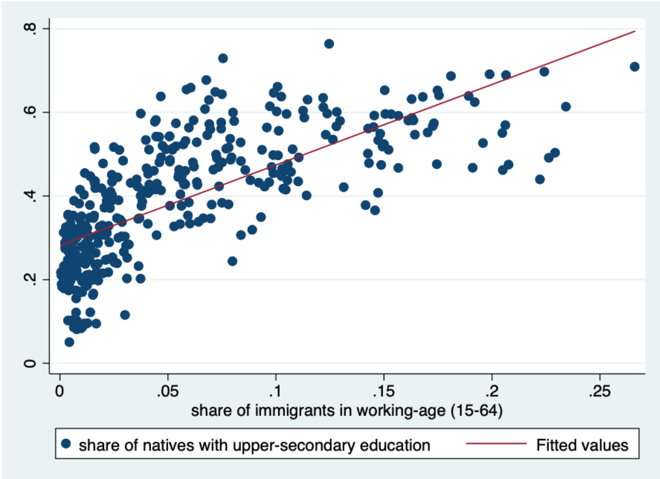
Together with the presence of skilled or unskilled immigrants, another aspect that typically influences the propensity of natives to invest in education is the industrial composition of the regions where they live (Diaz-Serrano and Nilsson, 2020). More precisely, natives residing in regions specialized in high-skilled sectors may have more incentives to invest in education, as they may want to acquire those skills required to obtain a job position in such sectors. On the contrary, in regions characterized by the presence of low-skilled industries, natives may decide to drop-out from school and start working, particularly if the wage offered in those sectors is higher than the expected return to schooling[4] (Gutiérrez-Domènech, 2011). However, the other way around is also true. The literature on technical change indeed indicates that the endowment of the factors of production of an economy (or region) directly affects the capital-labor ratio of the same, as well as its technology adoption (Acemoglu, 1998, 2002, 2010). A disproportionate increase of one factor type, e.g. low-skilled workers, may induce firms to adjust their capital-labor ratio in a way that may foster the productivity of low- or rather high-skilled workers. Which factor type experience the higher productivity gains ultimately depends on the elasticity of substitution between low- and high-skilled workers. If these are imperfect substitutes, as the existing literature seems

[3]Notable exceptions are the analyses of D'Amuri and Peri (2014), Foged and Peri (2016), Ottaviano et al. (2013), Peri and Sparber (2009) and Sparber and Zavodny (2020) that show how natives avoid (low-skilled) immigrants' competition by moving to communication-intensive type of occupations.

[4]As it happened in Spain during the property boom.

to suggest (Ciccone and Peri, 2005; Mollick, 2011), a low-skilled biased labor supply shock may induce firms to adjust their capital-labor ratio in favor of the latter (i.e. employ of more low-skilled labor and less capital) by adopting technologies that make prevalent use of low-skilled labor (Lewis, 2011; Clemens et al., 2018).

Figure 4.1: Correlation between native education and immigration



Note: *Correlation between the share of natives with upper-secondary education and the immigrants share over the total population in the regions of Greece, Ireland, Portugal and Spain over the period 1981-2011.*

In this context, the objective of this study is two-fold. On the one hand, the present chapter explores the effects of immigration on native education in the EU context, particularly in countries that become recipients of immigrants in the last decades. In this sense, the hypothesis to test is that immigration have an impact on the decisions of natives to acquire human capital. On the other hand, as a novel contribution to the extant literature, it also analyzes the effect of the interaction between immigration and specialization of the local economies in low-skilled activities. In this regard, the hypothesis to test is that the native education response to immigration is mediated by the specialization in low-skilled sectors.

The hypotheses are tested for a specific group of European countries, Portugal, Ireland, Greece and Spain. The rationale behind the choice of these countries is based on different peculiarities that characterize them[5]. First, they quite suddenly changed their role from immigrant-sending to immigrant-receiving countries, also because (i) their geographical position, and (ii) their colonial past (particularly Portugal and Spain). Second, they are characterized by strong spatial disparities that are driven by differences in: (i) the economic performance of the local labor markets that underly particularly sharp discrepancies in their productive structure; (ii) the skill distribution of the local population between regions; and (iii) the spatial distribution of the immigrant population, together with the high variability in the countries of origin distribution. Third, they have different education systems that represent an interesting source of variation to be exploited empirically.

Using the IPUMS census samples relative to the period 1981-2011 (Minnesota Population Center, 2019), I build on the two-steps strategy implemented by Hunt (2017). Specifically, in the first step, I estimate a set of country-specific regional schooling variables adjusted for

[5]Despite the availability of data, I decided not to include the countries that recently joined the EU, such as Hungary, Poland and Romania. Indeed, their experience as country of immigration is very recent (if any) and some of them are still net-emigration countries (Dustmann and Frattini, 2013).

gender and age. In the second step, I use the latter as dependent variable in a canonical spatial correlation approach in which I estimate the impact of immigration on natives degree of education.

In the first part of the chapter, I initially assess the direct and indirect effects of immigration on native schooling decisions. I do so by estimating a model that includes as independent variables different age- and education-specific immigrant shares. More precisely, I include the share of immigrants in upper-secondary age, whose associated coefficient corresponds to the direct effect. As previously indicated, this effect is expected to be negative, as immigrants in schooling-age tend to reduce the resources available and therefore induce native to drop-out from school earlier. The indirect effect is instead captured by the share of immigrants aged 19-64 with, respectively, less and equal or more than upper-secondary education. It is an indirect effect as it captures the native education responses based on the perceived complementarity, or rather substitutability, with working-age immigrants of different skills. The sign of the coefficients associated to these two immigrant measures are *ex-ante* ambivalent. Indeed, the presence of low-educated immigrants may, on the one hand, reduce the relative supply of high-skilled workers and therefore increase the return to skills, thus providing incentives for natives to invest in human capital. On the other however, low-skilled immigration may also induce an unskilled-biased technological adjustment of the local economy that can foster the productivity of low-skilled workers and increase their average wages, therefore inducing an increase in the drop-out rate among natives. Also the presence of medium- or high-skilled immigrants may have opposite effects. First, they can create knowledge-spillovers that typically exert a positive impact of both innovation and wages of high-skilled workers and this may induce natives to acquire further education. At the same time, however, it can increase the labor market competition for high-qualified positions that ultimately discourages natives to invest in human capital. Then, I address the high-correlation between the different immigrant measures (and between their specific instruments, too) by estimating a model that includes only one immigrant measure that comprehends all the different age- and education-specific versions, namely the share of immigrant of all education levels aged 15-64. The coefficient associated to it captures the overall effect of immigration on native schooling and its sign depends on the direction of the direct and indirect ones.

In the second part of the analysis, I investigate how the links between immigration and regional sectorial composition affect the schooling decisions of the native population. To be clear, I consider the specialization in sectors that require a low amount of human capital. As indicated above, the hypothesis to test is that the combination between the regional specialization in sectors that require a low amount of human capital (that, for simplicity, I call low-skilled sectors) and the presence of immigrants that, in the countries considered are mostly low-skilled, is expected to exert a negative effect on natives' propensity to invest in human capital. The rationale behind this is that the specialization of a region in low-skilled sectors is typically associated with lower incentives for its population in schooling-age to invest in education as they can easily find a job position. This is further amplified by the inflows of low-skilled immigrants that are often absorbed by local economies through an unskilled-biased change in the productive structure. To this extent, I propose an empirical methodology that considers the dynamics of sector specialization, more that its degree. This, I believe, is a more appropriate way to estimate the native education response to changes in the local employment structure. These are indeed not immediate, but occur over a medium/long time horizon. In other words, I assume that natives, when taking their schooling decisions, do not consider the extent of sector specialization (that is, its degree), but rather its expected evolution. I test this hypothesis by estimating an empirical model in which I interact the immigrant share with an arguably exogenous Bartik-like variable that captures the change in regional specialization in low-skilled sectors.

I conduct the empirical analyses of both parts of the chapter initially considering the whole native population, and then distinguishing by gender (males vs females). This allows to verify whether the native education response to immigration is common to the entire native population, or is rather gender-specific.

Throughout the analysis, I address the endogeneity issue deriving from the non-random immigrant settlement across regions using different instruments. Specifically, on the one hand I build a set of age- and education-specific versions of the shift-share (or enclave-based) instrument, very popular in the migration literature. Then, also to address the recent criticisms that this type of IV has received, I propose an instrument based on the combination between a set of push-factors of the countries of origin and the distance between these and the regions of destination.

In the first part of the analysis, the results indicate that in the period and countries considered, the presence of immigrants is associated with a reduction in the probability of natives to acquire at least upper-secondary education that, in my preferred specifications, varies between -0.6 to -1.3 percentage points. This effect is mostly driven by native females and survives to a battery of robustness tests. Furthermore, when taking into account the local employment structure, the estimates indicate that in regions specialized in low-skilled sectors the native population has lower incentives to accumulate education. Finally, these results appear not to be influenced by some sort of native relocation responses induced by immigrant shocks.

The rest of the chapter is structured as follows. Section 4.2 provides the background of the study. Section 4.3 depicts the empirical models implemented and the variables involved. Section 4.4 presents in detail the data used and the corresponding descriptive statistics. Section 4.5 presents the results, with a battery of robustness tests. Section 4.6 analyzes more deeply the relationship between immigration and regional specialization in low-skilled sectors and how their interaction affects native education. Section 4.7 shows that the results obtained are unlikely to be driven by native internal migration and, finally, section 4.8 provides the concluding remarks.

4.2 Background of the Study

4.2.1 Related Literature

The first contribution of this chapter is to the literature on the effects of immigration on native education. As for the existing studies, Betts (1998) indicates that immigration induces the U.S. native minorities to drop-out from high-school. However, Hunt (2017) argues that the negative result found in Betts (1998) was due to the definition of the outcome of interest that was non-homogeneous over time[6]. Analogously, Borjas (2007) estimates a negative relation between the presence of foreign students enrolled in a particular university and the presence of white native men in the graduate program of the same university. Other examples of a negative impact of immigration on native educational attainments are the

[6]For more details, see footnote 11 in Hunt (2017).

studies of Orrenius and Zavodny (2015), which shows that immigration adversely affects the probability of college-graduated native women to major in a science or engineering field, and Ransom and Winters (2020) who indicate that the Immigration Act of 1990 has caused (i) native black males to go out from STEM majors; (ii) native white males graduated in STEM to go out from STEM occupation; and (iii) native white females STEM graduates to leave the workforce. More recent evidences for the U.S. argue instead in favor of a positive impact of immigration on natives' education. Specifically, McHenry (2015) shows that the presence of low-skilled immigrants induces natives to increase their attendance to high-school and has a positive impact on their average grades and tests scores. In the same spirit, Jackson (2015) shows that an increase in the number of immigrant students in colleges does not harm the enrollment rates of their natives counterparts. Similarly, Hunt (2017) indicates that a unitary percentage increase in the share of immigrants in the population aged 11-64 increases the probability that natives aged 11-17 complete 12 years of schooling (i.e. high-school) by 0.3 percentage points. Along these lines, Figlio et al. (2021) show that immigrants students exert a positive effect on the school performances of natives and this effect is stronger for individuals with more disadvantaged backgrounds.

The evidence from European economies is less abundant. In addition, it has analysed shorter and more recent periods. In fact, to the best of my knowledge, the present chapter is the first study that provides evidence on the native response to immigration for a set of European economies by covering a long time span. More precisely, in the case of Norway, Røed and Schøne (2016) find that in the period 2001 to 2008 an increase in the immigrant share had a negative effect on the enrollment rate of natives in vocational programs specialized in providing building and construction skills. As for Turkey, by exploiting the (arguably) exogenous increase in low-skilled immigrants as consequence of the Syrian refugees crisis of 2012, Tumen (2018) finds that a unitary percentage increase in the refugee-to-population ratio is associated with a 0.4 percentage points increase in the native high-school enrollment rate. This positive response seems to be mainly driven by young males with low-skilled parents, while no impact is instead found on native females. Finally, Brunello et al. (2020) indicate that the recent increase in low-skilled immigration to Italy has triggered a human capital polarization among natives. Specifically, natives with a high-school diploma are induced to further increase their education (this is particularly valid for males in the whole of Italy, and for females in the Northern regions), and, at the same time, high-school drop-out natives tend instead to avoid to enter education or vocational training programs.

The native education response to immigration has also been considered in the structural analysis of Llull (2018*a*). More in detail, the author estimates that natives adjust to immigration shocks by changing their education, labor market participation and occupation. Limiting to the education response, this study reveals that some natives tend to increase their education by around three years, while some others reduce instead their education by also three years[7].

This chapter also contributes to the literature on the links between sector specialization and education. To this extent, Goldin and Katz (1997) for the U.S. find a negative relationship between the share of jobs in the manufacturing sector and the percentage of individuals with secondary education. Along the same lines, Black et al. (2005) find that the coal boom of the 1970s induced an increase in the drop-out rate in the U.S. counties specialized in the production of coal. Similarly, Weber (2014) and Rickman et al. (2017) find a negative

[7]These adjustment mechanisms are crucial as they mitigate the otherwise potentially detrimental effects of immigration on native wages.

association between the booms in the extraction of shale gas and oil occurred in different U.S. states and the high-school and college attainments of their populations.

In the case of Spain, Gutiérrez-Domènech (2011) estimates a positive correlation between the drop-out rate and the employment share in non-technical services and construction. Similarly, Aparicio-Fenoll (2016) finds that the return to education for men experienced a considerable reduction as a result of the housing boom of the period 1997-2008 and, as a consequence, a 10 percent decrease of the skilled to unskilled wage ratio was associated with a 2 percent reduction of the enrollment probability and a 0.2 increase in the drop-out rate among individuals aged 16 to 18. More recently, Diaz-Serrano and Nilsson (2020) identify the presence of a strong (causal) relationship between the employment structure of the local labor markets and the educational attainments of their youth population (i.e. aged 16-24).

However, the main contribution of the present study is to empirically provide long-term evidences on how the combination between sector specialization in low skilled-activities and immigration affects native education in the context of some European countries. To the best of my knowledge, this is the first analysis that attempts to do so.

Finally, I also contribute to the literature on the native (internal) migration responses to immigrant shocks. This is done to verify whether the native education responses are somehow altered by a sort of relocation of the native population. To this extent, the previous studies, which are particularly focused on the U.S. economy, have failed to provide clear-cut results. On the one hand indeed, Borjas et al. (1997) and Borjas (2006) indicate that natives do respond to increased immigration by moving to regions offering better economic opportunities. This mechanism is crucial as it underlies the small impact of immigration on native labor market outcomes estimated by the previous literature (e.g. Card, 1990)[8]. On the contrary, Card and DiNardo (2000) and Peri and Sparber (2011) provide evidences in favor of small (or even absent) native relocation responses to immigrant-driven labor supply shocks. In an attempt to reconcile these conflicting results, Albert and Monras (2020) indicate that these mixed evidences are due to the fact that immigrants disproportionately settle in more expensive areas, where it is easier for natives to relocate. In addition, the shift-share instrument typically used in this literature tends to give more weights to smaller regions or cities that are close to the Mexican border, and this may somehow confound the estimates of the effect of interest.

As for Europe, Hatton and Tani (2005) show that over the period 1982 to 2000 the non-immigrant population in the U.K. tends to respond to immigration shocks by moving out of regions characterized by the larger inflows. As for the Southern economies, Mocetti and Porello (2010) analyze the Italian context and find that, on the one hand, high-skilled natives tend to migrate to regions characterized by higher immigrant density. On the other however, they also identify a displacement effect on their low-skilled counterparts. More in detail, immigration seems to have somehow replaced the well-know South-to-North migration of low-educated Italians. In the case of Spain, Sanchis-Guarner (2017) estimates a positive native relocation responses to immigrant shocks across provinces, that is likely to be due to the presence of skill complementarity between immigrants and natives. Also for Spain, Özgüzel (2021) shows that, during the recent economic crisis, natives were attracted to provinces characterized by the departure of immigrants. The author identified a three to one relationship: for every three immigrants leaving, the native population grew by around one unit.

[8]These results have been recently confirmed by Amior (2021) and Monras (2021).

4.2.2 Education Systems, Immigration Trends, and Specialization in Low-skilled Activities

The education systems of the different countries around the world are classified based on the International Standard Classification of Education (ISCED henceforth). This facilitate cross-countries comparisons as it homogenizes the different education systems under a common structure. This classification is also valid for the European economies under analysis. Overall, ISCED range from early childhood education (ISCED 0) to Master's or equivalent level (ISCED 7). There exist three main models of organization of primary and lower-secondary education, specifically the single structure education, the common core curriculum provision and the differentiated lower-secondary education. The countries under analysis are characterized by the common core curriculum provision. This implies that, after having successfully completed primary education, pupils enter into the lower-secondary education level (ISCED 2) that is characterized by a common curriculum. Afterwards, students move to the upper-secondary education level (ISCED 3), that, contrarily to the lower, is characterized by different curricula depending on the type of school chosen. In any case, in all the countries considered the curricula relative to this level are typically aimed to complete secondary education and to prepare students for tertiary education or to provide them with the minimum skills required to enter the job market. In all the countries considered, pupils enter this level of education at age 15, and lasts until the age of 18 in Greece, Portugal and Spain, while 19 in the case of Ireland.

