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# The effects of immigration in frictional labor markets: Theory and empirical evidence from EU countries

Eva Moreno-Galbis [a,b,c,d], Ahmed Tritah [d,*]

[a] GRANEM, Université d'Angers, Avenue Francois Mitterrand, 49000 Angers cedex, France
[b] CREST, France
[c] IRES, (Catholic University of Louvain)
[d] GAINS-TEPP, Université du Maine, Avenue Olivier Messiean, 72085 Le Mans cedex, France

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ABSTRACT

Immigrants are newcomers in a labor market. As a consequence, they lack host-country-specific labor market knowledge and other country-specific and not directly productive valuable assets affecting their relative bargaining position with employers. We introduce this simple observation into a search and matching model of the labor market and show that immigrants increase the employment prospects of competing natives. To test the predictions of our model, we exploit yearly variations between 1998 and 2004 in the share of immigrants within occupations in 13 European countries. We identify the impact of immigrants on natives' employment rate using an instrumental variable strategy based on historical settlement patterns across host countries and occupations by origin country. We find that natives' employment rate increases in occupations and sectors receiving more immigrants. Moreover, we show that this effect varies depending on immigrants' characteristics and on host country labor market institutions which affect relative reservation wages.



## 1. Introduction

The consequences of immigration on labor market outcomes and host country welfare have been extensively discussed in the economic literature, both theoretically and empirically. This paper analyzes the impact of an immigration-induced labor supply shock on the employment opportunities of native European workers. Theoretically, the labor market consequences of immigration have been framed within a standard neoclassical labor supply, labor demand framework (see Borjas (2003); Card (2001, 2005, 2009), and Ottaviano and Peri (2012)). In such a framework, in the short run, a labor supply shock generated by the arrival of immigrants triggers a reduction in the wages of competing natives, which in turn may discourage labor force participation. As a consequence, the crucial problem in this literature is determining against which natives immigrants are competing, and then, analyzing the distributional consequences of an immigration inflow (see for example Friedberg and Hunt (1995)). Yet, this framework has somewhat been challenged by empirical findings over the last two decades. Exploiting various experiences of immigration, in the US first, and more recently in Europe, the literature has failed to find a consistent negative impact of immigrants on natives' labor market outcomes.[1]

* Corresponding author.
*E-mail addresses:* Eva.morenogalbis@univ-angers.fr (E. Moreno-Galbis), Ahmed.Tritah@univ-lemans.fr (A. Tritah).

[1] Note that an overwhelming majority of this literature has been focused on the impact over less-skilled natives, with the experience of the US following the 1965 Immigration Act that shifted immigrants' composition towards poorer countries and notably Mexicans; Card (1990), Altonji et al. (1991),

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A more recent literature has put some explanations forward. On the one hand, one stream argues that natives and immigrants are never perfectly substitutable. Notably, according to Ottaviano and Peri (2012), immigrants and natives, in spite of having similar "observable" skills, are not perfectly substitutable in production. According to their estimates, newly arrived immigrants are substitutes for older immigrants whereas they are imperfect substitutes for natives. Peri and Sparber (2011, 2009) or D'Amuri and Peri (2014) justify this finding by considering different relative skill endowments between natives and immigrants. Whereas natives have a comparative advantage in communication skill intensive jobs, immigrants have a comparative advantage in manual skill intensive jobs. A complementary relation arises then between both types of workers. Following an immigration-induced labor supply shock, natives reallocate towards communication and language intensive tasks, while immigrants become specialized in manual intensive tasks. A second stream of literature analyzes how immigration may affect the labor market decisions of natives with different skills with respect to immigrants through general-equilibrium effects. For the US data, Cortes and Tessada (2011) find that the reduction in the price of services (being close substitutes for household production) stemming from recent waves of low-skilled immigration has provided incentives for high-skilled women (earning above the median of the wage distribution) to substitute their own time invested in the production of household goods with work hours on the labor market. A similar study using Spanish data is proposed in Farré et al. (2011). They find that over the last decade immigration has led to the significant expansion of the household service sector and to an increase in the labor supply of women in high-earning occupations. A third stream of literature focuses rather on the endogenous nature of technological change. Lewis (2011) looks at labor demand side adjustment, and shows that firms adjust to unskilled labor supply shocks by adopting less skilled-biased technology.

We propose an alternative factor explaining the absence of a negative impact on natives' labor market outcomes: whatever the labor market considered, immigrants are newcomers. As a consequence, they lack host-country-specific labor market knowledge and other, although non directly productive, valuable assets. For instance, one such asset is the eligibility and amount of unemployment benefits which are conditional on past employment experience in host countries. These characteristics affect immigrants' outside options and put them in a weaker bargaining position as compared to natives when they negotiate their wages with employers, making them more profitable workers. Following an inflow of immigrants, the average expected profit of firms operating in the receiving labor market increases, raising incentives to open more vacancies. We claim thus that, even if they are perfectly substitutable with natives in the production process, immigrants are at the origin of a positive externality.

Another important difference with our paper is that most papers are focused on relatively longer-term effects, typically using time series of decennial censuses. These long-term effects may not be the same as short-term effects. Indeed, one prediction of our model for which we find some empirical support is that long-term stayers do no exert any effect on natives: they affect the size of the economy without any impact on its structure.

The idea of a divergent reservation wage between natives and immigrants, leading immigrants to accept lower wages, seems to be widely supported by the empirical evidence on wage inequality between natives and immigrants (see Algan (2010); Kee (1995) or Card (2005)). Moreover, using UK data, Nanos and Schluter (2014) estimate that, when controlling for the divergence in the reservation wage between natives and immigrants, the migrant effect of the wage differential is reduced by almost 55%. Gonzalez and Ortega (2011) also suggest the role of the divergent reservation wage as a determinant of their empirical finding: a 10% increase in labor supply in the skill-region cell is associated with a change in the cell employment rate of 0.884, which is indeed not that far from our own estimates.

Our paper proposes a search and matching framework à la Pissarides (1990) to analyze the labor market impact of immigrants. In labor markets with search and matching frictions and relatively generous institutions, disparities in outside options (reservation wage) may account for a substantial share of wage differentials between (eligible and protected) natives and (non-eligible and unprotected) immigrants. Interestingly, immigrants' wage gap with respect to similar natives seems to be more important in European countries than in the US (see Card, 2005 and Algan, 2010). In the former, labor market institutions may play a specific role in increasing the relative reservation wage of natives, especially since afterwards, with years of residence in the host country (when immigrants become eligible), wage differentials (for similar characteristics) tend to disappear (see Chiswick, 1978; Borjas, 1994 or Borjas, 1999 for the US, Chiswick et al., 2005 for Australia, Friedberg and Hunt, 1995 for Israel or Lam and Liu, 2002 for Hong Kong).

With the notable exceptions of Ortega (2000), Chassamboulli and Palivos (2014), Chassamboulli and Peri (2014) or Liu (2010), we are not aware of any other study analyzing the labor market impact of immigrants on host countries using a search and matching model of equilibrium unemployment. Ortega (2000) is interested in the equilibrium distribution of workers in the host and origin countries and the employment consequences for natives in the host country. Ortega (2000) shows that, provided they have higher search costs, immigrants can improve the employment prospects of natives. Chassamboulli and Palivos (2014) also consider that immigrants have a lower outside option with respect to natives, and distinguish between skilled and unskilled immigration to analyze the consequence of skill-biased immigration in the US between 2000 and 2009. Chassamboulli and Peri (2014) employ a theoretical setup to show that legalization stimulates firms' job creations by increasing the number of workers with a low reservation wage. In all three papers, the lower outside

(footnote continued)
Card and DiNardo (2000), Card (2001), and Borjas (2003) are the most influential papers. On Europe, see Dustmann et al. (2013) for the UK, Glitz (2011) for Germany, Gonzalez and Ortega (2011) for Spain, and Ortega and Verdugo (2014) for France. Longhi et al. (2006) offer a summary and perform a meta-analysis on the wage effect of immigrants. For a literature review, see Borjas (1999) or the more recent work by De la Rica et al. (2013).

option of immigrants is the source of the positive externality exerted by immigrants on natives and provides the main explanation for the absence of a short-run negative impact of immigrants on native employment.[2] However, no empirical evidence is provided. In our paper, we try to fill this gap.

Using data from the European Labor Force surveys from 1998 to 2004, we estimate the elasticity of natives' employment rate to changes in the share of immigrants within a labor market defined at the occupational and country level (every year, for every country, we define 9 broad occupations).[3] Labor supply shocks are analyzed within occupations to ensure that immigrants and natives are close substitutes as postulated in our model and that mobility across labor markets is limited. We deal with the endogeneity issue raised by the non-random sorting of immigrants across country-occupation labor market by adopting an instrumental variable strategy. Our instrument, inspired from Altonji et al. (1991), is based on immigrants' past historical settlement patterns by origin country across occupations and host countries.[4] We find that immigrants exert a small but positive impact on male natives' employment rate. In our preferred specification, we find that an increase in the number of immigrants by 10% in a country-occupation labor market is associated with an increase in natives' employment rate in that occupation by around 0.5%. Although small, the potential employment gain may be substantial in light of the large increase in the number of immigrants experienced in EU countries over the period considered.

Next, based on a two-sector extension of our theoretical framework, we find that immigrants lead to natives within a labor market reallocating towards sectors receiving more immigrants. This crowding-in displacement effect points to employment creation for natives in the immigrant-receiving sector within the occupation.

In a third step, we exploit heterogeneity in terms of reservation wage across immigrants and across host countries, to provide evidence in support of our key mechanism. We find that the positive effect on native employment is all due to non-EU15 and recently arrived immigrants and is found to be larger in host countries where the gap in the unemployment benefit take-up rate between similar immigrants and natives is higher.

Lastly, using standard parameter values, we simulate the model and find that for almost all values of immigrants' outside option lower than that of natives, the model is able to replicate the elasticity of natives' employment rate to an immigration-driven labor supply shock. The numerically estimated elasticities are not statistically different from the empirically estimated elasticities.

The rest of the paper is organized as follows. The next section presents our basic one-sector matching model that allows us to provide a rationale for our empirical results. Section 3 discusses the data and gives some relevant descriptive statistics. Sections 4 and 5 present the empirical specification, identification strategy and estimation results, considering two alternative definitions of the labor market: country-occupation *vs.* country-occupation-sector. The relevance of our mechanism (divergent outside option) is assessed in Section 6. The numerical simulations are presented in Section 7. Section 8 concludes. We provide additional empirical results and an extension to a two-sector model in an Appendix.

## 2. The model

We employ a simple theoretical framework as a guideline for our empirical investigations. We present a particular case of Ortega (2000)'s framework with exogenous migration and rigid wages.[5] For the sake of generality, we consider a labor market composed of two types of labor suppliers, immigrants and natives who, as in Ortega (2000), are perfect substitutes. Immigrants and natives differ with respect to their outside opportunities of employment. For instance, immigrants arriving in a host country are likely to be non-eligible to unemployment benefits, they are likely to have a lower value of domestic production or leisure than natives, and they certainly lack other valuable assets. As a result, when considering the immigrant population as a whole, their average outside opportunity of employment is lower than that of natives.

Due to their lower outside opportunity of employment, immigrants accept a lower wage from firms. Immigrants are thus a more profitable group of workers from the firm's point of view, since they are as productive as natives but they are willing to work for lower wages. When the proportion of immigrants in the active population increases, firms' average expected profits increase, promoting the opening of new vacancies.[6] This positive externality could still arise even if immigrants had a relatively lower (or even higher) productivity. As long as the difference between the value of productivity and wages is greater for immigrants than for natives, immigrants remain more profitable workers and are thus the source of a positive externality. For simplicity, we consider here that both types of workers are identical apart from the reservation wage.

[2] These papers also introduce other sources of heterogeneity.

[3] To our knowledge, Angrist and Kugler (2003) and more recently D'Amuri and Peri (2014) are the only studies that exploit variations across European countries to identify the impact of immigrants on natives' employment. However, no study has looked at the impact at the occupational level.

[4] Such an instrument has proven to be a strong determinant of contemporaneous inflows in the single-country case (see Card, 2001, 2009; Patel and Vella, 2007 or Cortes and Tessada, 2011 for the US and Gonzalez and Ortega, 2011 or Farré et al., 2011 for Spain, among others) and in a multi-country setting (see D'Amuri and Peri, 2014).

[5] Most authors and notably Chassamboulli and Palivos (2014) or Chassamboulli and Peri (2014) calibrate their model to simulate the effect of observed immigration on wages of different skill groups of natives. Instead, our aim is to identify the impact of immigrants on natives' employment rate using reduced form estimates guided by our theoretical framework.

[6] Ortega (2000), Chassamboulli and Palivos (2014) or Chassamboulli and Peri (2014) have also emphasized disparities in search costs as the source of a positive externality of immigrants on native employment.

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### 2.1. *The matching process*

We use the subscript $j = N, I$ when referring to a native or an immigrant (foreign-born) worker. The workforce $P$ is such that $P = P_I + P_N$. Natives and immigrants may be employed ($n_j$) or unemployed ($u_j$), and the number of vacancies is denoted by $v$. The matching function can be written as: $M = m(v, u_N + u_I)$. We assume that a standard Cobb–Douglas matching function of the form $M = m_0(v)^{1/2}(u_N + u_I)^{1/2}$. Labor market opportunities are described by the market tightness variable $\theta = v/(u_N + u_I)$. The probability for a firm to fill an empty vacancy equals $q(\theta) = M/v$. The probability of finding a job for an unemployed worker is given by $p(\theta) = M/(u_N + u_I)$.[7]

### 2.2. *The agents' behavior*

#### 2.2.1. *Workers*

Employed workers are paid a wage $w_j$. Jobs are destroyed at the exogenous probability $s$. For $j = N, I$, the asset value of employment for natives and immigrants is given by $rE_j = w_j + s(U_j - E_j)$, where r stands for the interest rate, $rU_j = b_j + p(\theta)(E_j - U_j)$ for the asset value of unemployment and $b_j$ for the outside option of employment. Note that $b_N > b_I$.

#### 2.2.2. *Firms*

From the firm's point of view, the asset value associated with an empty vacancy is given by minus the cost of posting this vacancy, $\gamma$, plus the surplus obtained by the firm if it manages to fill the vacancy. The firm can only observe the worker's type at the time of the match and cannot discriminate between unemployed natives or unemployed immigrants. Firms thus cannot select their applicants. The possibility of rejecting an applicant that provides a positive surplus is not considered here. Actually, it is optimal for firms to fill the vacancy as long as the surplus associated with the match is positive rather than leaving it unfilled and bearing a per period cost $\gamma$ while waiting for a better worker to come in.[8] The value of an empty vacancy is given by

$$rV = -\gamma + q(\theta)(\bar{J} - V) \tag{1}$$

where $\bar{J}$ represents the expected value of a filled vacancy. The value of a filled vacancy is defined by the instantaneous profit $h - w_j$ associated with the job (productivity minus the wage), plus the expected loss if the vacancy becomes empty due to an exogenous job destruction shock:

$$rJ_j = h - w_j + s(V - J_j) \quad \text{j} = \text{natives, immigrants} \tag{2}$$

The expected value of a filled vacancy equals a weighted average $\bar{J} = \omega_1 J_I + (1 - \omega_1) J_N$, where $\omega_1 = \frac{u_I}{(u_N + u_I)}$. Firms open vacancies until no more profit can be obtained so that, at equilibrium, the free-entry condition $V=0$ applies, i.e.:

$$\frac{\gamma}{q(\theta)} = \bar{J} = \frac{h - \omega_1 w_I - (1 - \omega_1) w_N}{r + s} \tag{3}$$

The cost born by the firm while the vacancy remains empty must equal the expected value of a filled vacancy.

### 2.3. *Wages*

As in recent works by Pissarides and Vallanti (2007), Mortensen and Nagypal (2007) and Nagypal (2007), we adopt the rigid wage definition proposed by Hall and Milgrom (2008). We suppose that the worker receives a payoff $b_j$ in case negotiations fail, but also when the agreement is delayed. For the firm, we assume that there is no cost while the bargaining continues. Firms and workers renegotiate the division of the match product $h$, so that the outcome of the symmetric alternating-offers game is simply

$$w_N = \eta h + (1 - \eta) b_N \quad \text{and} \quad w_I = \eta h + (1 - \eta) b_I \tag{4}$$

where $\eta$ can be interpreted as the bargaining power of each party.

Note that $h > b_N > b_I$ (otherwise workers will prefer to remain unemployed rather than to accept a job), which implies that $w_N > w_I$.