To this extent, Table 4A2 in the Appendix summarizes the duration of the upper-secondary level, as well as the age at which compulsory education ends in the countries under analysis[9]. Table 4A3 provides instead an overview of the evolution over the period analyzed of the share of natives with upper-secondary education. As is possible to verify, in all the countries analyzed there has been a stable increase in the share of natives with upper secondary education. The most remarkable case is represented by Ireland that experienced an increase in the share of natives with upper-secondary education from around 6% in 1981 to 47% just one decade later. This extremely sharp growth is quite shocking but is consistent with the expansion of the technological sector occurred in those years that encouraged an increase of the educational attainments of the Irish population. Greece and Spain have instead experienced a smother yet consistent increase and, by the year 2011, they had a share of natives with upper-secondary education of around 66 and 52 percent, respectively. Instead, Portugal lags behind the other countries in term of the educational attainment of its native population, and in 2011 just little more than 1 out of 3 natives reported to have upper-secondary education. The growing trend is confirmed in Figure 4A3, which shows the evolution over time of the share of natives with at least upper-secondary education at the regional level. Within each country, the Figure also highlights the presence of a great local variability in the magnitude of interest, and that this spatial dispersion increased over-time.

The phenomenon of immigration into the countries under analysis is somehow recent, especially if compared to countries like the U.S. or Australia, or even to other European countries with a longer tradition of immigration, such as Germany or the U.K. This is a peculiar feature of the countries considered that increases the interest of the present study. In the post-WWII period, some European countries experienced periods of rapid and substantial economic growth, particularly in the decades between the mid-1950's and mid-1970's.

[9]It is important to notice that, in all the countries under analysis, the completion of upper-secondary education goes beyond the compulsory education age.

This entailed large migratory flows from poorer and peripheral countries towards those countries (Dustmann and Frattini, 2013). Alongside, a period of de-colonization characterized Europe, during which many colonies obtained their independency. This has been the case, for example, of Angola and Mozambique, that gained their independency from Portugal between 1974 and 1976 (Carrington and De Lima, 1996; Mäkelä, 2017; Bohnet et al., 2021). This has caused large migratory waves towards the former "mother countries" (De la Rica et al., 2015). This period of mass-migration somehow slowed down in the years following the first oil crisis of 1973, that led to a period of quite severe economic downturn. In the period between 1973 and 1985 the migratory movement mostly consisted in family reunifications (Dustmann and Frattini, 2013). Subsequently, the fall of the Iron Curtain, induced by the fall of the Berlin wall in 1989, together with the outbreak of the Balkan war, triggered large population displacements from Central and Eastern Europe to Western Europe (Angrist and Kugler, 2003; Adjei et al., 2021). Alongside, countries historically less used to receive large migratory inflows - like those under analysis in the present chapter - in few years changed their role from immigrant-sending to immigrant-receiving countries. This has been the case in particular for Spain, that since the beginning of the XXIst century became one of the countries with the higher immigration rates in the world (Amuedo-Dorantes and De La Rica, 2011; González and Ortega, 2011, 2013). This sharp increase was also due to the fact that a large portion of the immigrant population entering Spain originated from former Spanish colonies, where Spanish was the official language[10].

A similar increasing trend also occurred, starting from the 1990s, in Greece (Chassamboulli and Palivos, 2013) and Ireland (Denny and O'Grada, 2013). In the first case, the main determinant has been the collapse of the communist regimes experienced by countries geographically close to Greece (Albania above all). In the second instead, the unprecedented increase in the migratory flows was particularly driven by the economic expansion experienced during that period by the technological sector. To this extent, it is also important to highlight the significant differences in the skill distribution of the immigrant population of the different countries under analysis. More in detail, Greece, Portugal and Spain mostly attract low-skilled immigrants, also because of their (local) industrial structures. On the contrary however, Ireland also hosts highs-skilled workers that were induced to migrate there by the economic boom of the technological sector previously described. To this purpose however, it is always important to remember that, despite their human capital levels, immigrants are often skill-mismatched because of the phenomenon of skill-downgrading that they experience upon arrival in the host country (Dustmann and Preston, 2012; Nieto et al., 2015).

More recently, the EU enlargement waves of 2004 and 2007 induced important migratory flows from the newly entered countries[11] towards other member states, such as Ireland or Spain (Elsner, 2013). Furthermore, despite the outbreak of the global financial crisis that severely hit many European economies, the presence of conflicts in the Arab world (also known as Arab Spring) increased the inflows of refugees and asylum seekers towards the mediterranean and Southern border, particularly Italy, Greece and Spain (Labanca, 2020; Hanson and McIntosh, 2016).

As for the data used, Table 4A1 of the Appendix reports the evolution over time of the

[10]Indeed, the knowledge of the language spoken in the destination countries plays a crucial role in shaping international migratory flows (Adserà and Pytliková, 2015; Lewis, 2013).

[11]In 2004 the countries that joined the EU were Czech Republic, Estonia, Cyprus, Latvia, Lithuania, Hungary, Malta, Poland, Slovakia and Slovenia, while in 2007 Romania and Bulgaria. For more details, see https://ec.europa.eu/neighbourhood-enlargement/policy/from-6-to-27-members_en

immigrant population in the countries considered. Some aspects need to be highlighted. First, in all countries the immigrant population increased between 1981 and 2011. Second, the countries that experienced the largest relative inflows were Ireland, that experienced an increase of their immigrant share from 7% in 1981 to around 22% just three decades later. Similarly, Greece and Spain moved from a share of immigrants that was around 1% and 0.6% of the population in 1981, to about 10% and 12% in 2011, respectively[12]. A considerably lower increase occurred instead in Portugal, from around 1% in 1981 to little less than 5% three decades later. These figures (particularly those for Ireland, Greece and Spain) are in line with countries with a longer tradition of immigration, like the U.S., Canada or Australia (for more details, see Figure 1 in Peri, 2016). Finally, Figure 4A2 shows that the increasing trend in the evolution over time of the immigrant population also occurs at the local level, highlighting fairly large regional disparities, particularly in the 2010s

As already indicated, the employment structure of a local economy plays a crucial role in the schooling determinants of its population. In regions whose production system is skewed towards more low-qualified labor-intensive type of sectors there are lower incentives to invest in human capital than in regions characterized by the presence of firms that are more capital-intensive or that make a prevalent use of high-skilled labor. This is particularly valid in the countries under analysis, whose sectoral composition varies strongly between regions, and that are characterized by an uneven distribution of economic activities. In addition, given the counter-cyclicality of education, this implies that an economic downturn that affects a particular sector may induce the local population to invest in human capital in regions specialized in that specific sector.

In the empirical analysis of the present article, I consider the specialization of a region in low-qualified jobs of the low-skilled sectors. I base this on the share of low-educated workers (that is, with less than upper-secondary education) in low-skilled sectors. These are defined based on Eurostat's classification. Specifically. I consider as low-skilled sectors medium-low- and low-technology manufacturing, less-knowledge-intensive services, agriculture, construction and mining[13]. Figure 4A4 reports the evolution over time of the share of low-educated workers in low-skilled sectors. As is possible to verify, in all the countries considered the number of low-educated workers in low-skilled sectors represents an important portion of the working-age population, although there are important within-country differences. Some regions are indeed characterized by an amount of low-educated workers in low-skilled sectors that represents around 50 or 60% of the working-age population, while others have values even lower than 20%. Moreover, the Figure also shows a stable decline over time in the magnitude of interest, particularly in the last decade. This decreasing trend is consistent with the recent development of new technologies that substitute routine-based activities and complement high-qualified jobs (Autor et al., 2003; Basso et al., 2020)[14].

4.3 Empirical Strategy

This section sketches the empirical strategy that I use to estimate the impact of immigration on natives education. In line with Hunt (2017), after presenting the outcome of interest

[12]Interestingly, these two countries experienced most of the increase in their immigrant population between 2000 and 2010.

[13]For more details, see https://ec.europa.eu/eurostat/cache/metadata/Annexes/htec_esms_an3.pdf.

[14]It must be noted however, that in the empirical analysis I consider the one-census lagged values of the variable and, therefore, I do not directly capture its most recent dynamics.

and the way in which I estimate it, I distinguish two channels through which immigration may affect natives' investment in human capital. Specifically, I initially estimate the direct and indirect effects of immigration on native education. The first captures the effect of immigrants students on the school performance of their native counterparts, and is assumed to be negative. The second captures instead the impact that working-age immigrants with different education levels exert on the tendency of natives to acquire human capital. The sign of this second effect is not clear *ex-ante* and needs to be estimated empirically. Then, I estimate the overall effect, that is, the impact that immigrants aged 15-64 exert on natives education. This effect is the combination of the direct and indirect ones and its sign depend on the sign of these two effects and, in case of opposite effects, on which of the two prevails. This strategy is complemented by the analysis in section 4.6 about the complementary effect of immigration and local specialization in low-skilled activities on natives education.

4.3.1 Outcome of Interest

The main objective of this chapter is to estimate the impact of immigration on the probability of natives to invest in human capital. However, instead of using the "simple" shares of natives in upper-secondary education, following Hunt (2017) I construct a composition-adjusted regional schooling variable. In other words, I perform a two-steps procedure that allows me to obtain a measure of regional schooling after controlling for individual observable characteristics. Given that natives tend to sort by skills (Albert and Monras, 2020; Combes and Gobillon, 2015; Hean et al., 2020), this procedure is crucial as it allows to control for changes over time and across local labor markets of the composition of the native population. To this extent, in the first-step for each native individual i, residing in region r at time t, I estimate the following linear probability model:

$$P(H_{irt} \geq up\_secondary) = \alpha_0 + \alpha_1 Z'_{irt} + \sum_r \sum_t \kappa_{rt}(\varphi_r \times \psi_t) + \eta_{irt} \tag{4.1}$$

where H_{irt} is a dummy variable that takes the value of 1 if the individual i has upper-secondary education or more, and 0 otherwise (that is, less than upper-secondary education). Therefore, it captures all those individuals with at least upper-secondary education completed. Z'_{irt} is a set of individual characteristics such as gender and age[15]. These variables are included because the probability of an individual to attain (at least) upper-secondary education varies depending on his/her gender and age. Finally the term $\varphi_r \times \psi_t$ represents an interaction between region and time dummies (φ_r and ψ_t, respectively). The main coefficient of interest is κ_{rt}, which represents an indicator of regional schooling net of observable micro-level characteristics. As in e.g. Hunt (2017), I perform the regression using as weights the census weights provided by IPUMS[16]. I compute the regional schooling variable for the whole native population and for men and women separately. In these last cases, I adjust the regional schooling variable just for age. As is possible to verify from Figure 4A1, the variable so-constructed is highly correlated with the un-adjusted share of native with upper-secondary education.

[15]Some studies (e.g. McHenry, 2015) also control for parents' education as is often correlated with children's investment in human capital. Unfortunately however, such information is not available for all the countries considered and therefore a measure of parents' education is not included in the vector of individual controls of equation (4.1).

[16]The unweighted regressions yield similar results.

4.3.2 Direct and Indirect Effects

In the second step, I use the previously estimated coefficients $\hat{\kappa}_{rt}$ to compute the change between two censuses and use it as dependent variable in a spatial correlation model that identifies the impact of immigration on the educational attainments of the natives[17]. Initially, I empirically assess the direct and indirect effect of immigration, by estimating the following model:

$$\Delta\hat{\kappa}_{rt} = \beta_0 + \beta_1 \Delta m^{us\ age}_{r,t-10} + \beta_2 \Delta m^{H<us}_{r,t-10} + \beta_3 \Delta m^{H\geq us}_{r,t-10} + \beta_4 \Delta X_{r,t-10} + \psi_t + \varepsilon_{rt} \tag{4.2}$$

where $m^{us\ age}_{r,t-10}$ indicates the share of immigrants in upper-secondary age over the total population also in upper-secondary age[18], $m^{H<us}_{r,t-10}$ is the share of immigrants aged 19-64 with less than upper-secondary education, while $m^{H\geq us}_{r,t-10}$ represents instead the share of immigrants aged 19-64[19] with exactly equal to, or more than upper-secondary education. Both these last regressors are computed as share of the total population aged 19-64. $X_{r,t-10}$ is a vector of region-level control variables, that are aimed at capturing, at least partially, the effects of some local characteristics. These are the youth unemployment rate, defined as the unemployment rate for individuals aged 19-24[20], and the prime-age unemployment rate, defined as the unemployment rate for individuals aged 25-54. The rationale behind their inclusion is that the investment in education is countercyclical, that is, it decreases in periods of economic prosperity and it increases during downturns (Card and Lemieux, 2001; Clark, 2011; Di Pietro, 2006; Petrongolo and San Segundo, 2002; Reiling and Strøm, 2015; Diaz-Serrano and Nilsson, 2020). This implies that the higher is the unemployment rate of workers aged 19 and older (that is, that have completed upper-secondary education), the more individuals still at school are expected to further invest in human capital. Then, I include the native cohort size, defined as the share of natives in upper-secondary age over the total working-age population, as it allows me to control for changes in the composition of the native population that may affect the regional schooling variable. It is important to notice that all the right-hand-side variables are lagged one census, so as to capture the time when natives are completing upper-secondary education, that is when they decide whether to further invest in education or not. Finally, ψ_t are census-year fixed-effects, and η_{rt} is an *i.i.d.* distributed error term, with zero mean and variance σ^2_η. As is common practice in econometrics, the second-step regression is weighted by the inverse of the squared standard errors of $\hat{\kappa}_{rt}$ estimated in the first-step[21].

In this setting, the coefficients of interest are β_1, β_2 and β_3. The first is the estimate of the direct effect, while the other two of the indirect one. More in detail, β_1 is expected to be negative as the presence of foreign-born students tends to reduce the quality of schooling (particularly in the case in which they have a poor knowledge of the language spoken in the destination country). As indicated by Hunt (2017), this coefficient may be interpreted as an elasticity of native educational attainment with respect to immigration. The sign of β_2, is instead more ambiguous. On the one hand, it might be positive, particularly in

[17] Throughout the chapter, all the empirical analyses are conducted at the year-region cell level, procedure that ensures to identify the total effect of immigration that, as indicated by Dustmann et al. (2016), is the most policy-relevant effect of immigration.

[18] i.e. 15-18 for Greece, Portugal and Spain, and 15-19 in the case of Ireland.

[19] 20-64 in the case of Ireland.

[20] 20-24 for Ireland.

[21] This procedure also allows me to correct for heteroskedasticity (Solon et al., 2015). However, I show that the results are robust to the modification of the weighting scheme and even to their removal. More details on the weights used are provided in the notes of the Tables reporting the estimated effects.

the case in which the presence of low-skilled immigrants alters the skills distribution of the destination regions and increase the return to skill, therefore providing more incentives for native to acquire human capital (as shown by Hunt, 2017). On the other hand however, in reaction to the immigration-induced increase in low-skilled labor, firms may change their capital-labor ratios and make use of technologies that enhance the productivity of the more abundant factor type (i.e. low-skilled workers), as shown in Clemens et al. (2018), Lewis (2011), Lafortune et al. (2019), and González and Ortega (2011). As a consequence, natives have less incentives to invest in human capital. The sign of β_3 is assumed to capture the effect that medium- or high-skilled immigrants exert on native investment in human capital. Again, the direction of this effect may be ambivalent. Indeed, the presence of medium- or high-skilled workers is typically associated with productivity gains at the local level (Moretti, 2004; Iranzo and Peri, 2009). This may create positive externalities (also known as knowledge-spillovers) that boost innovation (Hunt and Gauthier-Loiselle, 2010; Kerr and Lincoln, 2010; Peri, 2005) and create wage gains for high-educated natives (Peri et al., 2015). The combination of these factors, may ultimately induce natives to invest in human capital. On the other hand however, medium- or high-skilled immigrants may somehow displace their native counterparts by increasing labor market competition for high-skilled job positions and thus discourage the latter to invest in education (Borjas, 2009; Borjas and Doran, 2012, 2015).

4.3.3 Overall Effect

A potential limitation of the model sketched in equation (4.2) is that it has the three immigration endogenous regressors that are highly correlated between them. This could represent a problem, particularly when estimating the model by IV, as the first-stage F-statistic is likely to be very low in presence of highly correlated multiple instruments (Angrist and Pischke, 2009). A way to overcome the high correlation between the different immigrant measures is to estimate what I call the "overall" effect. This allows to capture the total effect of immigration on native education by using only one endogenous regressor. In other words, as alternative to (4.2), I estimate the following model:

$$\Delta\hat{\kappa}_{rt} = \gamma_0 + \gamma_1 \Delta m^{15-64}_{r,t-10} + \gamma_2 \Delta X_{r,t-10} + \psi_t + \nu_{rt} \tag{4.3}$$

where $m^{15-64}_{r,t-10}$ indicates the lagged share of immigrants aged 15-64 over the total working age population (i.e. also aged 15-64). The other variables are defined as before. The coefficient of interest is γ_1 which is an estimate of the overall effect of immigrants on native education and captures the combination of the direct and indirect effects. The sign of γ_1 ultimately depends on (i) the direction of both effects, and (ii) in case of opposite effects, which of the two prevails.