[7] The hypothesis of equal probability of finding a job is consistent with the estimates of Datta-Gupta and Kromann (2014) on Danish data.

[8] In the standard one job-one firm framework, if firms are allowed to direct their vacancies towards a particular group, that is, if discrimination is possible, two independent labor market segments arise. At equilibrium, the free-entry condition must be respected in each labor market segment, *i.e.* $V_j = 0$ so that $J_j = \frac{y - w_j}{r + s}$, and the expected profit of posting a vacancy must be equalized across both segments, i.e. $q(\theta_I)\frac{y - w_I}{r+s} = q(\theta_N)\frac{y - w_N}{r+s}$. Because $\frac{y - w_I}{r+s} > \frac{y - w_N}{r+s}$, at equilibrium $\theta_I > \theta_N > 0$. In this context, the arrival of an immigrant wave into the immigrant segment will promote a reallocation of vacancies towards the immigrant segment (where the probability of filling a vacancy increases with the immigrant wave). In the presence of discrimination, the arrival of an immigrant wave is no longer the source of a positive externality on native employment.

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### 2.4. Employment opportunities

Employment opportunities are measured by the labor market tightness which is determined by the free-entry condition $\frac{\gamma}{q(\theta)} = \bar{J} = \frac{h-\overline{w}}{r+s}$, where $\overline{w} = \omega_1 w_I + (1-\omega_1) w_N$. Since $\frac{\gamma}{q(\theta)} = \frac{\gamma}{m_0}(\theta)^{1/2}$, we find

$$\theta = \left(\frac{m_0(h-\overline{w})}{\gamma(r+s)}\right)^2 \tag{5}$$

The larger the share of immigrants among workers, the lower the average wage paid by firms, which increases the expected profit of a filled vacancy. The consecutive boost in firms' labor demand raises natives' probability of finding a job. Natives' employment rate is thus increasing with the share of immigrants.[9]

### 2.5. Natives' steady-state employment rate

At the steady state, inflows and outflows from the labor market must be equalized, so that the number of unemployed and employed natives and immigrants is constant at equilibrium. Without loss of generality, the total population is normalized to 1 so that $P = P_I + P_N = 1$. Exits from employment equal the proportion of employed people losing their job, $s \cdot n_j$. Entries to employment correspond to the proportion of unemployed workers finding a job, $p(\theta)u_j = p(\theta)(P_j - n_j)$. Equalizing both, we obtain the labor market employment rate for group $j$:

$$s \cdot n_j = p(\theta)\left(P_j - n_j\right) \Rightarrow \frac{n_j}{P_j} = \frac{p(\theta)}{s+p(\theta)} \tag{6}$$

$\frac{n_j}{P_j}$ is increasing in the probability of finding a job, $p(\theta)$, which is increasing with the share of immigrants in the labor market (since the labor market tightness, $\theta$, increases with the share of immigrants). Because the total population is normalized to 1, the unemployment rate equals $\frac{u_j}{P_j} = 1 - \frac{u_j}{P_j} = \frac{s}{s+p(\theta)}$.

### 2.6. Testing the assumptions and predictions of the model

From this basic theoretical model, we deduce that the arrival of an immigrant wave exerts a positive externality since immigrants are more profitable workers, which pushes firms to open more vacancies. An increase in the share of immigrants in a particular labor market where immigrants and natives are substitutes, should improve the employment prospects of natives in that market. We design a strategy to empirically test the following set of results derived from the model.

First, we focus on the country-occupation labor market level, where occupations are defined at the one-digit level. Unlike age-education cells used in the literature, by defining skills at the occupation level, we make sure that immigrants and natives are close substitutes in the tasks performed.[10] According to our theoretical model, native employment should increase in occupations that experience an exogenous increase in the share of immigrant workers.[11] We seek to exploit time variations in the cross country-occupation distribution of immigrants to test this result.

Secondly, within country-occupation labor markets, we consider different economic sectors. Workers can move across sectors within a country-occupation labor market. According to an extension of our model to two sectors (see Appendix C), occupational employment should become more than proportionally concentrated into immigrants' receiving sectors. First, better employment prospects within a sector stimulate the hiring of unemployed natives. Second, provided that immigrants sort into sectors with higher sector-specific productivity, they may also stimulate the job-to-job transition of already employed natives in the same occupation from non-receiving sectors. That is, within occupations, we are likely to observe a crowding-in effect towards immigrants' receiving sectors.

Finally, while not directly testable, we seek to provide empirical support in favor of our main theoretical hypothesis: the disparities in the outside option of employment between immigrants and natives are the source of a positive externality on native employment. We first exploit heterogeneity among immigrants according to dimensions that affect their relative outside option, namely their duration of residence and country of origin. Second, we look for a differential impact across host countries due to institutional characteristics that also affect immigrant-native relative outside options. We consider

[9] Note that we are considering a short-run effect. In the long run, immigrants' outside option is likely to converge to that of natives. Labor opportunities (denoted by $\theta$) will then return to their initial level.

[10] As shown by D'Amuri and Peri (2014); Manacorda et al. (2012) or Steinhardt (2012), within age-education cells, immigrants and natives work in different occupations. Moreover, as noted by Dustmann et al. (2013), upon arrival, immigrants may work in occupations that do not correspond to their observed skill distribution. This downgrading positions recent immigrants at different percentiles of the native wage distribution than where we would expect them to be, based on their observed skills. Due to this downgrading, the pre-allocation of immigrants to skill groups leads to considerable misclassification errors.

[11] Our results are not in contradiction with the findings of Peri and Sparber (2011) and Peri and Sparber (2009), who find that less-skilled foreign-born workers specialize in occupations which are intensive in manual physical skills, while natives pursue jobs which are more intensive in communication-language tasks. The authors consider a very detailed definition of occupation which stands in stark contrast with our broad definition of occupation. Their finding concerning the reallocation of natives towards occupations which are more intensive in communication skills could perfectly correspond to movements within one of our broad occupations.

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disparities in immigrants' impact that could be attributed to differences across host countries in the immigrant-native unemployment benefit take up rate.

## 3. Data and descriptive statistics

The main dataset we use is the harmonized European Labour Force Survey (ELFS), which homogenizes and gathers country-specific surveys at the European level (see EUROSTAT (2009)). Due to data availability, we restrict our analysis to the 1998–2004 period. Our sample comprises the working age population (age 15–64) of Western European countries only. The data includes information on present occupation for employed individuals and past occupation for the unemployed, working status (employed, unemployed or inactive) and demographic characteristics. Unfortunately, the ELFS does not include any information on wages. We drop observations with missing data on country of birth, which are fundamental for our empirical analysis. In line with the previous literature, we classify foreign-born individuals as immigrants.

Individuals are grouped into cells on the basis of their occupation (used as a proxy for skills) and their country of work.[12] We combine the two to constitute our labor market. Occupations are broadly defined in 9 groups which are (1) senior officials and managers, (2) professionals, (3) technicians and associate professionals, (4) clerks, (5) service workers and shop and market sales workers, (6) skilled agricultural and fishery workers, (7) craft and related trade workers, (8) plant and machine operators and assemblers, and (9) elementary occupations. Because occupations are defined broadly, we cannot claim that immigrants and natives in a given country-occupation cell are perfect substitutes. However, defining occupations in a narrower way may be problematic for two reasons: first, workers can easily move from one occupation to another and, second, given the limited size of our sample, since we work with a labor force survey and not with a census extract (as is usual in the literature), there simply may not be a sufficient number of observations to reliably measure our dependent variable. While they may not be perfectly substitutable, immigrants and natives employed in the same occupation are more likely to have similar skills than those employed in the same group of education-experience level.

Defining a labor market at the national level, as in the seminal contribution of Borjas (2003), but in a multi-country context and with a broad definition of occupations within each country, has two key advantages. First, we can easily understand that moving from one country to another or from one occupation to another, even within the same country, is very costly for natives in the short run.[13] Therefore one can mitigate the spurious correlation introduced by the possibility for natives to “vote with their feet” by moving outside the labor market whose employment prospects worsen: the so-called *displacement effect* (see Card, 1990; Borjas, 2003 or Peri and Sparber, 2011). Therefore we can concentrate our analysis on studying the impact of yearly variations from 1998 to 2004 in the share of immigrants within labor markets defined at the country-occupation level.[14] Second, as shown in the following section, a grouping of individuals based on a measure of skill (occupation) and geography allows us to use an instrumental variable identification strategy based on historical settlement patterns across both host countries and occupations by origin country.

Consistently with our theoretical setup, we define our main labor market outcome of interest as natives' employment rate within country-occupation cells. We consider that unemployed natives belong to the occupation they had in their last job.[15] We exclude those who have never worked. Because the last occupation of unemployed workers is missing for France and the Netherlands, these countries are dropped from the main analysis.[16]

Considering the thirteen European countries in our sample,[17] from 1998 to 2004, the share of immigrants in the labor force increased by 6 percentage points from 5.7% to 11.8%, which is a large increase even compared with the US. Comparatively, in the US, this share increased from 12.7% to 14.7% over the same period (Migration Policy Institute, 2006).

The increase in the foreign labor force in Europe is even more striking if one considers the heterogeneity across occupations as shown in Fig. 1. While the increase extends across all occupations, it is higher for less-skilled occupations. However, contrary to conventional wisdom, the contribution of immigrants to more-skilled occupations is also rising and significant.

Fig. 2 provides a scatter plot of the share of immigrants in 2004 against the share at the beginning of the period (1998 for all countries, except for Germany whose data start in 2002). Three points from this graph are worth mentioning. First, confirming the results from Fig. 1, the points in Fig. 2 lie above the 45° line for most countries and occupations, which is an evidence of the pervasive increase in the share of immigrants over time. Second, there are important disparities in the share of immigrants across occupations within countries, and across countries within occupations. Third, and central to our difference-in-differences identification strategy, the increase over time in the share of immigrants has not been equal across

[12] Previous papers have already exploited cross-occupation variation (see Card and DiNardo (2000), Card (2001), Hunt (1992), Patel and Vella (2007) or Steinhardt (2012)).

[13] The choice of such broadly defined occupations depends rather on initial training and other long-term investments, which is unlikely to be affected by small changes in employment opportunities.

[14] It would be interesting to investigate possible heterogeneous effects across age groups (see Smith, 2012). However, with data in hand, it would be impossible to implement our instrumental variable approach since we do not have the age of immigrants at entry.

[15] Given the broad definition of occupations, it is reasonable to assume that unemployed workers are essentially searching in the same occupation.

[16] They are reintroduced in the country-occupation-sector analyses, where we only require data on the employment level.

[17] Countries included in the country-occupation analysis are: Germany, Belgium, Portugal, Sweden, Norway, Ireland, Spain, Italy, Denmark, Luxembourg, Austria, Finland and Greece.

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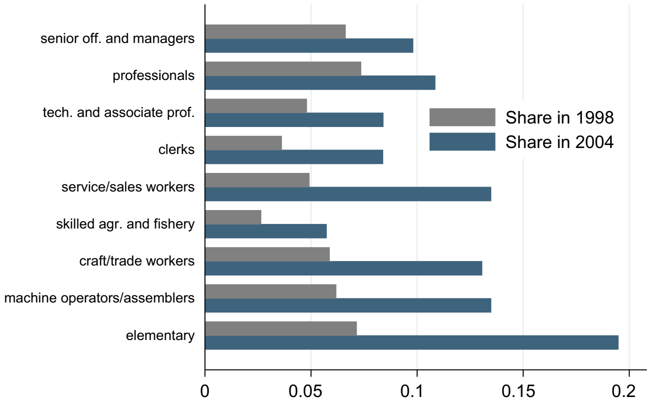


**Fig. 1.** Share of immigrants in the labor force by occupation.

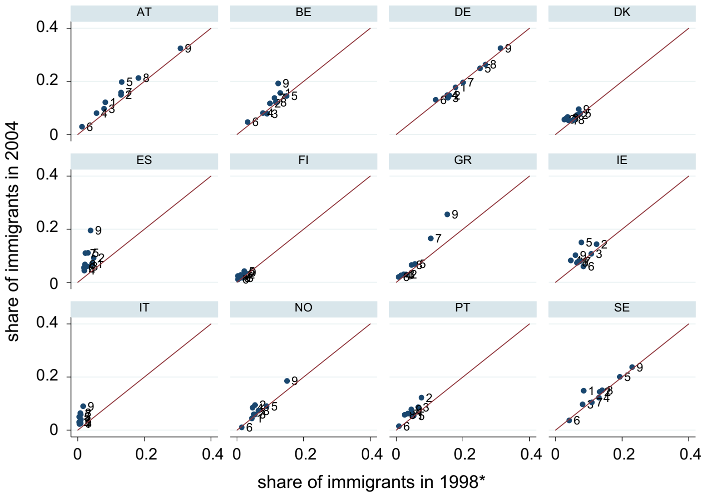


**Fig. 2.** Share of immigrants in country-occupation cells between 1998 and 2004. The data for Germany is from 2002 to 2004. For ease of reading, Luxembourg is not featured in the figure, although it is included throughout the econometric analysis. The 45° line is represented in each graph. Numbers refer to occupations, with (1) senior officials and managers, (2) professionals, (3) technicians and associate professionals, (4) clerks, (5) service workers and shop and market sales workers, (6) skilled agricultural and fishery workers, (7) craft and related trade workers, (8) plant and machine operators and assemblers, and (9) elementary occupations.

occupations within countries. This can be seen in the graph by noting that the further above the bisecting line the country-occupation cells, the larger the immigration share change. Depending on the country, we can see substantial variation in the extent to which occupations have been affected by immigration.

Overall, Fig. 2 shows that an identification strategy which relates the differential changes over time in the share of immigrants across occupations within countries to the corresponding time variation in the natives' employment rate (a difference-in-differences strategy) may unravel the impact of immigrants on natives' employment rate.[18] The rest of the paper will then seek to exploit changes over time in this heterogeneity across occupations within countries to identify the impact of immigrants on natives' employment rate.

## 4. Empirical specification, identification strategy and results at the occupation level

### 4.1. Empirical specification and identification strategy

We start by explaining the empirical specification and identification strategy for the impact of immigration on natives' employment at the occupational level. Our point of departure is Eq. (6) in our model which easily delivers the positive

[18] In the web Appendix, we plot the log change in the share of natives' employment rate between the first and last available year for each country-occupation cell against the corresponding change in the share of immigrants. The graph displays a slightly positive relationship between the two variables. Interestingly, the slope of the relationship (0.017), which is statistically significant at 5%, is very similar to the one we get in the more data-demanding fixed effect estimates that are implemented in the econometric analysis.

relationship between natives' employment rate in a particular labor market and the probability of finding a job:

$$\frac{n_N}{P_N} = F\left(\underbrace{p(\theta)}_{+}\right), \tag{7}$$

Because the probability of finding a job in a particular labor market segment increases with labor market tightness, $\theta$, and the labor market tightness itself increases with the proportion of immigrants in the market (see Eq. (5)), we can easily[19] derive our baseline estimating equation which establishes the positive link between the native employment rate and the share of immigrants in the cell:

$$\ln y_{cot} = \beta_0 + \beta_1 \ln shim_{cot} + \alpha_o + \alpha_c + \alpha_t + \alpha_{co} + \alpha_{ct} + \alpha_{ot} + u_{cot} \tag{8}$$

where $y_{cot}$ denotes the employment rate of natives in occupation $o$, in country $c$, at time $t$. The natives' employment rate in a particular cell is defined as the ratio between the number of employed natives in the cell over the total native active population (which corresponds to the sum of employed and unemployed natives in the cell). The key explanatory variable, $shim_{cot}$, is the share of immigrants in the labor market $co$ at time $t$. The numerator of this variable includes both employed and unemployed immigrants in cell $co$, while the denominator equals the active population in the cell (natives and immigrants, both employed and unemployed). The parameter $\beta_1$ then measures the elasticity of the native employment rate in a particular country-occupation labor market $co$ with respect to the share of immigrants in that market segment.

Heterogeneity across labor markets that is country- and/or occupation-specific is absorbed by the country and occupation fixed effects. Aggregate labor market shocks are absorbed by year-specific fixed effects. We also allow labor demand shocks to have a country-specific component and an occupation-specific component (due for instance to technological changes) by including country-by-year and occupation-by-year fixed effects ($\alpha_{ct}$ and $\alpha_{ot}$, respectively). More importantly, in all our estimations we will introduce country-by-occupation fixed effects to control for the sorting of immigrants into labor markets whose structural determinants of employment are better (higher $m_0$ for instance) and, at the same time, have a higher native employment rate. Our model allows us to relate changes in the share of immigrants in a particular labor market to changes in natives' employment rate in that labor market. Thus we seek to achieve identification by exploiting the cross-sectional time variation within labor markets defined at the country-occupation level. The model with the full set of fixed effects amounts to a triple difference estimation strategy: we exploit variation over time in the share of immigrants across occupations and countries.