4.3.4 Identification Issues

The regressions expressed in equations (4.2) and (4.3) are likely to suffer from endogeneity due to different reasons. The first is the potential omission of relevant variables that can confound the estimates of the effects of interest. Those that are time-invariant are filtered-out by means of the specification in first-differences. Those that instead vary over time are captured by the local factors in $X_{r,t-10}$. In addition, differences in the evolution of education are controlled for by the inclusion of country-specific trends (country dummies in the first-differences specifications). The second is a reverse causality issue. Indeed, the decision of immigrants to settle in a particular region may be driven by the same (economic) factors that could also influence natives educational attainments. To clarify, an unexpected increase in the demand for low-qualified workers due, for instance, to the expansion of a sector that make prevalent use of this type of labor[22], may (i) attract new immigrants; and (ii) induce natives to drop-out from school and start working. On the other hand, a sudden economic downturn may push immigrants towards other regions, looking for better job opportunities, and induce natives to invest in education in order to acquire a qualification that could somehow reduce the likelihood of unemployment, once in the labor market. Depending on the sign of the correlation between the dependent variable and the unobservable determinants of immigrants settlement, I could over- or under-estimate the effect of interest. OLS estimates may further suffer from attenuation bias due to measurement error in the estimation of the different age- and education-specific immigrant shares (Aydemir and Borjas, 2011).

All these reasons may work against the identification of the causal effect of immigration on native education when estimated by means of the OLS estimator.

4.3.5 The Enclave-based Instrument

To overcome these issues, I estimate equations (4.2) and (4.3) with an instrumental variable approach using the canonical enclave-based shift-share variable, pioneered in the migration literature by Altonji and Card (1991) and Card (2001) and widely used henceforth[23]. This instrument is based on the fact that immigrants tend to settle in areas characterized by previous immigrants coming from the same country of origin (Bartel, 1989; Munshi, 2003)[24] and takes the following form:

$$\Delta\widehat{M}_{r,t} = \sum_{o} \left(\frac{M_{o,r,t_0}}{M_{o,t_0}}\right) \cdot \Delta M_{o,t}$$

where the first element is the share component and is defined as the stock of immigrants coming from country o, residing in region r in a base year (i.e. M_{o,r,t_0}), divided by the stock of immigrants coming from the same country o and in the same year t_0 at the national level (i.e. M_{o,t_0}). The second part of the equation, $\Delta M_{o,t}$, is the shift component and is defined as the decadal change between the stock of immigrants coming from country o at the national level - that is, $\Delta M_{o,t} = (M_{o,t} - M_{o,t-10})$. Since the endogenous variable is defined as the share of immigrants over the total working-age population, I also divide the predicted immigrant change previously defined (i.e. $\Delta\widehat{M}_{r,t}$) by the total working-age population of the region[25]. The resulting instrument is therefore defined as:

$$\Delta\widehat{m}_{r,t} = \frac{\Delta\widehat{M}_{r,t}}{L_{r,t-10}}$$

where $L_{r,t-10}$ is the regional working-age population at the start of the period over which

[22] As it happened during the Spanish property bubble.

[23] Recent examples in the context of the impact of immigration on native education are Hunt (2017) and Brunello et al. (2020).

[24] This is what is called "network" or "enclave" effect.

[25] As in Hunt (2017), I use the regional population at the start of the period over which the change is computed. For instance, for $\Delta M_{o,2000} = (M_{o,2000} - M_{o,1990})$ I use the working-age population of 1990.

$\Delta\widehat{M}_{r,t}$ is computed. I construct a specific instrument for each age- and education-specific endogenous regressor. However, the share component is common to all age- and education-specific instruments since it is more likely to be exogenous with respect to the economic conditions of the local labor markets of destination (Hunt, 2017). This is likely to represent a problem when estimating equation (4.2) by means of the IV estimator.

The first-stage regressions are reported in Table 4A4. Specifically, Panel A reports the first-stage relative to equation (4.2). As is possible to see, all the age- and education-specific instruments display positive and statistically significant first-stage correlations[26]. However, the Sanderson and Windmeijer (2016) F-statistics are always fairly low. This is likely due to the high correlation between the different age- and education-specific instruments. As indicated in the results section of this chapter, this implies that the estimates used by means of these IV estimators should be taken with caution. Things are considerably better in Panel B of Table 4A4, where I report the the first-stage relative to equation (4.3). In this case, the instruments appear not to be weak, as the first-stage correlations are always positive and statistically significant, and the F-statistics always fairly high. Moreover, all the first-stages are robust to the inclusion of the country-specific trends[27].

In Table 4A5, I instead report the results of the reduced form where I regress the composition-adjusted regional schooling variable on the different age- and education-specific enclave-based instruments. This allows me to verify the extent of correlation between the dependent variable and the different instruments. Interestingly, in columns (1) and (2), all instruments show a negative and significant correlation, that is in line with the initial hypothesis. Things change slightly in column (3). Specifically, the coefficient associated to the enclave-based instrument for immigrants in upper-secondary age is negative and significant. This is consistent with the presumably negative effect of immigrants in schooling-age on their native counterparts. In addition, the sign of the two instruments for immigrants aged 19-64 with, respectively, less and equal or more than upper-secondary education is in line with the initial hypotheses (that is, negative the former and positive the latter). However, in both cases point estimates are not statistically significant

Instrument Validity

Despite being widely used[28], the enclave-based shift-share instrument has recently received numerous criticisms (Borjas et al., 2012; Adão et al., 2019; Jaeger et al., 2019; Borusyak et al., 2020; Goldsmith-Pinkham et al., 2020). The main drawback of this type of variable is that, if the determinants behind the immigrants' location decisions in the initial year are persistent over time (as it is often the case), then they may influence the immigrants' settlement patterns in the subsequent years, too. This is somehow addressed by using the country-wide growth of the stock of immigrants by countries of origin. However, the same economic factors that drive the inflows of immigrants at the national level may also affect the local labor market outcomes[29], thus producing a bias in the estimates of the effect of

[26] The only exception is represented by the shift-share variable for people aged 19-64 with less than upper-secondary education whose associated coefficient is only marginally significant. However, when estimating the first-stage using as independent variable this specific instrument alone, the correlation increases to 0.304, with s.e.=0.087 (significant at 1% level), F-statistic=12.6 and R-squared=0.615.

[27] For the sake of space, I do not report these results, but they are available upon request.

[28] See Appendix Table 1 in Jaeger et al. (2019), for a detailed review of the existing literature using this type of identification strategy.

[29] A period of economic expansion of a country driven by a specific sector, may (i) attract new immigrants, and (ii) produce an increase in wages and employment levels in the regions specialized in that sector.

interest.

The validity of the enclave-based shift-share variable as described previously therefore depends on its two components (i.e. the "shift" and the "share") being orthogonal to the local labor markets conditions that may affect the outcome variable. As for the share component, the exclusion restriction requires it to be uncorrelated with the local economic conditions of the period under analysis[30]. This is because it is assumed that the only factors that drive immigrants' location decisions are the network effect or their idiosyncratic tastes. I test this in Table 4A6 where I regress the share component of the instrument on a set of economic and education indicators of the local labor markets under analysis. The results show the complete absence of correlation.

With respect to the shift component, Jaeger et al. (2019) indicate that the instrument might not be fully exogenous if the location patterns and the countries of origin distribution of the immigrant population are persistent over time. In the case of Europe, however, these concerns are likely to be less strong than in the U.S. (Jaeger et al., 2019). This is particularly valid in the set of countries considered that suddenly change their role from immigrant-sending to immigrant receiving countries. In addition, the E.U. enlargement episode of 2004 and 2007 considerably reduced the cost of migrating and resulted in an unexpected huge increase in the immigrants inflows from CEE, notably towards Spain and Ireland (De la Rica et al., 2015; Elsner, 2013). In addition, the autocorrelation coefficient of the stock of immigrants by countries of origin is around 0.56 for both working-age individuals and for the whole immigrant population. This is considerably lower than the values of 0.9-0.99 estimated by Jaeger et al. (2019) for the U.S., and also than the value of 0.67 estimated by Özgüzel (2021) for Spain. Finally, the first-stage regressions typically display fairly high F-statistics (see Table 4A4). This implies that, even in presence of some correlation with the error term, the resulting bias in the estimates is expected to be small. All these aspects are reassuring in terms of the validity of the instruments.

4.3.6 The Push-factors and Distance-based Instrument

As an alternative to address the concerns relative to the shift-share instruments based on the settlement patterns in a base year, in this section I propose an instrument that is instead based on two arguably exogenous determinants of immigrant inflows, namely distance and push-factors of the countries of origin[31]. This variable builds on the "geographic shift-share" instrument used by Sparber and Zavodny (2020)[32]. Specifically, I obtain the share component of the instrument by estimating the following regression:

$$\frac{M_{r,o,t_0}}{M_{r,t_0}} = \varrho_0 + \varrho_1 \log(dist_{o,r}) + \varphi_r + \chi_o + \zeta_{r,o,t_0} \tag{4.4}$$

where the dependent variable is the share of immigrants from each country of origin o that reside in region r in year t_0, $dist_{o,r}$ is the distance in kilometers between the capital city of each country of origin o and the capital city of each region r, while φ_r and χ_o are region and country of origin fixed-effects, respectively.

I then compute the predicted values from this regression that represent the expected share of immigrants in each region from a particular country in the base year driven by the distance

[30] The share component of this type of variable plays an important role as the identification mostly comes from it (Goldsmith-Pinkham et al., 2020)

[31] This instrument is used only in th estimation of the overall effect.

[32] A push-factors-based instrument is also used in Boustan (2010) and Heepyung (2019).

between the region of residence and the country of origin of the immigrant population. This component provides the regional variation required to instrument the actual regional immigrant shares and is arguably more exogenous than the historical measure of immigrant shares by origin, as distance is not correlated with economic factors (and, moreover, is fixed-over time)[33]. The predicted share component in the base year is then augmented by the national growth rate of the immigrant population by countries of origin in the subsequent years obtained from a regression of this type:

$$\log(M_{o,t}) = \varrho_0 + \varrho_1 \Pi'_{o,t} + \chi_o + \psi_t + \xi_{o,t} \tag{4.5}$$

where $\Pi'_{o,t}$ is a vector of push-factors for the countries of origin and year (for every year t, I compute it as the average of the 10 previous years). I use different types of push factors, that capture different aspects of the origin countries that could induce migration and that are typically used in gravity models that predict immigrant stocks and flows. These are (i) GDP (constant in 2010 USD), capturing the economic conditions; (ii) total population and the share of the total population which is rural, capturing the demographic dynamics; (iii) school enrollment for primary and secondary education (gross, gender parity index), capturing immigrants' education distribution; (iv) infant mortality rate (per 1000 births), capturing health conditions; and (v) military expenditure as share of the GDP, used as proxy for the tendency of each country of origin to participate in wars or armed conflicts.

Then, I obtain the predicted national immigrant stocks by taking the exponentials of the predicted values from equation (4.5), and obtain the predicted immigrant stock in each region-year cell ($\widehat{M}^{push}_{r,t}$) by multiplying the predicted national immigrant stocks with the predicted regional shares by origin. The instrument is finally defined as:

$$\Delta \widehat{\mathcal{M}}_{r,t} = \frac{\Delta \widehat{M}^{push}_{r,t}}{L_{r,t-10}}$$

The first-stage and reduced form regressions are reported in Tables 4A4 and 4A5. Specifically, the last column of Panel B of Table 4A4 shows that the instrument so constructed is a good predictor of the actual immigrant share in working-age, as the first-stage correlation is positive and statistically significant, and the F-statistic is around 31.2. If I further include country trends the correlation remains positive and significant (around 0.106) and the F-statistic increases to 47.5. The last column of Table 4A5 shows instead the reduced-form regression where I estimate the correlation between the composition-adjusted regional schooling variable (i.e. the dependent variable) and the instrument. The point estimate is around -0.07, although only marginally significant. However, if I include country trends, the reduced form correlation increases to -0.120 and becomes significant at 1% level. Taken together these results identify the presence of a negative correlation between the dependent variable and the immigrant share that is confirmed in the rest of the analysis.

[33] In the migration literature, distance-based instrument is used, for instance, in Mocetti and Porello (2010) and Peri (2012).

Instrument Validity

In this section, I investigate the validity of the push-factor and distance-based instrument previously defined. Specifically, in Table 4A6 I show the correlation between the share-component of the instrument and some economic and education indicators of the local labor markets under analysis. As is possible to verify, there is an almost absence of correlation between the magnitudes under analysis. The only exception is represented by the share of natives with upper secondary education that, in the specification without country trends of column (3), displays an associated coefficient significant at 5%, although negative. However, once I control for country trends, the coefficient become not statistically significant. As for the validity of the shift component, the autocorrelation coefficient of the stock of immigrants by countries of origin is around 0.69, again considerably lower than the one estimated for the U.S. by Jaeger et al. (2019) and in line with the one for Spain estimated by Özgüzel (2021).

The presence of two available instruments (i.e. the enclave-based shift-share and the push-factor and distance-based) allows me to perform the test of overidentifying restrictions, in the case of equation (4.3). The joint null hypothesis is that both instruments are valid, that is, they are orthogonal with respect to the error term. The corresponding results are reported in Table 4A7. Reassuringly, for all the demographic groups considered, I cannot reject the null hypothesis of the instruments being valid, as the Hansen's J-statistics are typically low and have fairly high p-values. It is worthwhile to highlight that the fact that the rationales behind the construction of the two instruments are different[34] makes the results of the overidentification test more credible.

4.4 Data and Descriptive Analysis

The main data used for the analysis are drawn from the IPUMS samples of the population censuses for all the countries considered over the period 1981-2011 (Minnesota Population Center, 2019). More in detail, the dependent variable is computed using the data for the period 1991 to 2011, while all the independent variables are lagged one census and therefore are computed using the data relative to the period 1981 to 2001[35]. As indicated in section 4.3, the dependent variable is computed using a two-step procedure. I define the outcome of interest as the completion of upper-secondary education (ISCED 3 level). In this way, the dependent variable is based on post-compulsory education.

The regressors of interest in equations (4.2) and (4.3) are based on the share of immigrants over the total population, $m_{r,t} = M_{r,t}/L_{r,t}$, where $M_{r,t}$ is the stock of immigrants in each region and year and $L_{r,t}$ is the total population. In both cases, I only considered working-age individuals (i.e. people aged 15 to 64). I distinguish between immigrants and natives based on their citizenship. Specifically, I define as immigrants those individuals that are non-citizen, or without any citizenship (i.e. "stateless"). I choose to distinguish between immigrants and natives based on their citizenship status because of different reasons. First of all, to be consistent with a strand of the previous literature on immigration (e.g. Amuedo-Dorantes and De La Rica, 2011; Borjas, 2003; Basso and Peri, 2015; Cortés and Tessada, 2011; Edo, 2015, just to cite few). Besides, the Southern European countries, for whom

[34]Indeed, one is based on the network-effect resulting from the presence of other immigrants coming from the same origin, while the other is based on the distance and push-factor of the countries of origin.

[35]In the case of Ireland, the census of the beginning of the 2000s was carried out in 2002 instead of 2001.

the phenomenon of immigration is relatively recent, during the period under analysis had a somewhat complex and rigid naturalization process, implying that that most immigrants tend to keep their own citizenship for a fairly long period of time (OECD, 2011). In addition, the possession of the host country's citizenship is often seen as a "positive signal" from the potential employer and as an indicator of better predisposition to integration. Indeed, naturalized immigrants tend to have a better knowledge of the language spoken in the destination country compared to their non-naturalized counterparts. This thus implies that naturalized immigrants are closer in terms of labor markets characteristics and school performances to natives, and therefore their inclusion in the analysis may somehow confound the estimates[36].

Finally, the data on the push factors for the countries of origins that I used to compute the instrument sketched in the previous section are drawn from the World Development Indicators provided by the World Bank[37].