This wide set of fixed effects distinguishes our approach from standard cross-area studies that cannot control for such factors as they either use single cross-sectional data (Card, 2001) or from studies that exploit single country aggregate time series data (Borjas, 2003).

Because serial correlation within a particular labor market $co$ is a concern, in all regressions we adjust standard errors for the clustering of observations at the country-occupation level. We also use weighted least squares with weights equal to the native population size in each occupation in the base year period 1998.[20] It is important to note that the active native labor force in a cell appears in the denominator of both sides of Eq. (8), which may potentially create a spurious positive correlation between the immigrants' share within an occupation and the natives' employment rate. For this reason, the share of immigrants is computed by setting the denominator to its 1998 value, our first period for data.[21] In some specifications, we also directly control for the size of the native labor force in the cell. In this way, time variations of $shim_{cot}$ within a particular labor market stem only from changes in the number of immigrants and not from natives' inflow or outflow.

Despite our effort to control for non-time varying unobservable determinants of natives' employment rate potentially correlated with the immigrants' share within an occupation, endogeneity biases still remain a concern. This is the case for instance if changes in the immigrants' share within a labor market are correlated with changes in unobserved determinants of employment within that labor market. It is indeed plausible that within a country, immigrants would sort into occupations whose demand is growing. In that case, country-specific occupation fixed effects are not sufficient, since occupation-specific employment rates are not fixed within countries. We address this issue with two strategies. First, we control partly for labor-market-specific productivity shocks using a labor-market-specific demand shift index. If an occupation is concentrated in an industry whose employment has grown more than average over the period, we expect the labor demand for this occupation to have grown more than average and, at the same time, to have drawn more immigrants and natives within that occupation. To control for this possibility, in our estimated equation, we introduce an occupation-specific labor demand

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[19] The probability of finding a job is given by $p(\theta) = m_0\theta^{1/2} = \frac{m_0^2}{\gamma(r+s)} \cdot \Omega(shim)$, where $shim$ stands for the share of immigrants in the labor market. $\Omega(shim) = h - \overline{w}$ is the expected profitability margin of employers which increases with the share of immigrants. The term $\frac{m_0^2}{\gamma(r+s)}$ captures the structural determinants of employment (i.e. matching efficiency, separation rate and the cost of opening a vacancy) which we assume that are specific to a given labor market and are non-time varying. Given a population composed of various segmented labor markets observed over several years, the natives' employment rate in a given labor market in a given year can be decomposed as

$$\log\left(\frac{n_N}{P_N}\right)_{cot} = E(\log F(p(\theta_{cot})) | shim_{cot}) + \varepsilon_{cot}$$

where $\varepsilon_{cot}$ is a random error component which is independent from $shim$. Taking logs on both sides and assuming that the conditional expectation function for log $(F(p(\theta)))$ admits a linear approximation, we obtain our baseline estimating equation.

[20] Using a fixed weight ensures that our results are not affected by changes in the native population size across occupations due to immigrants.

[21] This first period has been set to 2002 for Germany since earlier data were not available for this country.

shift index driven by the sectoral composition of natives' occupational employment at the national level : $(Demand\ shift)_{cot}$. Thus, we achieve identification using employment rate deviations from occupation-specific trends determined by the initial sectoral composition of occupations in every host country. In the spirit of Katz and Murphy (1992) or Katz and Blanchard (1992), this labor demand shift index is constructed as follows:

$$(Demand\ shift)_{cot} = \sum_j \gamma_{coj1998} L_{cjt} \tag{9}$$

where $L_{cjt}$ is aggregate natives' employment at the two-digit industry level $j$ at date $t$ in country $c$,[22] and $\gamma_{coj1998} = \frac{E_{coj1998}}{\sum_j E_{coj1998}}$ is the share of native workers in occupation $o$ employed in industry $j$ in 1998 in country $c$,[23] excluding immigrants. We interpret the demand shift index as the predicted employment for native workers in an occupation given the distribution of that occupation across sectors.

The estimation of our coefficient of interest, $\beta_1$, will not be biased by the correlation between immigrants' inflows into a particular labor market and better employment prospects for natives due to the labor-market-specific demand shocks driven by the sectoral composition of that particular labor market.

Our second approach to deal with endogeneity biases uses an instrumental variable strategy. This requires a variable correlated with the influx of immigrants into a given labor market but uncorrelated with unobserved factors driving employment growth among natives. We extend the strategy originally developed by Altonji et al. (1991) to a multi-country-occupation setting and use historical settlement patterns across both host countries and occupations by origin country as an instrument for current inflows. Because of informational networks, immigrants tend to cluster into labor markets with a higher share of their country peers.[24] Such an instrument has proven to be a strong determinant of contemporaneous inflows in the single-country case.[25] To date, Angrist and Kugler (2003) and D'Amuri and Peri (2014) are the only ones who use a similar instrument in a multi-country setting, although they do not focus on occupational choices within countries.

Due to the lack of data on the past share of immigrants across occupations by origin country, our instrument is constructed in two steps. First, following Altonji et al. (1991), we construct a predicted number of immigrants having three different levels of education (primary, secondary, and tertiary) for each country and year as follows:

$$\phi_{cst} = \sum_m Stock_{cms1990} * \frac{Flow_{OECDmt}}{Stock_{OECDm1990}}, \quad t = 1998, \ldots, 2004$$

where $Flow_{OECDmt}$ is the flow of immigrants from country $m$ in year $t$ into the OECD, $Stock_{OECDm1990}$ is the number of immigrants from country $m$ in the OECD in 1990,[26] and $\mathrm{Stock}_{cms1990}$ is the number of immigrants from country $m$ in country $c$ with education level $s$ in 1990. Data on immigrant yearly flows from origin into destination countries are gathered from the OECD and those on 1990 stocks are from Docquier et al. (2007).

Finally, we distribute these predicted immigrant flows by educational level across occupations according to the natives' educational distribution by occupation in 1998.

Specifically, our predicted inflow of immigrants across occupations, countries and over time is

$$(Predicted\ inflow)_{cot} = \sum_{s=1}^{3} \phi_{cst} * \eta_{cso1998} \tag{10}$$

where $\eta_{cso1998}$ is the share of natives of education level $s$ employed in occupation $o$ in country $c$ in 1998. The validity and quality of our instrument relies on two assumptions: (*i*) the skill distribution of natives across occupations in 1998 is not affected by immigrants, which amounts to an exclusion restriction and (*ii*) for the quality of our first stage, the educational distribution of immigrants across occupations is correlated with that of natives, although it need not be the same.[27] For our exclusion restriction to be valid, we require the natives' distribution within each educational group across occupations, $\eta_{cso1998}$, to be independent from immigrants' labor supply shock. To check the robustness of our results to that assumption,

[22] The labor demand shift variable has also been computed using the real level of production instead of aggregate employment at the two-digit industry level. Industrial production data is obtained from the EUKlems consortium (http://www.euklems.net/). Unfortunately, data on the real level of production is not available for all countries, so we are missing some observations when using this variable.

[23] We have also constructed an index with the average level of occupation share over the whole 1998–2004 period. This index gives similar results.

[24] The importance of social networks for the location decisions of migrants coming from Mexico to the US and their labor market outcomes has been examined by Munshi (2003). Working with 1980, 1990 and 2000 US Census data, Patel and Vella (2007) find that the occupational share of certain ethnic groups grew drastically in particular labor markets over the period from 1980 to 2000. Moreover, the pattern of growth is consistent with the presence of network effects. The data also does not appear to suggest that the allocation observed is the result of sorting on the basis of a comparative advantage.

[25] See Card (2001), Card (2009), Cortes and Tessada (2011) or Patel and Vella (2007) for the US and Gonzalez and Ortega (2011) or Farré et al. (2011) for Spain among others.

[26] We consider the stock in the whole OECD which we believe is more exogenous than considering the stock of immigrants from country $m$ in country $c$. This stock is more influenced by the economic conditions of the host country.

[27] Note though that we do not need the educational distribution of immigrants across occupations to be the same as that of natives which, as shown by Dustmann et al. (2013), D'Amuri and Peri (2014), Manacorda et al. (2012) or Steinhardt (2012), may be unrealistic. The only required assumption is that the two are correlated. The lack of correlation between the two distributions will weaken our first stage, but alone it is not a violation of the exclusion restriction.


**Fig. 3.** Immigrants' predicted share (instrument) against immigrants' observed share. Male sample.

in the following section we experiment with alternative rules to distribute immigrants within skill groups across occupations.[28]

Fig. 3 shows the scatter plot of the (log) share of immigrants against our (log) shift-share instrumental variable. The latter is obtained by dividing the predicted inflows obtained in Eq. (10) by the predicted number of natives in the cell in 1990.[29] The figure illustrates the strong (unconditional) correlation between the two variables making, at first glance, our shift-share variable a good candidate to instrument changes in the share of immigrants within country-occupation labor markets. As shown further in this paper, this result is confirmed by the first-stage IV regression. Provided that flows into the OECD by origin country and the past distribution of immigrants by origin country across labor markets are independent from current demand shocks affecting a particular labor market, our predicted inflows can be used as an IV to identify the impact of immigrants on natives' employment rate.

### 4.2. *Results at the occupation level*

The first two columns of Table 1 present OLS estimates of the elasticity[30] of native males' employment rate[31] with respect to the immigrants' share within a country-occupation labor market, controlling for all the fixed effects specified in Eq. (8). Appendix B shows how the results change when we successively add country-by-occupation, occupation-by-year, and country-by-year fixed effects (columns 2–4).

To make sure that changes in the immigrants' share within a labor market are driven by immigrants and not by changes in the number of natives, the denominator of our independent variable is set to its 1998 value.[32] Further, introducing the log number of natives in an occupation into the regression does not alter this result. We are therefore confident that, given the broad definition of labor markets, the bias due to displacement across occupations is negligible in our context, and does not confound our estimated impact.

A potentially more serious concern is the unobserved, time-varying, labor market demand shocks that may be correlated with inflows of immigrants. As a first attempt to assess and partially control for this possibility, in column (2) of Table 1 we introduce the occupation-specific industry-driven labor demand shift presented in Eq. (9). The coefficient associated with

[28] In order to deal with the endogeneity of occupation choices by immigrants and natives, we have implemented additional checks. More precisely, in the spirit of Autor et al. (2003) and following a referee's suggestion, we have defined a dependency index of occupation $o$ in country $c$ on immigrant labor. First, we compute the dependency index as the share of immigrants in the country-occupation cell $co$ in 1990 (we have also tested for 1998). This index is then multiplied by the log share of immigrants in country $c$ in year $t = 1998, 1999, \ldots, 2004$. We then replace our explanatory variable $\log(shim_{cot})$ by $\log(shim_{ct}) \cdot shim_{co,1990}$. The estimation results were consistent with our benchmark IV estimates. The results are presented in a web Appendix on the authors' websites.

[29] The number of natives in each occupation in 1990 is computed using information on the number of natives in each schooling level in 1990 and the distribution of natives in each schooling level across occupations in 1998.

[30] When assuming logs for the independent variable, small changes in the share of immigrants in an occupation, i.e. from zero to a positive share, imply a huge percent change. However, there should not be a large effect on the overall employment rate of natives since the number of additional immigrants is small. In a web Appendix (available on the authors' websites), we propose additional robustness checks where (i) we eliminate the log of the independent variable, (ii) we drop cells with a low immigrant share, and (iii) we eliminate the log from the dependent variable. Qualitatively, all results provide an impact which is positive and imply an elasticity of a similar magnitude. In the paper, we choose to stick to the log-share specification since it allows straightforward computations of elasticities.

[31] Due to the specificities of female labor force participation, we decided to restrict the analysis to males.

[32] The estimated coefficient on the immigrants' share remains statistically unchanged if instead we use the native national workforce for the denominator of the dependent variable and the total workforce for the denominator of the independent variable. Additional robustness tests are available on the web Appendix on the authors' websites.

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**Table 1**
The effect of immigrants on the labor market (country-occupation labor market). Male sample 1998–2004.

| Estimation method | Dependent variable: log (natives' employment rate $)_{cot}$ | | | | | |
|---|---|---|---|---|---|---|
| | OLS | | 2SLS | | | |
| | | | IV1 | IV1 | IV2 | IV3 |
| | (1) | (2) | (3) | (4) | (5) | (6) |
| Independent variable | | | | | | |
| ln(immigrants/workforce $98)_{co}$ | 0.023*** | 0.021*** | 0.052*** | 0.047*** | 0.047*** | 0.040*** |
| | (0.006) | (0.006) | (0.019) | (0.017) | (0.017) | (0.015) |
| Additional control labor demand shift index | NO | YES | NO | YES | YES | YES |
| Observations | 654 | 654 | 654 | 654 | 654 | 654 |
| First stage on excluded instrument (Kleibergen-Paap rank Wald F statistic) | | | 9.121 | 8.840 | 9.557 | 10.11 |

*Note*: units of observation are country-occupation, *co*, in each year *t* from 1998 to 2004. The dependent variable is the log share of employed natives in the country-occupation cell in a given year. The explanatory variable of interest is the log of the share of immigrants in the workforce of a country-occupation cell in a given year. The size of the workforce cell has been set to its initial value in the first period (1998 for all countries, except Germany for which it is 2002). In addition to country, year and occupation fixed effects, country-by-year, country-by-occupation and occupation-by-year fixed effects are included in all regressions. The labor demand shift is based on the sector distribution of occupations with the** weights of sectors based on national sectoral employment. Standard errors clustered at the country-occupation level are reported in parentheses. Regressions are weighted by the total number of natives in a country-occupation cell in 1998.
Statistical significance level: * $p<0.1$, ** $p<0.05$, *** $p<0.01$.

the immigrants' share remains largely unaffected. This result suggests little correlation between changes in the immigrants' share within an occupation and the labor demand shift driven by the sectoral composition of occupational employment at the national level.

In spite of controlling for country-specific occupation fixed effects and labor demand shift across occupations, OLS estimates may still be contaminated by time-varying unobservable demand shocks correlated with the immigrants' share. Indeed, our labor demand shift index controls only for changes in employment within an occupation driven by the sectoral composition of occupations. However, it is well known and documented (see Acemoglu and Autor, 2011) that occupational employment has evolved over time due to other factors. One predominant factor is technological change that fosters relative demand for some occupations, in all sectors, in which immigrants may be flowing. We then turn to IV estimates, presented in the last four columns of Table 1. The first-stage *F* test on our instrument is presented in the lower panel of Table 1. Not surprisingly, as already suggested by Fig. 3, the first-stage IV regression coefficients associated with the instrument are fairly significant given the number of fixed effects introduced. In all specifications, the first-stage F-stat is close to 10.

Columns (3) and (4) present estimation results when employing our preferred instrument (described in the previous section). Column (4) introduces the labor demand shift driven by the sectoral composition of occupations as an additional control. IV estimates suggest a small but positive effect of immigration on male natives' employment rate: a 10% increase in the share of immigrants within an occupation increases natives' employment rate within that occupation by 0.47%. Reassuringly, this estimate is robust to the introduction of the industry labor demand shift, which comforts the exogeneity of our instrument. Estimates by IV give a larger point estimate than those by OLS, however standard errors are also much larger, so that there is no statistically significant difference between both estimations. Although the estimated impact is small, it is important to note that the share of immigrants in most occupations more than doubled over our sample period, suggesting that the employment creation effect due to immigrants, over this short period of time, may have been substantial.[33]

For our exclusion restriction to be valid in our benchmark estimation, we require the natives' distribution within each educational group across occupations to be independent from immigrants' labor supply shock in the same country. As a robustness check, we reproduce the estimation with two alternative instruments using different rules to distribute immigrants within skill groups across occupations.