In addition to that commented in the previous sections, here I focus on the descriptive of the overall sample of local labour markets included in the analysis. Specifically, Table 4A8 presents the descriptive statistics of the main variables used in the analysis. It is worth to highlight some aspects. First, in the period and countries under analysis there has been an overall increase in regional schooling of around 15%. The increase has been stronger for native females (around 17%) than for their male counterparts (increase of around 13%). Second, the average region has experienced an increase in the working-age immigrant population by around 2.2%. Moreover, an increasing pattern is also found when considering immigrants in upper-secondary age (+1.4%), immigrants aged 19-64 with less than upper-secondary education (+0.9%) and with exactly equal to, or more than upper-secondary education (+1%). Third, the period considered has been characterized by relatively good economic conditions[38]. Indeed, the unemployment rate for people aged 19-24 has only increased by 0.8%, while slightly higher was the increase for people aged 25-54 (+2%). However the large regional disparities that characterize the countries considered are particularly evident in this last case. Indeed, in the case of the youth population, there are region-years for which the unemployment rate has experienced a decrease of 15% together with others in which it has increased by around 17%. Similarly, the prime-age unemployment rate has decreased in some regions-year by roughly 6%, while in others it has grown by slightly less than 12%. Moreover, the native cohort size has experienced an average decrease by about 1.8%. Finally, the regions under analysis appear to experience a switch in their employment structure from low- to high-skilled sectors. Indeed, in the period under analysis the share of low-educated workers in low-skilled sectors experience, on average, a decrease of around 6%. This declining trend is particularly accentuated in some region-year cells that experience a decrease of around 22%.

4.5 Results

In the this section, I will first provide the estimates of the direct and indirect effects and of the overall effect. Then, I present a set of robustness checks that test the validity of the results obtained.

[36] A similar circumstance is, for example, emphasized by Edo (2015) in the case of immigrants in France.

[37] See https://databank.worldbank.org/source/world-development-indicators.

[38] At this purpose, it is important to notice that all the right-hand-side variables are computed over the period 1981-2001, and, therefore, I do not capture the Global Financial Crisis.

4.5.1 The Direct and Indirect Effects

Table 4.1 presents the results of the direct and indirect effects as sketched in equation (4.2). Panel A is relative to the whole population, while Panels B and C to men and women, respectively.

Table 4.1: Direct and indirect effects.

	OLS			2SLS	
	(1)	(2)	(3)	(4)	(5)
Panel A: Whole population					
$\Delta m^{us\ age}_{r,t-10}$	0.162	-0.115	-0.467**	-1.408	-0.999
	(0.204)	(0.248)	(0.207)	(1.458)	(1.170)
S.W. F-statistic	-	-	-	3.9	3.7
$\Delta m^{H<us}_{r,t-10}$	-0.178	0.073	1.030***	-0.466	-0.047
	(0.458)	(0.393)	(0.340)	(2.386)	(3.326)
S.W. F-statistic	-	-	-	3.9	2.6
$\Delta m^{H\geq us}_{r,t-10}$	-0.202	-0.587*	-0.466	0.820	0.339
	(0.325)	(0.345)	(0.350)	(1.490)	(1.038)
S.W. F-statistic	-	-	-	4	9.9
Controls	NO	YES	YES	YES	YES
Country trends	NO	NO	YES	NO	YES
R^2	0.375	0.515	0.698	0.362	0.642
Panel B: Men					
$\Delta m^{us\ age}_{r,t-10}$	0.156	-0.054	-0.505**	-1.555	-1.383
	(0.220)	(0.281)	(0.218)	(1.647)	(1.299)
S.W. F-statistic	-	-	-	4.3	3.9
$\Delta m^{H<us}_{r,t-10}$	-0.301	-0.089	1.226***	-1.011	0.616
	(0.521)	(0.456)	(0.394)	(2.687)	(3.552)
S.W. F-statistic	-	-	-	4.3	2.7
$\Delta m^{H\geq us}_{r,t-10}$	0.353	-0.100	-0.335	1.728	0.962
	(0.390)	(0.412)	(0.386)	(1.679)	(1.148)
S.W. F-statistic	-	-	-	4.5	11.7
Controls	NO	YES	YES	YES	YES
Country trends	NO	NO	YES	No	YES
R^2	0.206	0.366	0.605	0.095	0.537
Panel C: Women					
$\Delta m^{us\ age}_{r,t-10}$	0.160	-0.168	-0.404*	-1.127	-0.462
	(0.223)	(0.255)	(0.241)	(1.440)	(1.297)
S.W. F-statistic	-	-	-	3.4	3.5
$\Delta m^{H<us}_{r,t-10}$	-0.049	0.241	0.827**	-0.113	-1.288
	(0.427)	(0.389)	(0.361)	(2.292)	(3.732)
S.W. F-statistic	-	-	-	3.5	2.6
$\Delta m^{H\geq us}_{r,t-10}$	-0.803**	-1.105***	-0.615	-0.190	-0.277
	(0.317)	(0.347)	(0.382)	(1.480)	(1.096)
S.W. F-statistic	-	-	-	3.4	7.5
Controls	NO	YES	YES	YES	YES
Country trends	NO	NO	YES	No	YES
R^2	0.493	0.589	0.711	0.529	0.645

***Note**: The table reports the estimates of the direct and indirect effects, as indicated in equation (3.2). All the RHS variables are lagged one census. All regressions include time fixed-effects and are weighted by $1/(1/w_t + 1/w_{t-10})$. In all specifications the number of observations is 258. Standard errors, in parentheses, are clustered at the province level.*

*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

More in detail, columns (1) to (3) report the OLS estimates, while columns (4) and (5) report the 2SLS ones (when using the age- and education-specific enclave-based instrument). The first column is relative to the baseline specification with only the different age-and education-specific immigrant shares but without controls and country trends. In general, point estimates are not statistically significant, with the only exception of the share of immigrants aged 19-64 with at least upper-secondary education that has a negative and significant (at 5% level) coefficient in the case of women. Column (2) reports the estimates of the direct and indirect effect from the specification that includes the controls (youth and prime-age unemployment rates, and native cohort size). Again, the coefficients associated to the share of immigrants in upper-secondary age and to the share of immigrants aged 19-64 with less than upper-secondary education are not statistically significant. On the other hand, the coefficient on the share of immigrants aged 19-64 with equal or more than upper-secondary education is negative and significant for the whole population and for native women. These estimates are in line with the idea that the presence of high-educated immigrant workers has a displacement effect on natives and induce them to drop-out from school (as shown in Borjas, 2009; Borjas and Doran, 2012, 2015). When adding country-specific trends (column 3), however, the results change considerably and are in line with those reported in Hunt (2017). Specifically, the estimate of the direct effect (proxied by the coefficient on the share of immigrants with upper-secondary age) is negative and significant. In line with the idea that immigrant students lower the quality of education, this implies that the presence of immigrants in upper-secondary age is associated with a reduction in native probability to invest in human capital by around -0.5 percent. On the other hand, the coefficient on the share of immigrants aged 19-64 with less than upper-secondary education is also statistically significant but positive in sign. Point estimates imply that, overall, the presence of low-skilled immigrants is associated with an increase in native education of about 1 percent. Interestingly, the gender distinction indicates that the education responses are stronger for men (1.2%) than for women (0.8%).

However, the non-random settlement of immigrants is likely to bias the OLS estimates. To overcome this problem, in columns (4) and (5) I present the 2SLS results obtained by using the age- and education-specific versions of the enclave-based instrument, in the specifications with controls and both without and with the country-specific trends. In this case, however, the estimates of the effects of interest are never significant. The Sanderson and Windmeijer (2016) F-statistics of excluded instruments are usually very low in most cases, which points to the weakness of the instruments[39]. It should be noted that I aim to estimate the effects of interest using three endogenous regressors that are highly correlated between them and whose specific instruments have the share component in common[40]. To this purpose, however, Angrist and Pischke (2009) indicate that to estimate a 2SLS model with more than one endogenous regressor is often problematic, and therefore the 2SLS estimates of this section should be taken with caution.

4.5.2 The Overall Effect

Due to the problems with the identification of both the direct and indirect effects, I therefore provide evidence on the overall effect, as presented in equation (4.3). Table 4.2

[39] I report the Sanderson and Windmeijer (2016) F-statistic as, in case of multiple instruments, is more reliable than the Kleibergen and Paap (2006) F-statistic.

[40] However, to build different age- and education-specific share components would be problematic for two reasons. The first is that the use of the entire immigrant population (that is, irrespective on age or education), gives greater prominence on the network effect, rather than on the local economic conditions that might instead attract immigrants of different age groups or education levels. The second is that this procedure allows me to make use of all the available regions because it ensures to have information for all the regions of the countries under analysis, and therefore it maximizes the precision of the estimates.

reports the corresponding results. Again, I perform the regressions for the whole population (Panel A) and for men and women separately (Panels B and C, respectively). The first three columns report the OLS estimates, columns (4) and (5) the corresponding 2SLS estimates obtained using the enclave-based instrument[41], and, finally, columns (6) and (7) the 2SLS estimates obtained using the change in the push-factor and distance-based instrument. More in detail, column (1) of each Panel shows the estimation of the coefficients of interest from the baseline specification (i.e. OLS with no controls or country-specific trends). The only significant coefficient is estimated for females. According to it, a 1 percent increase in the share of immigrants aged 15-64 is associated with a reduction in the probability of native women to acquire upper-secondary education by around -0.42 percent. Similarly, the OLS estimates of the specification with the controls (but not the country-specific trends) of column (2) present a negative and significant effect for the whole population, which seems to be fully driven by the impact on women. To be clear, the estimated effect in the whole population is -0.47 , and in the case of women -0.72. In the case of men the coefficient of interest is also negative although much lower in magnitude and not precisely estimated. The inclusion of the country trends reduces quite substantially the precision and the magnitude of the estimates, which suggests that part of the effect reported in column (2) may correspond to the (country-specific) trends in education and immigration.

As for the 2SLS estimates relative to the enclave-based instrument, the effect of interest is negative and statistically significant for the three groups considered. In the case of the whole population, point estimates identify a negative native education response to immigration that is -0.98 in the specification that does not control by country-specific trends, and that decreases slightly when these trends are included (to -0.82), although remains significant.

Interestingly, the effect of immigration on education is more intense for women than for men (the effects for women is indeed double in magnitude than that of men). The estimates are robust to the inclusion of country trends, although their magnitude is slightly reduced (the point estimates now are -0.64 for men and -1 for women). Also in the case of the push-factor and distance-based instrument the effect of interest is negative and significant even after including the country-specific trends. Specifically, point estimates imply that a unitary percent increase in the share of immigrants aged 15-64 is associated with a reduction in the probability to invest in education of around -1.3 percent for the whole population, -0.9 for men and -1.4 for women. Reassuringly, both instruments used appear not to be weak, as the Kleibergen and Paap (2006) F-statistics are always fairly high[42]. In addition, it is important to highlight that the estimated impacts when using the two instruments is quite similar and fairly different to that of the OLS. This will somehow work in favor of the validity of the two instruments.

[41] It is important to notice that the 2SLS estimates of columns (4) and (5) are obtained using the enclave-based instrument for people aged 15-64. However, they are robust to the use of the instrument considering all the individuals reported in the census. I do not show the estimates relative to this last option to be more consistent with Hunt (2017), but they are available upon request.

[42] It is important to notice that the Kleibergen and Paap (2006) F-statistics vary between the different groups because each of them have a specific regression weights, and also because they are different samples.

Table 4.2: Overall effect.

	OLS			2SLS			
				enclave-based		push-distance	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Panel A: Whole population							
$\Delta m^{15-64}_{r,t-10}$	-0.185	-0.471**	-0.140	-0.982***	-0.815***	-0.643*	-1.136***
	(0.205)	(0.210)	(0.225)	(0.216)	(0.257)	(0.356)	(0.402)
Controls	NO	YES	YES	YES	YES	YES	YES
Country trends	NO	NO	YES	NO	YES	NO	YES
R^2	0.377	0.519	0.684	0.497	0.654	0.517	0.619
First-Stage F-stat	-	-	-	67.4	34.4	31.2	47.5
Panel A: Men							
$\Delta m^{15-64}_{r,t-10}$	0.024	-0.234	-0.013	-0.671***	-0.636**	-0.238	-0.867**
	(0.226)	(0.224)	(0.235)	(0.236)	(0.266)	(0.355)	(0.367)
Controls	NO	YES	YES	YES	YES	YES	YES
Country trends	NO	NO	YES	NO	YES	NO	YES
R^2	0.197	0.397	0.587	0.351	0.558	0.370	0.532
First-Stage F-stat	-	-	-	67.8	35.2	28.7	46.4
Panel A: Women							
$\Delta m^{15-64}_{r,t-10}$	-0.417**	-0.724***	-0.264	-1.311***	-0.995***	-1.043***	-1.360***
	(0.196)	(0.221)	(0.239)	(0.225)	(0.274)	(0.402)	(0.470)
Controls	NO	YES	YES	YES	YES	YES	YES
Country trends	NO	NO	YES	NO	YES	NO	YES
R^2	0.496	0.589	0.704	0.567	0.677	0.583	0.645
First-Stage F-stat	-	-	-	71	35.9	36	50.9

***Note**: The table reports the estimates of the overall effect, as indicated in equation (4.3). All the RHS variables are lagged one census. All regressions include time fixed-effects and are weighted by $1/(1/w_t + 1/w_{t-10})$. In all specifications the number of observations is 258. Standard errors, in parentheses, are clustered at the province level.*

*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Considering that the overall effect is the net effect of the direct and indirect ones, a negative overall effect can indicate two things. First, assuming that direct and indirect effects have opposite signs (as, for instance, estimated by Hunt, 2017, or as found in the OLS specifications that control for country trends of Table 4.1), a negative overall effect can mean that the schooling competition mechanism (i.e the direct effect) dominates the labor market one (i.e. the indirect effect). Second, this result may also imply that the two effects in the case of the European countries considered go in the same direction and are both negative. This may be driven by the fact that, in the countries considered, immigrants are less substitute to natives in the labor markets. Indeed, Europe in general, and the countries under analysis in particular, typically attract low-skilled immigrants, also because of the skill-downgrading mechanism that these experience upon arrival (Hanson, 2009). Therefore, natives do not feel threatened by their presence and do not need to acquire further human capital to avoid immigrants' competition once in the labor market. Additionally, it is interesting to notice that the impact of immigrants (both OLS and 2SLS) seem to be particularly large by native females. This might be driven by the stronger complementarity in production that exists between immigrants and native females (also confirmed by the two previous chapters of the thesis).

Robustness Checks

To be sure that the previous results are not driven by choices made on the empirical specification, I carry out a set of robustness checks.

First, I modify equation (4.3) by substituting the share of immigrants aged 15-64 with the share of immigrants aged 03-64. This new measure of immigrant-driven supply shock allows me to consider immigrants students of all schooling levels and may provide more accurate estimates of the impact of immigrants on native investment in human capital. Table 4A9 shows the 2SLS estimates that refer to this new specification, where I instrument the share of immigrants aged 03-64 with the shift-share variable for people of the same age (columns 1 and 2) and with the push-factors and distance based instrument (columns 3 and 4). The results are in line with the original ones as the coefficients associated to the change in the lagged share of immigrants aged 03-64 are always negative and statistically significant, and similar in magnitude to the ones reported above.

As previously indicated, in the estimations of the effects of interest I use a weighting scheme based on the inverse of the squared standard errors of the regional schooling variable estimated in equation (4.1). This is a fairly standard procedure and is in line with Hunt (2017). However, the use of regression weights is also relatively controversial (see Solon et al., 2015). To address this issue, and also to verify that the results are not driven by the use of weights, I provide some estimates of the overall effect for the different demographic groups where (i) I change the weighting scheme; and (ii) I estimate the effects of interest from the unweighted regressions. Table 4A10 provides the results. In columns (1), (2), (5) and (6) I use the population in each region-year cell as weight, while in columns (3), (4), (7) and (8) I run the unweighted regressions. As for the first case, the results are in line with the original ones for both the enclave-based and the push-factor and distance-based instruments[43]. When I instead remove the regression weights, in the case of the enclave-based instrument the sign and the magnitude of the coefficients of interest are not affected significantly, although the standard errors increase and the first-stage F-statistics reduce quite considerably. In the case of the push-factor and distance-based instrument the estimates of interest are always not significant. I therefore conclude that the use of an appropriate weighting scheme is preferred as, besides controlling for heteroskedasticity (see Solon et al., 2015), allows me to obtain more precise estimates.

The extant literature indicates that the investment in education is strongly connected with the business cycle. So, next I check whether the negative overall effect previously estimated is driven by particularly favorable economic conditions in the countries under analysis. To do so, I include as an additional control the Bartik variable (Bartik, 1991) that is typically used as an arguably exogenous proxy for shocks on the labor demand side (Basso and Peri, 2015; Özgüzel, 2021). I build the Bartik variable as follows. I initially obtain the predicted employment for every industry[44] by interacting the initial regional employment share in every industry with the total employment by industry at the national level in every census year. Then, I sum-up the interactions across industries and obtain the predicted employment-to-population rate by dividing the predicted employment in each region-year cell with the corresponding working-age population.

I report the estimates relative to the inclusion of the Bartik control in Table 4A11. Some interesting aspects emerge. First, there is a fairly clear negative relationship between

[43]The results also hold when I use as weights the number of observations used to compute the regional schooling variable (I do not show these results here, but are available upon request).