[33] Previous studies have already exploited cross-occupation variation. Card and DiNardo (2000) is one of the few and seminal papers that focuses explicitly on occupations (three occupational groups observed at the city level) with US data. In their preferred specification choice (controlling for city-specific time trends) they find that immigrants actually pull natives into sectors and occupations in which they work, rather than push them out. Using 1990 US census data, Card (2001) also finds that occupation-specific wages and employment rates are systematically lower in cities with a higher relative supply of workers in a given occupation. Hunt (1992) examines the impact of people repatriated from Algeria to France in 1962 on non-repatriate wage cells defined at the occupation-department level (geographical unit below the region). Using data on the US Census, Patel and Vella (2007) exploit cross-occupational variation to analyze the role of networks as an explanatory factor for the increase in the occupational share of certain ethnic groups in particular labor markets from 1980 to 2000. In a more recent paper, Steinhardt (2012), for the case of Germany, concludes that an analysis based on education-experience cell grouping underestimates the impact of immigration compared to an analysis based on occupation grouping. While our paper exploits also cross-occupation variation, coefficient estimates are difficult to compare since the explanatory and the dependent variables are systematically defined differently from ours.

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First, we use the distribution of natives in each education group across occupations computed only on the subset of workers over 40 years old. We think that the exclusion restriction for these workers is more likely to hold because they are less mobile than younger workers. Therefore, immigrants are very unlikely to trigger a reallocation of native workers over 40 years old across occupations within a given level of education (high, middle, or low). This is particularly true given the relatively broad definition of occupations. The predicted number of immigrants is then computed as follows:

$$\text{Predicted Flows } 2_{cot} = \sum_{1}^{3} \phi_{cst}\eta_{cso1998(age>40)},$$

where $\phi_{cst}$ is the predicted number of immigrants of skill level $s$ ($s=$ High, Low, Middle) in country $c$ at period $t$, computed using the ethnic-network-based shift-share procedure. The variable $\eta_{cso1998(age>40)}$ denotes the share of skilled natives $s$ employed in occupation $o$ in country $c$ in 1998 among native workers over 40.

Second, we define a new instrument where, instead of using the distribution of natives of a given skill across occupations in the country, we use the average distribution computed over immigrants across all countries, except country $c$ for which we want to identify the impact. Our aim here is to catch occupation-specific characteristics which are common across all immigrants' host countries, explaining why this particular occupation employs a particular mix of workers according to their skills. The predicted number of immigrants is then computed as follows:

$$\text{Predicted Flows } 3_{cot} = \sum_{1}^{3} \phi_{cst}\eta_{EUso1998},$$

where $\eta_{EUso1998}$ denotes the share of immigrants with skill $s$ employed in occupation $o$ in 1998 in EU 15 countries, excluding country $c$.

The results provided by $IV2$ and $IV3$ in columns 5 and 6 are perfectly consistent with the results provided by our preferred instrument in columns 3 and 4. We are thus confident of the quality of the instruments. Moreover, in a web Appendix (available on the authors' websites), we implement additional robustness checks for a larger set of countries by only using data on individuals who are employed (this avoids losing data for France and the Netherlands). The results from these alternative specifications confirm our benchmark conclusions.[34]

## 5. Empirical specification, identification strategy and results at the occupation-sector level

### 5.1. Empirical specification and identification strategy

The assessment of immigrants' impact on a particular labor market has been blurred by the possibility for natives to leave a labor market hosting more immigrants. In our context, the labor market is defined by broad occupations across countries, so that moving across these labor markets is too costly. However, workers in a country, within a given occupation, can move across sectors in response to changing employment opportunities. Indeed, a noteworthy prediction of our search and matching model is that, within a particular labor market, the employment share of sectors receiving more immigrants should increase. Improved employment opportunities in immigrants' receiving sectors benefit unemployed native workers and may also pull native workers from other sectors (see the model extension presented in Appendix C).

While most of the literature has focused on a crowding-out effect, whereby native mobility has an offsetting effect on the supply shock created by immigrants, instead our model is consistent with a crowding-in effect.

To investigate the direction of natives' displacement within occupations, we group workers into nine occupations and three sectors: services, manufacturing and construction. The latter is traditionally an important employment sector for immigrants. Then, we keep a consistent definition of occupations throughout the analysis. This will ease the comparison of these results with those obtained at the occupational level.

To evaluate the possibility of a crowding-in effect, we adopt the same specification proposed in Card (2001) or Cortes (2008). We estimate the following equation:

$$\begin{aligned}&\frac{(\text{Total employment in occupation } o \text{ and sector } j)_{ct}}{(\text{Total employment in occupation } o)_{ct}}\\ &\quad=\beta_0+\beta_1\frac{(\text{Immigrants' employment in occupation } o \text{ and sector } j)_{ct}}{(\text{Total employment in occupation } o)_{ct}}\\ &\quad+G(c,t,s,o)+\varepsilon_{ojct}\end{aligned} \tag{11}$$

[34] We have also run regressions country by country. However, with the exception of a few countries, the coefficient estimates on immigrants' share were very imprecise and the first-stage results were much weaker. Cross-country variations allow contrasting time variation within the same set of occupations across different countries; these variations are then useful to estimate an average impact with precision and with limited bias. Using cross-country data also allows substantially improved precision since there is less correlation in outcomes across countries than within countries. Nevertheless, we made some progress in investigating cross-country heterogeneity by distinguishing the impact according to country-specific labor market institutions in Section 6.3.

**Table 2**
The crowding-in within sectors (country-occupation-sector labor market). Male sample 1998–2004.

| Dependent variable: | Share of occupational employment in a sector | | | | | |
|---|---|---|---|---|---|---|
| Estimation method | OLS | | | IV | | |
| Independent variables | (1) | (2) | (3) | (4) | (5) | (6) |
| Immigrants' share in a sector-occupation | 0.780*** | 0.784*** | 0.791*** | 0.973 | 1.326** | 1.460*** |
| | (0.044) | (0.042) | (0.042) | (0.133) | (0.151) | (0.155) |
| Additional fixed-effects | | | | | | |
| Country by occupation by sector | NO | YES | YES | NO | YES | YES |
| Country by sector by year | NO | NO | YES | NO | NO | YES |
| First-Stage F on excluded instrument | | | | 258.9 | 173.8 | 182.5 |
| Observations | 2303 | 2303 | 2303 | 2303 | 2303 | 2303 |

Units of observation are at the country-occupation-sector level, *coj*, in each year *t* from 1998 to 2004. In addition to the fixed effects indicated in the table, in all regressions we include fixed effects for all possible two-way interactions of country by year, country by occupation, country by sector, occupation by year, occupation by sector, and sector by occupation. The immigrants' share independent variable is equal to the ratio of immigrants' employment in a sector-occupation to the total employment in the occupation in a given country year. The shift-share instrumental variable is computed using the predicted number of immigrants in a sector occupation cell for every year and country.

*** Statistical significance with respect to the value of 1 at 1%

** Statistical significance with respect to the value of 1 at 5%

where $G(c,t,s,o)$ stands for different sets of fixed effects whose aim is to control for unobserved labor demand shocks. In our regressions, we control for all possible two-way interactions between country, occupation and sector. Additionally, by systematically including country-by-occupation-by-sector fixed effects, we identify the displacement effects by exploiting time variation in the contribution of immigrants to the sectoral concentration of occupations within countries.

There is no displacement of natives if $\beta_1 = 1$. If the arrival of one immigrant leads some natives to leave the sector, $\beta_1 < 1$. There is a crowding-in effect (i.e. natives are pulled into sectors hosting more immigrants) if $\beta_1 > 1$.

Despite our efforts to control for non-observable labor demand shocks with this set of fixed effects, immigrants may still not be randomly located across sectors and occupations within countries. We deal with this bias due to unobserved time-varying and country-specific sectoral and occupational labor demand shocks using an IV strategy which generalizes the previous approach. We now consider the case in which past immigrants' networks are defined at the sector and occupation level, instead of occupations only. Specifically, once we compute the predicted number of immigrants by educational level, we obtain the predicted number of immigrants by occupation and sector by replacing the parameter $\gamma_{cso}$ by $\gamma_{csoj}$, in Eq. (10), where $\gamma_{csoj}$ is the share of education level $s$ employed in occupation $o$ and sector $j$ in country $c$ in 1998. The first-stage relation for this instrument is, as previously, strong and highly significant, suggesting that network effects do not only determine the sorting of immigrants across countries and occupations, but also their distribution across the different sectors of the economy within occupations in each country.