[44]I use 15 different industries.

the Bartik shock and the propensity of natives to acquire human capital, confirming the counter-cyclicality of the accumulation of education. Second, in the case of the shift-share instrument, the coefficients on the share of immigrants aged 15-64 are always negative and statistically significant at 5% for the whole population and women and only at 10% for men, confirming the negative relationship between immigrants and native schooling in the countries and during the period considered. However, the magnitude of the coefficients is slightly reduced, highlighting the importance of controlling for labor demand shocks in this type of models. In the case of the push-factors and distance-based instrument, when controlling for local demand shocks the impact of immigration is somehow reduced. However there is still evidence of a negative effect in the case of the full sample that seems to be particularly driven by the response of native women.

The results on the overall effect show a negative impact of immigration on natives' education. To be sure that this effect is not driven by some changes in the composition of the workforce, nor by population (geographical) displacements, in this section I estimate the model sketched in equation (4.3) by adding two additional variables that typically control for this, namely immigrants' and natives' average ages (Altonji and Card, 1991; Dustmann et al., 2013; Özgüzel, 2021). Table 4A12 shows the corresponding estimates. As it is possible to verify, the coefficients associated to the immigrant share are in line with those presented in Table 4.2, implying that in general the estimates of the impact of immigration on natives education are not driven by changes in the composition of the workforce. The coefficients associated to immigrants' and natives' average ages are negative and very close to zero, and are only significant in the specifications that control for country trends.

The literature on the determinants of school drop-outs indicates that the demographic structure of the population might also be correlated with its schooling decisions (Gutiérrez-Domènech, 2011). To address this concern, in the main analysis I control for the native cohort size. However, to further test the robustness of the main results, in Table 4A13 I additionally control for changes in the (lagged) population share, that is the ratio between individuals aged 15-24 and the total working-age population and in the (lagged) male share, defined as the share of young individuals (i.e. 15-24) that are men. Again, the results are not substantially affected by the inclusion of these variables, whose associated coefficients are, in turn, never statistically significant (with the exception of the one on male share in the specification that does not control for country trends estimated for women).

Another important aspect that may influence the schooling decision of the population is the employment structure of the local labor markets (Aparicio-Fenoll, 2016; Diaz-Serrano and Nilsson, 2020). To verify whether the results are robust to indicators of the (local) employment structure, in Table 4A14 I add as further controls the changes in the (lagged) share of workers in agriculture, construction and wholesale and retail services. Two interesting things emerge. First, when controlling for country-specific trends, the coefficients on the immigrant share are always negative and statistically significant and their magnitude is slightly increased. Second, the coefficients on the shares of workers in agriculture and services are negative and significant, except the coefficients on the share in agriculture of column (4). This confirms the presence of a negative relationship between low-skilled labor demand and the decision to invest in human capital. Nevertheless, it is important to consider that the employment shares are likely to be endogenous[45] and therefore these results

[45]The fact that the employment shares are lagged one census might not be enough to ensure their exogeneity, particularly in presence of serially correlated labor demand and supply shocks that affect the two sides of the regression equation.

of this section should be taken with caution.

4.6 Heterogeneity Analysis: Effect of Changes in Local Specialization in Low-skilled Sectors

4.6.1 Empirical Strategy

In the second part of this chapter, I perform an analysis of the heterogeneity of the impact of immigration on natives education. Specifically, I study how this impact varies with the specialization in low-skilled sectors of the local economy. In other words, I want to verify whether the different industrial structure of the regions, combined with the uneven intensity in the migratory inflows that these receive, may somehow modify the natives education responses estimated in the previous sections. This is further justified by the fact that the countries considered are characterized by large regional disparities not only in terms of the industrial composition in general and by their specialization in low-skilled sectors in particular, but also of the inflows of immigrants that receive, and of the education distribution of their population (Fonseca, 2017). Regions like Attiki in Greece, Dublin in Ireland, Grande Lisboa in Portugal, or Barcelona and Madrid in Spain are indeed expected to have a substantially different employment structure, a higher immigrant density, and a higher share of high-educated individuals, with respect to more peripheral regions in the same countries (European Commission, 2010; Fonseca and Fratesi, 2017). This fairly high regional disparities provide an interesting source of variation that can be exploited empirically to test whether differences in specialization in low-skilled sectors of the regions under analysis implies different native education responses to immigration.

This heterogeneity analysis is motivated by two separate strands of the extant literature. First, the literature on the determinants of school drop-outs indicates that the sector specialization of a region is a key factor in explaining the educational attainments of its population. Indeed, regions specialized in sectors that make a prevalent use of low-skilled labor are often characterized by a higher drop-out rate. This is because individuals less prone to studying may not have incentives to acquire additional education as they can easily find a job position. This phenomenon is even further magnified in periods of economic expansion of the low-skilled sectors, when the wage offered is often higher than the expected return to schooling, net of its costs (Aparicio-Fenoll, 2016; Diaz-Serrano and Nilsson, 2020). This is confirmed by the results reported in Table 4A14, where the coefficients associated to the share of workers in agriculture and retail services (that are typically sectors that require a low amount of education) are negative and significant. In addition, after their inclusion the coefficients on the share of immigrants is increased in magnitude (i.e. more negative).

Second, the literature of the skill-biased technical change shows that, in reaction to an increase in the supply of low-skilled labor, for instance induced by immigrants, firms adjust their capital-labor ratios by employing more low-skilled labor and less capital. There are two channels through which local economies may absorb a low-skilled-biased supply shock. The first is the between change: the increase in the supply of low-skilled labor is absorbed by changes in the industry mix of the local economy. In other words, firms of the low-skilled sector becomes more productive than those in the high-skilled one and experience higher productivity gains. The second is instead the within change: the supply shock is absorbed through changes in the technologies of production of each firm. More precisely, firms switch to technologies that make prevalent use of the more abundant factor type (Clemens et al., 2018; González and Ortega, 2011; Lewis, 2011; Peri, 2012). If these mechanisms occur, then natives have less incentives to acquire more human capital.

Against this background, I go a step further by joining these two strands of the extant literature and analyze the combination of increased immigration and regional specialization in low-skilled industries and activities on the natives' propensity of investing in human capital.

In this context, the hypothesis to test is that in regions specialized in low-skilled industries, natives have lower incentives to invest in human capital, as they can easily find a job position. The presence of immigrants further magnify this mechanism as host regions typically absorb the increase in labor supply through an unskilled-biased technical change (as it happened in Spain during the recent housing boom, see González and Ortega, 2011). As already indicated, the focus is not on the degree of sector specialization, but rather on its dynamics. Indeed, I assume that native decisions to invest in human capital are particularly influenced by the evolution of local specialization in low skilled activities. In practical terms, I capture this by means of a specification in first-differences based on equation (4.3), which also allows to cancel out the region specific unobserved heterogeneity. In other words, I estimate the following model[46]:

$$\Delta\hat{\kappa}_{rt} = \delta_0 + \delta_1 \Delta m^{15-64}_{r,t-10} + \delta_2 \Delta S^{low}_{r,t-10} + \delta_3 (\Delta m^{15-64}_{r,t-10} \times \Delta S^{low}_{r,t-10}) + \delta_4 \Delta X_{r,t-10} + \psi_t + \upsilon_{rt} \tag{4.6}$$

where $\Delta S^{low}_{r,t-10}$ is the change in the lagged share of low-educated workers in sectors that require a low level of education (that, for simplicity, I call low-skilled sectors), while the other variables are defined as before. It is important to notice that this measure expressed in first-difference somehow captures the dynamics of sector specialization: negative values mean that low-tech sectors were losing weight in local employment, whereas positive ones indicate the opposite. In this setting the effect of interest is given by $\delta_1 + \delta_3 \cdot \Delta S^{low}_{r,t-10}$ (that is, the effect of immigration shocks on natives education).

However, the share of workers in low-skilled sectors is likely to be endogenous as it may be correlated with the same unobservable factors that also influence the regional schooling variable (i.e. the dependent variable). To solve this problem, I control for region- and year-specific unobserved heterogeneity by means of the specification in first-differences and the inclusion of year fixed-effects. Moreover, there's might be a reverse causality bias as not only the share of workers in low-skilled sectors may influence the native educational attainment, but also the other way around. Indeed, an increase in the drop-out rate in some regions might be absorbed through a higher growth in the low-skilled sector (Diaz-Serrano and Nilsson, 2020). I address the deriving endogeneity issue by using the ten-year lagged share of low-educated workers in low-skilled sectors. Unfortunately, however, in presence of serially correlated shocks that affect the two sides of the regression equation, this strategy might not be enough (Jaeger et al., 2019; Goldsmith-Pinkham et al., 2020). Therefore, I also build a Bartik-like variable for only low-skilled sectors that gets rid of the local shocks that might cause the endogeneity of the original sectoral specialization measure. The variable is

[46] For the sake of simplicity, and also given the low first-stage F-statistics estimated when using different age- and education-specific instruments, I perform the heterogeneity analysis only on the overall effect. In any case, the interaction between the immigrant shocks and the local employment structure is assumed to influence only the indirect effect of immigration.

defined as follows:

$$B_{r,t}^{low} = \frac{\widehat{E}_{r,t}}{L_{r,t}} \quad \text{where} \quad \widehat{E}_{r,t} = \sum_j \left(\frac{E_{r,j,t_0}}{E_{j,t_0}} \right) \cdot E_{j,t}^{H<up-sec}$$

In other words, I obtain the predicted employment in low-skilled sectors in each region by multiplying the share of low-educated workers (i.e. with less than upper-secondary education) residing in region r employed in low-skilled industries j in year $t_0 = 1981$ ($E_{r,j,t_0}/E_{j,t_0}$) by the stock of workers with less than upper-secondary education employed in low-skilled industries j at the national level in all the subsequent years ($E_{j,t}^{H<up-sec}$). Then, to obtain the Bartik variable, I divide the predicted employment previously computed by the the total working-age population. There is a high, positive and statistically significant correlation (around 0.7) between the share of low-educated workers in low-skilled sectors and the Bartik variable for low-skilled sectors, implying that the latter is a good predictor for the former, whereas in principle it does not include (most of) the elements that induces endogeneity in the original regressor. Finally, I interact the (change in the) Bartik variable so computed with the immigrant variable. In other words, I modify equation (4.6) as follows:

$$\Delta\hat{\kappa}_{rt} = \lambda_0 + \lambda_1 \Delta m_{r,t-10}^{15-64} + \lambda_2 \Delta B_{r,t-10}^{low} + \lambda_3 (\Delta m_{r,t-10}^{15-64} \times \Delta B_{r,t-10}^{low}) + \lambda_4 \Delta X_{r,t-10} + \psi_t + \iota_{rt} \tag{4.7}$$

where $\Delta B_{r,t-10}^{low}$ is the change in the one-census lagged Bartik variable for low skilled sectors defined previously. In line with equation (4.6), the effect of interest is given by $\lambda_1 + \lambda_3 \cdot \Delta B_{r,t-10}^{low}$.

4.6.2 Results

I estimate equation (4.6) by OLS and equation (4.7) by 2SLS, where I instrument the immigrant share with alternatively the enclave-based and the push-factor and distance-based instruments[47]. The corresponding results are reported in Table 4A15. The focus of the analysis is the combined effect of immigration and the dynamics of sector specialization on native education. In other words, the more interesting aspect of the analysis is to verify whether, and to which extent, the effect of immigrants varies in regions characterized by a different evolution of low-skilled sectors specialization. I do so in Table 4.3, where I report (i) the average marginal effect of the change in the lagged immigrant share in working-age (first two rows of each panel), and (ii) the marginal effect of the change in the lagged immigrant share in working-age computed at different percentiles of the change in the lagged Bartik variable for low-skilled sectors. As for the first, in the case of the enclave based instrument (Panel A), the preferred specification (i.e. the one with country-specific trends), indicate that the native education response to immigration is, on average, negative. More precisely, in the case of the whole native population the point estimate indicate a reduction in the probability of acquiring upper-secondary education by around -0.5 percent, which is only significant at 5% level. Interestingly, the gender distinction indicates that the parameter estimate for men is not statistically significant, while for women is negative and significant

[47]In other words, in the 2SLS approach the only variable that I instrument is the share of immigrants. The Bartik variable for the low-skilled share in local employment is used as (exogenous) regressor instead of the original variable.

at 1% level, implying a reduction in the probability of acquiring upper-secondary education by around -0.8 percent. Analogously, the results relative to the push-factor and distance-based instrument imply eduction in the probability of acquiring upper-secondary education by around -0.9% for the whole population, -0.6% for men (although only significant at 10%), and -1% for women.

As for the second, the results of the preferred specification imply that in regions that are to some extent loosing employment in low-skilled sectors, the native education responses are positive but not significant in most cases. On the contrary, in regions that are loosing it a a slower pace, or even increasing their specialization in low-skilled sectors, the native education response is negative and statistically significant. In other words, the coefficients associated to the change in the lagged immigrant share get negative and increase in magnitude as long as the change in the lagged Bartik variable for low-skilled sectors increases. The estimated effects indicate a reduction in the probability of natives to acquire upper-secondary education that vary between -1.2 to -2.4 percent for the whole native population, -1 to -2 percent for men, and -1.4 to -2.6 percent for women. In line with the initial hypothesis, these results imply that in regions that experienced a switch in their employment structure towards low-skilled activities the negative native education responses to immigration are stronger.

As stated above, the previous results are obtained considering the change in the local specialization in low-skilled sectors. However, it may be the case that the regions must have a large change in the share of employment in low-skilled sectors in order to experience a higher impact of immigration shocks on natives education. To this extent, an alternative way to measure the industrial specialization of a region is to use a dummy variable that takes the value of 1 in specialized regions and 0 otherwise. Against this background, I define a set of different dummy variables based on different values of the change in the Bartik variable for low-skilled sectors distribution[48].

In Table 4.4 I report the average marginal effects, as well as the marginal effects of immigration depending on the regional specialization in low-skilled sectors[49]. In Panel A, I define the dummy based on the 50^{th} percentile of the change in the Bartik variable, while in Panels B and C based on the 75^{th} and 90^{th} percentiles, respectively. The results highlight different things. First, the average marginal effects are negative and significant in all Panels. Second, in my preferred specifications (the ones that control for country-specific trends), the coefficients associated to the dummy variable are typically negative. This is valid both in regions not specialized in low-skilled sectors (that is, those for which the dummy variable takes the value of 0) and in regions specialized in low-skilled sectors (those for which the dummy variable takes the value of 1). The only exception is represented by Panel B, where the effect in regions specialized in low-skilled sectors is only significant at 10% level. However, in the latter case the effect is considerably higher, implying that the combination between immigration and sector specialization induces a stronger native education response[50].

[48] I define the dummy based on the (change in the) Bartik variable, and not on the share of workers in low-skilled sectors, because is a more exogenous measure of regional specialization, as previously discussed.

[49] The 2SLS estimates (obtained with the enclave-based instrument) of the interaction effect between the change in immigrants in working age and the dummy variables are reported in Table 4A16.

[50] Table 4.4 only reports the 2SLS estimates produced instrumenting the immigrant share with the enclave-based variable. Those relative to the push-factors and distance-based instrument are in line with them and are available upon request.

Table 4.3: Marginal effects of immigration depending on the evolution of local specialization in low-skilled sectors.

	All		Men		Women	
	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: 2SLS estimates with enclave-based instrument						
AME	-0.841***	-0.541**	-0.441**	-0.307	-1.286***	-0.770***
	(0.186)	(0.256)	(0.199)	(0.267)	(0.195)	(0.279)
-0.20	1.344	0.535	2.791***	1.217	-0.246	-0.093
	(0.925)	(0.961)	(0.940)	(1.012)	(1.005)	(1.052)
-0.15	0.622	0.180	1.730***	0.717	-0.591	-0.318
	(0.639)	(0.702)	(0.655)	(0.743)	(0.694)	(0.767)
-0.10	-0.100	-0.176	0.670*	0.216	-0.936**	-0.543
	(0.367)	(0.455)	(0.385)	(0.486)	(0.396)	(0.496)
-0.05	-0.821***	-0.531**	-0.391*	-0.284	-1.282***	-0.767***
	(0.187)	(0.260)	(0.200)	(0.274)	(0.196)	(0.281)
-0.03	-1.110***	-0.673***	-1.452***	-0.784***	-1.627***	-0.992***
	(0.209)	(0.229)	(0.325)	(0.264)	(0.353)	(0.298)
0.03	-1.976***	-1.099***	-2.512***	-1.285***	-1.972***	-1.217***
	(0.487)	(0.383)	(0.587)	(0.470)	(0.645)	(0.525)
Country trends	NO	YES	NO	YES	NO	YES
First-stage F-stat.	16.3	7.6	16.6	7.9	17.4	8
Panel B: 2SLS estimates with push-factor & distance-based instrument						
AME	-0.608	-0.877**	-0.242	-0.636*	-0.980**	-1.079***
	(0.402)	(0.380)	(0.411)	(0.374)	(0.434)	(0.413)
-0.20	3.909***	2.040	4.278***	2.281*	3.400**	1.760
	(1.412)	(1.383)	(1.487)	(1.244)	(1.520)	(1.665)
-0.15	2.417**	1.077	2.794***	1.323	1.945*	0.818
	(1.001)	(0.971)	(1.060)	(0.909)	(1.066)	(1.141)
-0.10	0.925	0.113	1.310**	0.366	0.492	-0.124
	(0.626)	(0.594)	(0.666)	(0.598)	(0.655)	(0.659)
-0.05	-0.566	-0.850**	-0.172	-0.591	-0.962**	-1.067***
	(0.404)	(0.381)	(0.414)	(0.379)	(0.434)	(0.413)
-0.03	-1.163***	-1.235***	-0.766*	-0.974***	-1.543***	-1.444***
	(0.422)	(0.410)	(0.420)	(0.359)	(0.472)	(0.483)
0.03	-2.953***	-2.391***	-2.547***	-2.124***	-3.288***	-2.575***
	(0.767)	(0.775)	(0.765)	(0.573)	(0.880)	(0.999)
Country trends	NO	YES	NO	YES	NO	YES
First-stage F-stat.	10.9	23.5	10.7	22.6	11.3	25.5

***Note**: The table reports the marginal effects of the change in the (lagged) share of immigrants aged 15-64 computed at different percentiles of the distribution of the change in the (lagged) Bartik variable for low-skilled sectors. In all specifications the number of observations is 258. Standard errors, in parentheses, are clustered at the province level.*

*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table 4.4: Marginal effects. Specialization in low-skilled sectors based on dummy variables.

	All		Men		Women	
	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: dummy-low=1 if $\Delta B^{low}_{r,t-10} \geq 0.31$ (around 50^{th} percentile)						
AME	-1.620***	-1.696***	-1.032**	-1.235***	-2.206***	-2.154***
	(0.524)	(0.409)	(0.522)	(0.402)	(0.564)	(0.466)
dummy-low=0	-0.693***	-0.877***	-0.334	-0.665**	-1.068***	-1.089***
	(0.247)	(0.280)	(0.264)	(0.287)	(0.267)	(0.299)
dummy-low=1	-5.154**	-4.820***	-3.610*	-3.342***	-6.768***	-6.429***
	(2.063)	(1.315)	(1.952)	(1.199)	(2.319)	(1.653)
Country trends	NO	YES	NO	YES	NO	YES
R^2	0.572	0.632	0.462	0.551	0.600	0.643
First-stage F-statistic	3.1	2.9	3.2	2.9	3.1	2.9
Panel B: dummy-low=1 if $\Delta B^{low}_{r,t-10} \geq 0.39$ (around 75^{th} percentile)						
AME	-1.360***	-1.112***	-1.101***	-0.941***	-1.677***	-1.326***
	(0.340)	(0.352)	(0.353)	(0.362)	(0.362)	(0.380)
dummy-low=0	-0.826***	-0.840***	-0.512**	-0.660**	-1.163***	-1.023***
	(0.212)	(0.258)	(0.233)	(0.268)	(0.222)	(0.274)
dummy-low=1	-7.733**	-4.347*	-7.811***	-4.154*	-8.198**	-5.166*
	(3.120)	(2.483)	(2.884)	(2.353)	(3.625)	(2.987)
Country trends	NO	YES	NO	YES	NO	YES
R^2	0.576	0.653	0.439	0.555	0.625	0.677
First-stage F-statistic	13.4	17.2	12.3	15.5	14.4	18.5
Panel C: dummy-low=1 if $\Delta B^{low}_{r,t-10} \geq 0.47$ (around 90^{th} percentile)						
AME	-1.179***	-0.931***	-0.854***	-0.729***	-1.525***	-1.136***
	(0.224)	(0.274)	(0.245)	(0.284)	(0.234)	(0.291)
dummy-low=0	-0.977***	-0.829***	-0.666***	-0.648**	-1.307***	-1.011***
	(0.214)	(0.258)	(0.235)	(0.268)	(0.224)	(0.275)
dummy-low=1	-12.201***	-6.476***	-10.966***	-5.108**	-13.831***	-8.208***
	(1.627)	(2.325)	(1.612)	(2.323)	(1.856)	(2.461)
Country trends	NO	YES	NO	YES	NO	YES
R^2	0.519	0.659	0.372	0.561	0.585	0.682
First-stage F-statistic	33.3	17.1	33.5	17.5	35.1	17.8

Note: *The table reports the average marginal effects and the marginal effects depending on the regional specialization in low-skilled sectors. The latter is defined by means of a set of dummy variables based on the distribution of the change in the Bartik variable for low-skilled sectors. In all specifications the number of observations is 258. Standard errors, in parentheses, are clustered at the region level.*
*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

4.7 Adjustment Mechanisms: Internal Migration

Contrarily to Hunt (2017), I have estimated a negative effect of immigration on native education, which is particularly intense in regions highly specialized in low-skilled sectors. Assuming that this effect was mostly due to the indirect effect, one of the potential reasons behind such result is internal migration. The negative effect may indeed be driven by the fact that high-educated natives respond to immigrant shocks by moving away from regions that receive larger inflows. Indeed, it is well-known that education is key determinant of migration and high-educated individuals are often considered as more mobile than low-educated ones (Bover and Arellano, 2002; Sardadvar and Rocha-Akis, 2016; Amior, 2019). Moreover, the inflow of low-skilled immigrants may induce firms to move to labor-intensive technologies and reduce their investment in capital. This phenomenon often creates wage differentials across regions that may induce high-educated natives to migrate towards regions that instead are capital-intensive and where they can obtain a job position that would not imply a skill-downgrading (Berry and Glaeser, 2005; Hean et al., 2020; Betz et al., 2015). To this extent, immigrants are indeed assumed to *grease the wheel of the labor market*, not only because they are highly responsive to local economic shocks (as highlighted by Borjas, 2001; Basso et al., 2019; Cadena and Kovak, 2016; Amior, 2020), but also because they induce natives to relocate to regions offering better economic opportunities, as argued by Albert and Monras (2020), Monras (2020), Monras (2021) and Amior (2021). In order to assess the native internal migration responses to increased immigration, I follow the existing literature and estimate the following model for the group of countries under analysis:

$$\frac{\Delta N_{r,t}}{L_{r,t-10}} = \pi_0 + \pi_1 \frac{\Delta M_{r,t}}{L_{r,t-10}} + \psi_t + \mu_{rt} \tag{4.8}$$

where the dependent variable captures the inter-census change in the native working-age population ($\Delta N_{r,t} = N_{r,t} - N_{r,t-10}$), standardized by the lagged total working-age population ($L_{r,t-10}$), and the main independent variable is defined similarly for the immigrant population. Finally, ψ_t is a set of census-year fixed-effects that control for country-wide time-varying unobserved heterogeneity. The equation is assumed to identify the native relocation responses to immigration. Again, I control for the non-random immigrants settlement across regions using a shift-share instrument[51]. The parameter of interest is π_1. An estimate of $\pi_1 = 0$ implies absence of native internal migration in reaction to immigrant shocks, while an estimate of $\pi_1 < 0$ (or $\pi_1 > 0$) indicates that natives "vote with their feet" and move away (or into) regions characterized by the higher presence of immigrants. I report the estimates of the parameters of interest Table 4.5.

When I do not control for labor demand shocks (columns 1, 2, 4 and 5), both OLS and 2SLS identify a positive native relocation response. In other words, natives of all education groups seem to move into regions characterized by the higher presence of immigrants. However, in the case of the 2SLS specifications. When I control for Bartik-like labor demand shocks (column 6), the precision of the estimates reduces considerably. Indeed, the point estimate for the whole population becomes only marginally significant, and that of the low-skilled native becomes not significant. On the contrary, the effect on high-educated natives is not affected by the inclusion of the Bartik shocks and remains positive and statistically significant at 1% level. The magnitude of the coefficient implies that a unitary percentage increase in the standardized immigrant population is associated with an increase in the native population of around 0.3 percent. This is likely to be due to the fact that high-educated natives are more complementary in production with immigrants than their low-skilled counterparts. Immigrants are indeed assumed to induce a task-upgrade of native workers (Peri and Sparber, 2009; Foged and Peri, 2016; Cattaneo et al., 2015; D'Amuri and Peri, 2014) that might positively affect their productivity (Peri, 2012, 2016) and may ultimately underlie the tendency of natives (particularly high-skilled) to locate in regions with higher immigrant density. My estimates are in line with the analyses of Mocetti and Porello (2010) for the Italian economy, Sanchis-Guarner (2017) for Spain and Monras (2021) for the U.S.

[51] I define the instrument based on Basso and Peri (2015) and Sparber and Zavodny (2020). In other words,

$$\Delta \widehat{m}^{std}_{r,t} = \frac{\widehat{M}_{r,t} - \widehat{M}_{r,t-10}}{\widehat{M}_{r,t-10} + N_{r,t-10}} \quad \text{where } \widehat{M}_{r,t} = \sum_o \left(\frac{M_{o,r,t_0}}{M_{o,t_0}} \right) \cdot M_{o,t}$$

Table 4.5: Native relocation responses to immigration.

	OLS			2SLS		
	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: All Natives						
$\Delta m^{std}_{r,t}$	0.573***	0.558***	0.593***	0.534***	0.506***	0.551*
	(0.089)	(0.101)	(0.112)	(0.159)	(0.187)	(0.283)
Bartik shocks			-0.068			-0.049
			(0.118)			(0.170)
Country trends	NO	YES	YES	NO	YES	YES
R^2	0.257	0.289	0.291	0.256	0.288	0.291
First-stage F-stat	-	-	-	39.4	31.8	18.1
Panel B: High-educated Natives						
$\Delta m^{std}_{r,t}$	0.379***	0.353***	0.368***	0.362***	0.328***	0.346***
	(0.041)	(0.047)	(0.057)	(0.071)	(0.083)	(0.126)
Bartik shocks			-0.030			-0.020
			(0.042)			(0.062)
Country trends	NO	YES	YES	NO	YES	YES
R^2	0.415	0.425	0.426	0.415	0.424	0.425
First-stage F-stat	-	-	-	39.4	31.8	18.1
Panel C: Low-educated Natives						
$\Delta m^{std}_{r,t}$	0.311***	0.301***	0.317***	0.290**	0.292**	0.324
	(0.079)	(0.076)	(0.092)	(0.122)	(0.125)	(0.199)
Bartik shocks			-0.031			-0.034
			(0.108)			(0.143)
Country trends	NO	YES	YES	NO	YES	YES
R^2	0.112	0.273	0.273	0.112	0.272	0.273
First-stage F-stat	-	-	-	39.4	31.8	18.1

Note: *The table reports the estimates of the native relocation responses to immigration, as specified in equation (4.8). Panel A refers to the whole native population, Panel B to high-skilled natives (i.e. with upper-secondary education or more) and Panel C to low-skilled ones (i.e. with less than upper-secondary education). All regressions include time fixed-effects and are weighted by the region-year cell population at the beginning of the period. In all specifications the number of observations is 258. Standard errors, in parentheses, are clustered at the province level.*
*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

In terms of the analysis on native schooling, these results imply that the negative effect estimated in the previous sections is unlikely to be driven by some sort of internal relocation of native workers.

4.8 Conclusions

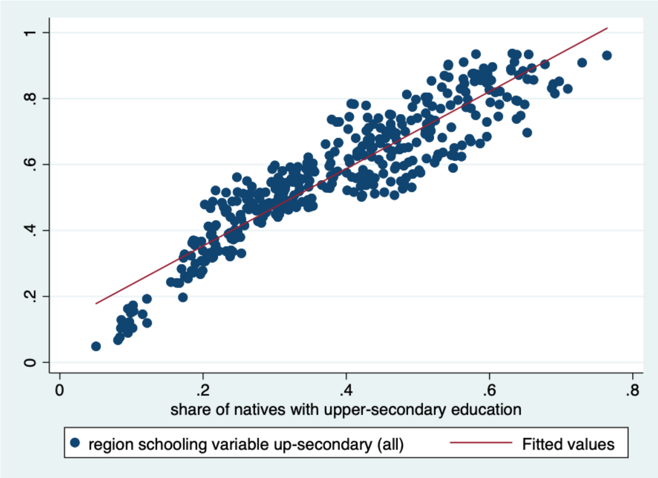
Over the last few decades, the European economies have experienced an increase in the migratory inflows. Surprisingly, this has occurred non only in countries with a longer tradition of immigration, such as Germany or the U.K., but also in countries less used to it, like Spain or Greece. In this chapter, I exploit this large and unprecedented increase in the migratory phenomenon and analyze the native education responses to immigration in a set of European countries - that is, Portugal, Ireland, Greece and Spain - over the period 1981-2011. Indeed, one of the potential consequences of immigration is that it may alters the skill distribution of the regions of destination. This is particularly valid in the countries under analysis that mostly host low-skilled immigrants, also because of the skill-downgrading that they typically experience upon arrival. This, in turn, implies an increase in the supply of low-skilled labor and a contemporaneous decrease in the relative supply of high-skilled workers. Both these changes imply an increase in the return to skills that can be matched by an increase in the investment in human capital of the native population.

From a methodological point of view, this chapter consists into two main parts. In the first, I initially assess the direct and indirect effect of immigration on native schooling. I do so by regressing a composition-adjusted regional schooling variable for natives on a set of age- and education-specific immigrant shares, whose endogeneity is addressed by a set of age- and education-specific shift-share instruments. The corresponding 2SLS results do not identify any significant impact. However, these results must be taken with caution because of the low first-stage Sanderson and Windmeijer (2016) F-statistics that is likely to be driven by the high correlation between the different instruments. I overcome this issue by estimating the overall effect of immigration, that is identified by means of a single immigrant measure, namely the share of immigrants in working-age. After adjusting again for the non-random immigrants' location decisions, I estimate a negative effect that implies that a unitary percentage increase in the share of immigrants in working-age is associated with a decrease in the probability of natives to acquire upper-secondary education that varies between 0.6 to 1.3 percent. All in all, assuming that the omission of relevant variable bias is minimized by the inclusion of a comprehensive set of control variables, census-year and region-specific fixed-effects as well as region specific trends, the instruments' validity suggests that these estimates are likely to identify a causal relationship between immigration and native education.

These results, however, are relative to the "average" region, and fail to take into account in a more systematic way another crucial determinant of native schooling besides immigration: the employment structure of the different local labor markets where they live. I address this concern in the second part of the analysis, where I estimate the combined effect of immigration and regional sectoral composition on native propensity to invest in human capital. Taking into account the set of countries under analysis, I consider the regional specialization in low-skilled sectors. After addressing the endogeneity concerns, I find that in regions specialized in low-skilled sectors the combined effect of immigration and specialization is negative. In addition, the native education responses to immigration tend to be stronger (i.e. more negative) as long as the specialization in low-skilled sectors increases. This is consistent with the idea that local economies absorb immigration-induced labor supply shifts by means of an unskilled-biased technical changes.

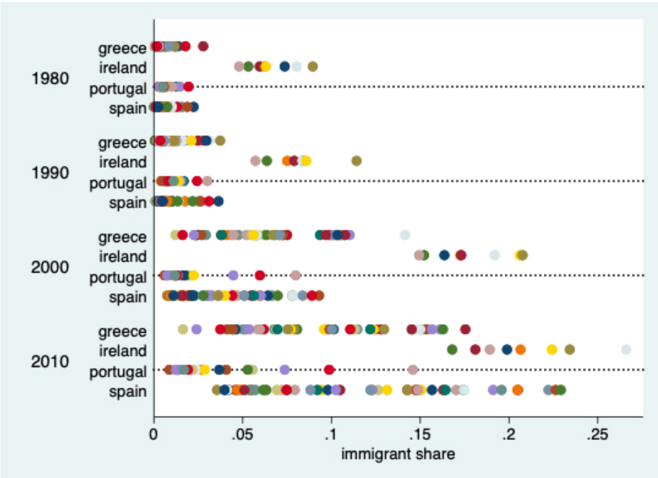
4.9 Appendix

Figure 4A1: Regional schooling variable vs simple share.



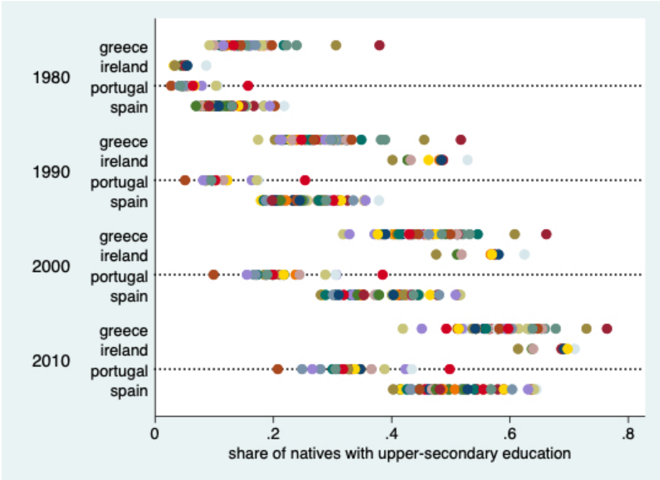
Note: *Correlation between the composition-adjusted regional schooling variable and the simple share of natives in upper-secondary education. The slope of the regression line is 1.17, with s.e.=0.025 and R-squared of 0.848.*

Figure 4A2: Immigrant share over the working-age population.



Note: *Evolution over time of the immigrant share over the working-age population. Each dot corresponds to a region within each of the countries under analysis.*

Figure 4A3: Share of natives with upper-secondary education.



Note: *Evolution over time of the share of natives with at least upper-secondary education. Each dot corresponds to a region within each of the countries under analysis.*

Figure 4A4: Evolution over time of the regional employment structure.

Note: *Evolution over time of the share of low-educated workers in low-skilled sectors. Each dot corresponds to a region within each of the countries under analysis.*

Table 4A1: Immigrant share over total population.

	Country			
Census	*Greece*	*Ireland*	*Portugal*	*Spain*
1981	0.0138	0.0699	0.0091	0.0059
1991	0.0148	0.0819	0.0128	0.0097
2001	0.0775	0.1810	0.0265	0.0446
2011	0.1013	0.2202	0.0472	0.1294

***Note**: The table reports the evolution over time of the immigrant share over the total population (working-age individuals only) in the period under analysis and in the countries considered. In the case of Ireland, I used the census of 2002 instead of 2001.* ***Source**: IPUMS.*

Table 4A2: Upper-secondary and compulsory education

	Age	
Country	*upper-secondary education*	*compulsory education*
Greece	15-18	15
Ireland	15-19	16
Portugal	15-18	18
Spain	15-18	16

***Note**: The table reports the duration of upper-secondary education, as well as the age at which compulsory education ends in the countries under analysis.* ***Source**: European Commission (2018).*

Table 4A3: Share of natives with at least upper-secondary education.

	Country			
Census	*Greece*	*Ireland*	*Portugal*	*Spain*
1981	0.2493	0.0565	0.0805	0.1367
1991	0.3814	0.4775	0.1428	0.2740
2001	0.5366	0.5675	0.2493	0.4024
2011	0.6639	0.6784	0.3638	0.5268

***Note**: The table reports the evolution over time of the share of natives with upper-secondary education (working-age individuals only) in the period under analysis and in the countries considered. In the case of Ireland, I used the census of 2002 instead of 2001.* ***Source**: IPUMS.*

Table 4A4: First-stage regressions.

Panel A: First-stage for equation (4.2)			
	Δ share imm.	Δshare imm. 19-64 with:	
	15-18	Less up-sec.	Equal/more up-sec
ΔEnclave 15-18	0.283***	0.087*	0.041
	(0.108)	(0.049)	(0.028)
ΔEnclave 19-64 with:			
Less up-sec.	0.187	0.244*	0.052
	(0.209)	(0.132)	(0.084)
Equal/more up-sec	0.196*	-0.227***	0.578***
	(0.113)	(0.051)	(0.057)
S.W. F-statistic	3.9	3.9	4
R^2	0.714	0.637	0.786

Panel B: First-stage for equation (4.3)			
	Δshare imm. in working-age (15-64)		
	(1)	(2)	(3)
ΔEnclave-based:			
whole	0.520***		
	(0.064)	-	-
15-65	-	0.444***	-
		(0.054)	
ΔPush-distance	-	-	0.116***
			(0.021)
F-statistic	65.4	67.4	31.2
R^2	0.700	0.702	0.622

***Note**: All regressions are weighted with the weights used in the second-stage regressions and include controls and year fixed-effects. In all specifications the number of observations is 258. Standard errors, in parentheses, are clustered at the province level.*
*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table 4A5: Reduced form regressions.

	(1)	(2)	(3)	(4)
ΔEnclave-based:				
Whole	-0.545***	-	-	-
	(0.127)			
15-64	-	-0.437***	-	-
		(0.105)		
15-18	-	-	-0.405***	-
			(0.139)	
ΔEnclave-based 19-64 with:				
Less up-sec.	-	-	-0.337	-
			(0.318)	
Equal/more up-sec	-	-	0.304	-
			(0.255)	
ΔPush-distance	-	-	-	-0.074*
				(0.044)
R^2	0.535	0.531	0.535	0.507

***Note**: The table reports the reduced form regression for all the age- and education specific enclave-based instruments used, and for the push-and distance-based one. All specifications are weighted with the weights used in the second-stage regressions and include controls and year fixed-effects. In all specifications the number of observations is 258. Standard errors, in parentheses, are clustered at the province level.*
*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table 4A6: Validity of the share component.

	Share component of			
	enclave-based		push & distance	
	(1)	(2)	(3)	(4)
Share natives with upper-secondary education	-0.087	-0.368	-0.179**	-0.090
	(0.103)	(0.508)	(0.084)	(0.175)
Share natives with tertiary education	0.010	0.254	-0.058	-0.169
	(0.201)	(0.679)	(0.110)	(0.256)
Youth unemployment rate	-0.042	-0.120	0.000	0.091
	(0.178)	(0.106)	(0.122)	(0.114)
Prime-age unemployment rate	0.044	0.066	-0.099	-0.037
	(0.200)	(0.232)	(0.152)	(0.184)
Country trends	NO	YES	NO	YES
Observations	1,498	1,498	1,668	1,668
R^2	0.752	0.753	0.882	0.883

***Note**: The table reports the correlation between the share component of the shift-share and push-factor distance-based variables and some education and economic indicators of the local labor markets under analysis. All regressions include year and region fixed-effects Standard errors, in parentheses, are clustered at the province level.*

*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table 4A7: Overidentification test on the overall effect.

	All		Men		Women	
	(1)	(2)	(3)	(4)	(5)	(6)
$\Delta m^{15-64}_{r,t-10}$	-0.932***	-0.918***	-0.613***	-0.706***	-1.268***	-1.118***
	(0.222)	(0.268)	(0.235)	(0.264)	(0.240)	(0.301)
Country trends	NO	YES	NO	YES	NO	YES
R^2	0.501	0.644	0.356	0.551	0.570	0.668
First-stage F-stat.	43.4	33.8	43.4	33.3	45.4	36.4
Overidentification test						
Hansen's J-stat.	1.897	1.009	2.702	0.599	0.938	0.996
p-value	0.169	0.315	0.100	0.439	0.333	0.318

***Note**: The table reports the estimate of the overall effect obtained with the two instruments, as well as the overidentification test. All the RHS variables are lagged one census. All regressions include controls and time fixed-effects and are weighted by $1/(1/w_t + 1/w_{t-10})$. In all specifications the number of observations is 258. Standard errors, in parentheses, are clustered at the province level.*

*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table 4A8: Summary statistics.

Variable	Mean	S.D.	Min.	Max.
Change in composition-adjusted regional schooling variable				
All	0.146	0.069	-0.049	0.301
Men	0.125	.0068	-0.064	0.309
Women	0.166	0.082	-0.037	0.364
Change in immigration related regressors				
Share of immigrants aged 15-64	0.022	0.029	-0.003	0.125
Share of immigrants aged 15-18	0.014	0.032	-0.050	0.143
Share of immigrants aged 19-64 with:				
Less than upper-secondary education	0.009	0.017	-0.027	0.074
Equal or more than upper-secondary education	0.010	0.014	-0.003	0.064
Change in controls				
Youth unemployment rate (19-24)	0.008	0.065	-0.150	0.172
Prime-age unemployment rate (25-54)	0.020	0.028	-0.056	0.117
Native cohort size	-0.018	0.024	-0.108	0.076
Share of low-educated workers in low-skilled sectors	-0.061	0.046	-0.225	0.049

Note: *Mean, S.D., minimum and maximum of the changes in the main variables used in the analysis.* ***Source***: *IPUMS.*

Table 4A9: Robustness checks, change immigrant measure (share of immigrants age 03-64).

	enclave-based		push-distance	
	(1)	(2)	(3)	(4)
Panel A: Whole population				
$\Delta m^{03-64}_{r,t-10}$	-1.114***	-0.815***	-0.678*	-1.189***
	(0.230)	(0.256)	(0.364)	(0.423)
Country trends	NO	YES	NO	YES
R^2	0.481	0.660	0.513	0.621
First-Stage F-stat.	65.5	36.4	24.4	42.9
Panel B: Men				
$\Delta m^{03-64}_{r,t-10}$	-0.823***	-0.639**	-0.251	-0.908**
	(0.252)	(0.262)	(0.368)	(0.385)
Country trends	NO	YES	NO	YES
R^2	0.335	0.565	0.369	0.537
First-Stage F-stat.	65.9	36.6	22.3	41.8
Panel C: Women				
$\Delta m^{03-64}_{r,t-10}$	-1.414***	-0.989***	-1.102***	-1.427***
	(0.236)	(0.276)	(0.407)	(0.496)
Country trends	NO	YES	NO	YES
R^2	0.555	0.681	0.576	0.642
First-Stage F-stat.	70.1	38.8	28	46.2

Note: *The table reports the 2SLS estimates of the robustness check on the overall effect, where I change the measure of the immigrant supply shock as indicated in the text. All regressions include controls and year fixed-effects. In all specifications the number of observations is 258. Standard errors, in parentheses, are clustered at the province level.*
*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table 4A10: Robustness checks, using alternative weighting scheme.

	enclave-based				push-distance			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Panel A: Whole population								
$\Delta m^{15-64}_{r,t-10}$	-1.042***	-0.742***	-0.899***	-0.897	-0.779*	-1.324***	0.359	-0.378
	(0.234)	(0.273)	(0.308)	(0.752)	(0.410)	(0.508)	(0.533)	(0.575)
Country trends	NO	YES	NO	YES	NO	YES	NO	YES
Weights	Pop	Pop	NO	NO	Pop	Pop	NO	NO
R^2	0.553	0.709	0.389	0.578	0.563	0.654	0.394	0.631
First-Stage F-stat.	56.5	26.2	14.6	4.9	19.2	31.4	11.4	21
Panel B: Men								
$\Delta m^{15-64}_{r,t-10}$	-0.711***	-0.597**	-0.471	-0.918	-0.318	-1.018**	0.845	-0.017
	(0.261)	(0.291)	(0.340)	(0.944)	(0.404)	(0.443)	(0.618)	(0.653)
Country trends	NO	YES	NO	YES	NO	YES	NO	YES
Weights	Pop	Pop	NO	NO	Pop	Pop	NO	NO
R^2	0.428	0.634	0.251	0.409	0.435	0.592	0.207	0.491
First-Stage F-stat.	56.5	26.2	14.6	4.9	19.2	31.4	11.4	21
Panel C: Women								
$\Delta m^{15-64}_{r,t-10}$	-1.386***	-0.902**	-1.349***	-0.898	-1.263***	-1.638***	-0.206	-0.797
	(0.243)	(0.289)	(0.349)	(0.668)	(0.478)	(0.624)	(0.542)	(0.611)
Country trends	NO	YES	NO	YES	NO	YES	NO	YES
Weights	Pop	Pop	NO	NO	Pop	Pop	NO	NO
R^2	0.604	0.717	0.457	0.627	0.610	0.658	0.497	0.636
First-Stage F-stat.	56.5	26.2	14.6	4.9	19.2	31.4	11.4	21

***Note**: The table reports the 2SLS estimates of the robustness check on the overall effect, where I change the weighting scheme. All regressions include controls and year fixed-effects. In all specifications the number of observations is 258. Standard errors, in parentheses, are clustered at the province level.*

*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table 4A11: Robustness checks, control for labor demand shocks.

	enclave-based		push-distance	
	(1)	(2)	(3)	(4)
Panel A: Whole population				
$\Delta m_{r,t-10}^{15-64}$	-0.833***	-0.560**	-0.334	-0.786*
	(0.217)	(0.277)	(0.380)	(0.424)
Bartik shock	-0.215**	-0.199**	-0.222**	-0.189**
	(0.094)	(0.084)	(0.095)	(0.093)
Country trends	NO	YES	NO	YES
R^2	0.532	0.692	0.543	0.674
First-Stage F-stat	67.5	33.1	36.2	42.1
Panel B: Men				
$\Delta m_{r,t-10}^{15-64}$	-0.551**	-0.482*	0.020	-0.656
	(0.231)	(0.290)	(0.384)	(0.417)
Bartik shock	-0.174*	-0.121	-0.183*	-0.113
	(0.094)	(0.078)	(0.097)	(0.086)
Country trends	NO	YES	NO	YES
R^2	0.377	0.579	0.383	0.563
First-Stage F-stat	68	33.8	33.6	40.7
Panel C: Women				
$\Delta m_{r,t-10}^{15-64}$	-1.128***	-0.614**	-0.680	-0.845*
	(0.240)	(0.297)	(0.419)	(0.465)
Bartik shock	-0.260**	-0.289***	-0.266**	-0.279***
	(0.107)	(0.098)	(0.105)	(0.106)
Country trends	NO	YES	NO	YES
R^2	0.604	0.727	0.616	0.715
First-Stage F-stat	71.3	34.4	41.9	45.1

***Note**: The table reports the 2SLS estimates of the robustness checks on the overall effect, where I control for labor demand shock by means of a Bartik-like variable. All regressions include the other controls as defined in section 4.4 and year fixed-effects. In all specifications the number of observations is 258. Standard errors, in parentheses, are clustered at the province level.*
*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table 4A12: Robustness checks, control for immigrants' and natives' average ages.

	enclave-based		push-distance	
	(1)	(2)	(3)	(4)
Panel A: Whole population				
$\Delta m_{r,t-10}^{15-64}$	-0.964***	-0.739***	-0.634*	-1.107***
	(0.215)	(0.258)	(0.355)	(0.393)
ΔImm. avg age	-0.005	-0.010*	-0.000	-0.002**
	(0.006)	(0.005)	(0.001)	(0.001)
ΔNat. avg age	-0.001	-0.001*	-0.003	-0.011**
	(0.001)	(0.001)	(0.007)	(0.005)
Country trends	NO	YES	NO	YES
R^2	0.501	0.670	0.518	0.635
First-Stage F-stat	70.2	39.9	35.8	50.1
Panel B: Men				
$\Delta m_{r,t-10}^{15-64}$	-0.647***	-0.549**	-0.227	-0.836**
	(0.230)	(0.258)	(0.360)	(0.356)
ΔImm. avg age	-0.006	-0.012**	-0.001	-0.002**
	(0.008)	(0.006)	(0.001)	(0.001)
ΔNat. avg age	-0.001	-0.002**	-0.005	-0.013**
	(0.001)	(0.001)	(0.008)	(0.006)
Country trends	NO	YES	NO	YES
R^2	0.358	0.582	0.372	0.557
First-Stage F-stat	70.5	40.8	33.4	48.6
Panel C: Women				
$\Delta m_{r,t-10}^{15-64}$	-1.300***	-0.932***	-1.042***	-1.343***
	(0.224)	(0.280)	(0.393)	(0.459)
ΔImm. avg age	-0.003	-0.010	-0.000	-0.001
	(0.006)	(0.006)	(0.001)	(0.001)
ΔNat. avg age	-0.000	-0.001	-0.002	-0.010*
	(0.001)	(0.001)	(0.006)	(0.006)
Country trends	NO	YES	NO	YES
R^2	0.568	0.686	0.583	0.653
First-Stage F-stat	72.8	64.7	40.7	53.5

***Note**: The table reports the 2SLS estimates of the robustness checks on the overall effect, where I control for immigrants' and natives' average ages. All regressions include the other controls as defined in section **??** and year fixed-effects. In all specifications the number of observations is 258. Standard errors, in parentheses, are clustered at the province level.*

*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table 4A13: Robustness checks, control for population and male shares.

	enclave-based		push-distance	
	(1)	(2)	(3)	(4)
Panel A: Whole population				
$\Delta m^{15-64}_{r,t-10}$	-0.979***	-0.839***	-0.629*	-1.132***
	(0.217)	(0.265)	(0.350)	(0.403)
ΔPop. share	-0.106	-0.162	-0.168	-0.181
	(0.246)	(0.235)	(0.261)	(0.244)
ΔMale share	0.435	0.580	0.475	0.638
	(0.572)	(0.487)	(0.571)	(0.499)
Country trends	NO	YES	NO	YES
R^2	0.498	0.653	0.519	0.621
First-Stage F-stat	61.9	34.9	32.7	47.8
Panel B: Men				
$\Delta m^{15-64}_{r,t-10}$	-0.668***	-0.634**	-0.232	-0.853**
	(0.233)	(0.269)	(0.363)	(0.366)
ΔPop. share	0.145	-0.010	0.067	-0.024
	(0.270)	(0.233)	(0.287)	(0.239)
ΔMale share	0.288	0.565	0.330	0.607
	(0.864)	(0.733)	(0.861)	(0.738)
Country trends	NO	YES	NO	YES
R^2	0.352	0.559	0.370	0.535
First-Stage F-stat	62	35.6	30.3	46.7
Panel C: Women				
$\Delta m^{15-64}_{r,t-10}$	-1.303***	-1.039***	-1.015***	-1.364***
	(0.228)	(0.287)	(0.385)	(0.471)
ΔPop. share	-0.384	-0.291	-0.437	-0.312
	(0.254)	(0.278)	(0.269)	(0.289)
ΔMale share	0.581	0.558	0.622*	0.617
	(0.356)	(0.350)	(0.359)	(0.376)
Country trends	NO	YES	NO	YES
R^2	0.572	0.676	0.590	0.647
First-Stage F-stat	36.6	41.6	37.3	51.1

***Note**: The table reports the 2SLS estimates of the robustness checks on the overall effect, where I control for population and male shares. All regressions include the other controls as defined in the main text and year fixed-effects. In all specifications the number of observations is 258. Standard errors, in parentheses, are clustered at the province level.*

*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table 4A14: Robustness checks, control for local employment structure.

	enclave-based		push-distance	
	(1)	(2)	(3)	(4)
Panel A: Whole population				
$\Delta m_{r,t-10}^{15-64}$	-1.296***	-0.898***	-1.142	-2.193***
	(0.329)	(0.377)	(0.763)	(0.829)
Δ*Employment shares in:*				
Agriculture	-0.228**	-0.206**	-0.233*	-0.137
	(0.115)	(0.090)	(0.122)	(0.124)
Construction	-0.005	0.228	0.013	0.182
	(0.267)	(0.196)	(0.276)	(0.262)
Retail services	-0.972***	-0.691***	-0.882**	-1.284***
	(0.263)	(0.232)	(0.446)	(0.432)
Country trends	NO	YES	NO	YES
R^2	0.514	0.678	0.526	0.495
First-Stage F-stat	35.5	23.8	8.7	19.3
Panel B: Men				
$\Delta m_{r,t-10}^{15-64}$	-1.065***	-0.819**	-0.801	-2.033***
	(0.321)	(0.361)	(0.769)	(0.782)
Δ*Employment shares in:*				
Agriculture	-0.228**	-0.179**	-0.239**	-0.115
	(0.112)	(0.084)	(0.118)	(0.116)
Construction	-0.133	0.223	-0.101	0.172
	(0.259)	(0.201)	(0.262)	(0.264)
Retail services	-1.117***	-0.780***	-0.963**	-1.339***
	(0.263)	(0.226)	(0.468)	(0.425)
Country trends	NO	YES	NO	YES
R^2	0.377	0.579	0.395	0.392
First-Stage F-stat	36	24.2	8.1	19
Panel C: Women				
$\Delta m_{r,t-10}^{15-64}$	-1.555***	-0.993**	-1.478*	-2.317**
	(0.387)	(0.432)	(0.848)	(0.918)
Δ*Employment shares in:*				
Agriculture	-0.236*	-0.247**	-0.239*	-0.178
	(0.129)	(0.107)	(0.136)	(0.139)
Construction	0.196	0.276	0.204	0.247
	(0.305)	(0.228)	(0.320)	(0.288)
Retail services	-0.852***	-0.635**	-0.806	-1.249***
	(0.327)	(0.296)	(0.507)	(0.486)
Country trends	NO	YES	NO	YES
R^2	0.583	0.704	0.588	0.560
First-Stage F-stat	35.8	24.2	9.8	20.4

***Note**: The table reports the 2SLS estimates of the robustness checks on the overall effect, where I control for the employment structure of the local labor markets under analysis. All regressions include the other controls as defined in the main text and year fixed-effects. In all specifications the number of observations is 258. Standard errors, in parentheses, are clustered at the province level.*

*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table 4A15: Overall effect of immigration and specialization in low-skilled sectors.

	All		Men		Women	
	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: OLS estimates						
$\Delta m^{15-64}_{r,t-10}$	-0.841**	-0.226	-0.592	-0.038	-1.113***	-0.402
	(0.340)	(0.373)	(0.366)	(0.383)	(0.352)	(0.403)
$\Delta S^{low}_{r,t-10}$	0.233**	-0.031	0.150	-0.067	0.326***	-0.002
	(0.096)	(0.110)	(0.105)	(0.110)	(0.105)	(0.131)
Interaction	-5.883	-3.098	-6.326	-2.401	-5.673	-3.868
	(3.990)	(4.008)	(4.473)	(4.122)	(4.211)	(4.351)
Country trends	NO	YES	NO	YES	NO	YES
R^2	0.529	0.686	0.376	0.589	0.603	0.705
Panel B: 2SLS estimates with enclave-based instrument						
$\Delta m^{15-64}_{r,t-10}$	-1.543***	-0.886***	-1.451***	-0.784***	-1.627***	-0.992***
	(0.331)	(0.271)	(0.325)	(0.264)	(0.353)	(0.298)
$\Delta B^{low}_{r,t-10}$	-0.096	-0.209	-0.015	-0.128	-0.200	-0.307*
	(0.212)	(0.143)	(0.198)	(0.130)	(0.252)	(0.172)
Interaction	-14.436**	-7.106	-21.212***	-10.008*	-6.907	-4.493
	(5.899)	(5.395)	(5.888)	(5.579)	(6.404)	(5.939)
Country trends	NO	YES	NO	YES	NO	YES
R^2	0.556	0.702	0.445	0.614	0.593	0.715
First-stage F-stat.	16.3	7.6	16.6	7.9	17.4	8
Panel C: 2SLS estimates with push-factor & distance-based instrument						
$\Delta m^{15-64}_{r,t-10}$	-2.058***	-1.813***	-1.657***	-1.549***	-2.415***	-2.010***
	(0.561)	(0.562)	(0.555)	(0.428)	(0.644)	(0)
$\Delta B^{low}_{r,t-10}$	0.158	-0.083	0.131	-0.032	0.166	-0.159
	(0.207)	(0.171)	(0.207)	(0.152)	(0.236)	(0.206)
Interaction	-29.838***	-19.266**	-29.675***	-19.154***	-29.075***	-18.849*
	(8.732)	(8.719)	(9.034)	(7.129)	(9.664)	(10.974)
Country trends	NO	YES	NO	YES	NO	YES
R^2	0.513	0.653	0.421	0.571	0.546	0.675
First-stage F-stat	10.9	23.5	10.7	22.6	11.3	25.5

***Note**: The table reports the OLS (Panel A) and 2SLS (Panels B and C) estimates of the overall effect for regions specialized in low-skilled sectors, as indicated in equations (4.6) and (4.7). All the RHS variables are lagged one census. All regressions include controls and time fixed-effects and are weighted by $1/(1/w_t+1/w_{t-10})$. In all specifications the number of observations is 258. Standard errors, in parentheses, are clustered at the province level.*

*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table 4A16: Regional specialization in low-skilled sectors based on dummy variables.

	All		Men		Women	
	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: dummy-low=1 if $\Delta B^{low}_{r,t-10} \geq 0.31$ (around 50^{th} percentile)						
Interaction	-4.460**	-3.943***	-3.277*	-2.677**	-5.699**	-5.341***
	(2.000)	(1.242)	(1.883)	(1.116)	(2.254)	(1.582)
Country trends	NO	YES	NO	YES	NO	YES
R^2	0.572	0.632	0.462	0.551	0.600	0.643
First-stage F-stat	3.1	2.9	3.2	2.9	3.1	2.9
Panel B: dummy-low=1 if $\Delta B^{low}_{r,t-10} \geq 0.39$ (around 75^{th} percentile)						
Interaction	-6.908**	-3.507	-7.299***	-3.493	-7.035**	-4.143
	(3.083)	(2.410)	(2.835)	(2.267)	(3.594)	(2.919)
Country trends	NO	YES	NO	YES	NO	YES
R^2	0.576	0.653	0.439	0.555	0.625	0.677
First-stage F-stat	13.4	17.2	12.3	15.5	14.4	18.5
Panel C: dummy-low=1 if $\Delta B^{low}_{r,t-10} \geq 0.47$ (around 90^{th} percentile)						
Interaction	-11.224***	-5.647**	-10.300***	-4.460**	-12.524***	-7.197***
	(1.553)	(2.228)	(1.517)	(2.215)	(1.794)	(2.358)
Country trends	NO	YES	NO	YES	NO	YES
R^2	0.519	0.659	0.372	0.561	0.585	0.682
First-stage F-stat	33.3	17.1	33.5	17.5	35.1	17.8

***Note**: The table reports the results relative to the robustness checks where I define specialization in low-skilled sectors by means of a set of dummy variables based on the distribution of the Bartik variable for low-skilled sectors. In all specifications the number of observations is 258. Standard errors, in parentheses, are clustered at the province level.*

*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Chapter 5

Conclusions

In the recent decades, in many OECD economies, the phenomenon of immigration has experienced a stable and consistent increase. In particular, this has occurred not only in countries with a long tradition of immigration, as the U.K., Germany, o the U.S., but also in countries that, until relatively recently, were more characterized by the emigration of its population. This is the case of the countries analyzed in the different chapters of this thesis, namely Greece, Ireland, Italy, Portugal and Spain. An important characteristics of the worldwide immigrant population is that the vast majority of individuals are in working-age and may be considered as economic immigrants, as they move looking for better job opportunities. This implies that foreign-born individuals have become an increasingly important component of the labor force of the countries of destination. In this context, the issue of immigration has gained a central stage in the political debate of the host countries, as well as among scholars. Related to the former, the recent years have witnessed the rise of a significant number of political parties, typically with a right-wing orientation, that fiercely oppose immigration.

In the academic world, immigration is one of the more discussed and controversial topics and, despite the large amount of studies carried out, the consensus on what is the actual economic impact of immigration (if any) is far from being reached. In this context, the objective of this thesis is to provide empirical evidences on the impact of immigration on different native outcomes, namely employment, wages and human capital accumulation. The analysis is particularly focused on the Italian context, which, for several reasons, represents an interest case study. This is because, first of all, Italy has only recently changed its role from being an immigrant-rending country. to be an immigrant-receiving one. To this extent, because of its strategic position in the middle of the Mediterranean Sea, Italy is a fairly popular destination particularly for immigrants coming from Africa, as well as Central and Eastern Europe. Second, Italy was also one of the European countries mostly affected by the recent economic crisis, aspect that has created a resurgence of an anti-immigration sentiment in both the public opinion and the political class. Third, Italy is characterized by a strong degree of employment protection and (downward) wage rigidity, aspects that play a crucial role in the extent to which Italian local labor markets are able to absorb an immigration-induced supply shift. All these elements together make an analysis on the impact of immigration in the Italian context particularly interesting, and this is further magnified by the surprisingly scarce studies on the subject in the previous literature.

To this extent, the second and third chapters of the thesis analyze the impact that the recent inflows of immigrants exert on both the employment prospects and the wage structure of the native workers. More in detail, the second chapter analyze the impact of immigration on native employment in the Italian administrative provinces over the period 2009-2017. The results contradict the idea of immigrants being responsible of a job-loss among natives, and highlight an overall negligible native employment response. In addition, the heterogeneity analysis has allowed to identify some interesting peculiarities. First of all, when distinguishing native workers based on their completed education, the estimates have revealed a positive impact on high-educated workers, while no effect is estimated for low-educated ones. Second, the results obtained when distinguishing by occupation indicate a positive impact on skilled-manual workers and a negligible one on both white and blue collars. Third, interestingly, when distinguishing by gender, the estimates has identified a positive employment response for females, irrespective of their skills, while the employment prospects of men appear not to be significantly affected by immigrant shocks.

If the second chapter provides evidences on how immigration affects "quantities" (i.e. the number of people employed), the third one is instead relative to the impact of immigration on "prices" (that is, wages). More precisely, the third chapter investigates the impact of immigration along the native wage distribution in the Italian context and during the period 2009-2017. In line with the second chapter and with a strand of the existing literature, the results oppose the idea of immigration exerting a negative impact on native wages. Instead, they suggest an overall negligible effect on both average wages and on the entire wage distribution. Interestingly, the heterogeneity analyses identify a positive, albeit marginally significant, wage response that is particularly driven by high-skilled women (that is, located in the upper part of the conditional wage distribution) residing in the Northern provinces.

The fourth chapter conducts instead a complementary analysis. Specifically, overall this chapter studies whether, and to what extent, the presence of immigrants induces natives to invest in human capital. To study the native education responses to immigration is of extreme interest because it is an important mechanism of adjustment that the native population may enact to react to the inflows of immigrants into the labor market. To this extent, to invest in education may ensure natives to acquire the skills that are complementary, rather than substitute to those of the immigrants. This, in turn, may allow them to somewhat minimize immigrants competition once in the labor market. This chapter contributes to the existing literature by providing evidences on a set of European countries, namely Greece, Ireland, Portugal and Spain, over the period 1981-2011. The rational behind the choice of the countries is that, on the one hand, they have some some specificities that render them particularly different from the countries typically studied in the existing literature on immigration (the U.K., the U.S. or Germany), and on the other, make them similar to Italy (which is the country analyzed in the two previous chapters). The first peculiarity is that they only recently became immigrant-recipient countries. The second is that they were particularly affected by the Global Financial Crisis. Similarly to what happened in Italy, the combination of these two aspects together induced the appearance of an anti-immigration sentiment. The third is that all the countries considered in the fourth chapter are characterized by fairly high degree of rigidity in both product and labor markets.

From a methodological point of view, the chapter is divided into two main parts. In the first, I assess the impact of immigration on the propensity of natives to acquire human capital. To do so, I estimate both the direct and indirect effects[1] of immigration on native

[1] The first is the effect that immigrants in schooling-age exert on the school performances of their native counterparts. The second is instead the effect that immigrants in working-age exert on the native decisions

schooling, as well as the overall effect that comprises the other two. The results of this first part, which survive to a set of robustness tests, identify a negative native education response to immigrant shocks that is somewhat stronger for native women. Then, in the second part I consider what is the effect that the combination between the immigrant penetration and the local specialization in low-skilled sectors exerts on the native propensity to invest in education. In line with the hypothesis, the results indicate that the native education response to immigration is particularly negative in regions specialized in low-skilled type of activities.

In the three core chapters of the thesis, I address the non-random immigrant settlements by means of an instrumental variable approach using different types of instruments. Specifically, the first is the canonical shift-share (or enclave-based) variable, which combines the spatial variation in the immigrants historical settlement by countries of origin with the national-level growth rate of the immigrant population, again distinguishing by countries of origin. Despite the numerous criticisms that this type of variable has recently received, I show its validity in the context of the countries under analysis. Second, I also make use of an alternative type of instrument, that is based, rather than on immigrants historical settlements, on the combination of the distance between the countries of origin and the region of destination, and the push-factors of the countries of origin.

The empirical models sketched in all these chapters have a clear spatial focus. This is because this empirical choice allows to exploit the uneven geographical distribution of the immigrant population, as well as the spatial heterogeneity in economic performance that characterizes the regions of the countries considered. Moreover, the main limitation of the spatial approach, that is the native relocation responses to immigrant shocks, are extremely limited (if not absent) in the countries and period analyzed. Finally, this type of approach allows to identify the *overall* effect of immigration, and therefore, the parameter that it estimates is more policy-relevant (Dustmann et al., 2016).

To this extent, the results obtained in the chapters of this thesis have important policy implications, particularly those of the second and third chapters. Specifically, in terms of the impact of immigration on native labor market outcomes, the second and third chapter indicate that (i) overall, immigration does not exert a significant impact on both native employment and wages (as instead claimed by some political actors); and (ii) there is evidence of a positive impact on both employment and wages of native females, particularly for the medium/highly skilled ones[2]. This last aspect is particularly relevant from a policy standpoint, especially given the still substantial gender disparities in economic performances and opportunities that characterize the Italian peninsula. Also the results of the fourth chapter may have important implications. Specifically, on the one hand they highlight the interest in conducting an analysis on the incentives to accumulate education in the context of the specific EU countries considered, that are fairly different in many aspects to the countries typically analyzed in the literature. On the other, they underline the relevance of considering the local specialization in low-skilled activities when analyzing th native education responses to immigration.

To conclude, it is also important to acknowledge some shortcomings that characterize the empirical exercises of the chapters of this thesis. Specifically, due to data limitations, the analyses of the second and third chapters focus on the short-run employment and wage

to invest in human capital.

[2]In tis regard, it is important to highlight that the positive impact on the wages of native females is only marginally significant, but it is still present.

effects of immigration. Conversely, the long-run responses, that may also involve different magnitudes (e.g. innovation, productivity, etc), are inevitably left aside. As for the fourth chapter, I consider the specialization in low-skilled sectors. This is because, these represent an important component of the total employment, particularly given the economies and period under analysis. However, it could be interesting to also consider the specialization in high-skilled (or technology-intensive) type of activities (although, perhaps, to obtain interesting results, the analysis should be performed considering other countries and a different period of time). In any case, I will try to address all these shortcomings in future research.

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