### *5.2. Results at the occupation-sector level*

In Table 2, we present the OLS and IV estimates of Eq. (11). The OLS estimates suggest a significant outward displacement. However, this effect disappears once we control for the endogeneity bias with IV estimates (country-occupation and sector-specific fixed effects are also included in the regression). First-stage estimates reveal that the instrument is strong. In all specifications, the coefficients have the expected sign with an F-stat on instrument exclusion above 100 in all cases. The IV estimates, in columns 5 to 6, clearly point towards a crowding-in effect. The results are robust to the inclusion of country-by-year and by-sector fixed effects to control for sector-specific labor demand shocks that may increase the concentration of workers from all occupations in those sectors. The coefficients are significantly different from unity in both specifications. These results suggest that following an inflow of immigrants, occupational employment becomes more than proportionally concentrated into immigrants' receiving sectors: there is a crowding-in of natives from other sectors or from unemployment into those sectors.

Although the coefficients cannot be compared with those found in the occupational-level analysis, the displacement effect is fairly important, even if the effect is estimated with large standard errors. This suggests that, potentially, immigrants' employment creation effect is also associated with job-to-job turnover among natives who are beyond transitions from unemployment to employment. This could explain the comparatively small effect of the employment gain at the occupational level. The impact of immigrants on natives' job turnover has also been found in some recent papers and helps to explain a potentially positive effect of immigrants on natives' wages as well (see Ortega and Verdugo, 2014 for the case of occupational turnover in France).

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The results on the displacement effect are consistent with a positive employment effect at the occupational level. Indeed, if the increasing concentration of natives came entirely from the job-to-job transitions of already employed workers, we would not find a net effect on employment at the occupational level. However, the results on the displacement effect, once combined with the positive impact found at the occupational level, are consistent with the view that employers respond positively to immigration inflows by opening more vacancies in the sectors concerned. These results provide some evidence of the fact that immigrants do not lead natives to reallocate towards other sectors. In the rest of the paper, we seek to provide further empirical arguments supporting the assumption of a lower outside option for immigrants, on which our theoretical arguments are based.

## 6. How relevant is our mechanism?

In this section, we provide some evidence that the lower reservation wage of immigrants is a plausible mechanism to explain the short-run positive impact of immigrants on natives' employment rate. We exploit the fact that the outside option varies across various groups of immigrants to gauge the relevance of our mechanism. First, we emphasize differences across immigrants with respect to their duration of residence and their country of origin. Next, we exploit institutional heterogeneity across host countries leading to diverging outside option gaps between natives and immigrants.

### 6.1. Distinguishing new immigrants from long-term stayers

Our baseline Eq. (8) does not distinguish between veterans and earlier immigrants which amounts to assuming that all immigrants have the same outside options. This is unlikely to be the case.[35] With years of residence in the host country, immigrants become eligible to unemployment benefits, they develop their social networks and know the labor market better. Therefore, the value of their outside option should converge to that of natives, and their positive impact on natives' employment rate should be lower.

To relax the assumption of identical outside options, within an occupation, we distinguish immigrants with less than 10 years of residence (low outside option group) from those with more than 10 years (high outside option group). In a way, if immigrants *assimilate*, their outside option improves over time, converging to that of natives. As immigrants become more substitutable with natives with respect to their profitability for employers, incentives to open more vacancies decrease and their positive impact on natives' employment rate should vanish.

Let $y_{cot}$ be the employment rate of natives in occupation $o$, country $c$ and year $t$, $shim_{cot1}$ the ratio of immigrants with less than or equal to 10 years of residence in cell $cot$ to the total population of the cell in 1998, and let $shim_{cot2}$ be the same ratio for immigrants with more than 10 years of residence.[36] The equation to be estimated then becomes

$$\ln y_{cot} = \gamma_0 + \gamma_1 * \ln(shim_{cot1}) + \gamma_2 * \ln(shim_{cot2}) + \alpha_o + \alpha_c + \alpha_t + \alpha_{ot} + \alpha_{co} + \alpha_{ct} + u_{cot} \tag{12}$$

This specification assumes a piecewise linear impact of immigrants on natives' employment rate that depends on their years of residence within a host country. Under the outside-option assimilation hypothesis, our testable assumption becomes $\gamma_1 > \gamma_2$.

Since we face the same identification issues as in Eq. (8), as an additional instrument we use an immigrant's specific labor demand shift index. We exploit the fact that sectoral labor demand shocks at the national level may have a differential impact on immigrants' sorting across occupations due to the past distribution of immigrants' occupational employment across sectors. Thus our instrument generalizes the labor demand shift index, but uses the past distribution of immigrants within occupations across sectors as fixed weight. To be specific, let $\tilde{\gamma}_{coj,1990}$ denote a proxy for the share of immigrants in occupation $o$, who worked in sector $j$ in country $c$ in 1990,[37] then our second instrument is

$$(Demand\ shift)^{immigrants}_{cot} = \sum_j \tilde{\gamma}_{coj,1990} Y_{cjt} \tag{13}$$

where $Y_{cjt}$ denotes the two-digit real industrial output in sector $j$.[38] Our identification assumption is the following: immigrants will be differently attracted to occupations due to (i) their past distribution across sectors within an occupation, and (ii) due to sectoral output shocks at the national level. Our identification assumption relies on the exogeneity of past immigrants' distribution across sectors and occupations.

[35] As suggested by Chiswick (1978), Borjas (1994) or Borjas (1999) for the US, Chiswick et al. (2005) for Australia, Friedberg and Hunt (1995) for Israel, or Lam and Liu (2002) for Hong Kong, with years of residence, immigrants' wages converge to those of natives. The implicit idea in these papers is that immigrant workers acquire language skills and other productive assets making them closer substitutes to natives.

[36] The choice of tenure spells in host countries is a trade-off between having a sufficient number of observations within each cell and a sufficient variation to allow for identification. The EU-LFS does not distinguish a duration of residence above 10 years.

[37] As previously, $\tilde{\gamma}_{coj,1990}$ is computed in two steps: first, we consider the distribution of immigrants across educational levels (low, medium, and high) in 1990; second, we distribute them across occupations according to natives' educational distribution by occupation in 1998.

[38] When using the "Demand shift" variable as an instrument, rather than as a control (as we have previously done), we prefer to use the real industrial output of the corresponding sector rather than aggregate employment in the sector, since total employment would also include immigrant workers.

**Table 3**
Heterogeneity among immigrants with different durations of residence and from different origin countries. Male sample.

| Estimation method | Dependent variable: log(employment rate of natives in the cell) | | | | | |
|---|---|---|---|---|---|---|
| | OLS | IV | IV | OLS | IV | IV |
| Independent variables | | | | | | |
| ln(New immigrants/worforce 1998)$_{cot}$ | 0.006* | 0.035** | 0.040** | | | |
| | (0.004) | (0.016) | (0.016) | | | |
| ln(Veteran immigrants/worforce 1998)$_{cot}$ | 0.007** | 0.010 | | | | |
| | (0.003) | (0.025) | | | | |
| ln(EU15 immigrants/workforce 1998)$_{cot}$ | | | | −0.009*** | −0.013 | |
| | | | | (0.003) | (0.023) | |
| ln(Non-EU15 immigrants/workforce 1998)$_{cot}$ | | | | 0.019*** | 0.053*** | 0.042*** |
| | | | | (0.005) | (0.015) | (0.013) |
| First-Stage $F$ test of excluded instruments | | | | | | |
| ln(New immigrants/worforce 1998)$_{cot}$ | | 17.84 | 15.10 | | | |
| ln(Veteran immigrants/worforce 1998)$_{cot}$ | | 5.47 | | | | |
| ln(EU15 immigrants/workforce 1998) | | | | | 14.09 | |
| ln(Non-EU15 immigrants/workforce 1998) | | | | | 9.72 | 14.15 |
| Observations | 519 | 519 | 519 | 522 | 522 | 525 |

Units of observation are country-occupation cells, $co$, in each year $t$ from 1998 to 2004. The dependent variable is the log share of employed natives in the country-occupation workforce cell $co$ for a given year. The size of the workforce cell has been set to its 1998 value. Country, occupation, year and country-by-occupation fixed effects are included in all regressions. Additionally, we control for the labor demand shift computed on the sector distribution of occupation, and the weights of sectors are based on national sectoral employment. Standard errors clustered at country-occupation levels are reported in parentheses under coefficient estimates. The instrumental variables in column (2) are the shift-share instrument based on historical settlement patterns and the immigrant-specific demand shift instrument computed according to Eq. (13). In column (3), veteran immigrants are dropped from the estimation and we use only the shift-share instrument. In column (6), we omit the share of EU15 immigrants from the estimation and use the share of non-EU15 immigrants as predicted by our shift-share instrumental variable computed on the set of non-EU15 immigrants' origin countries. First-stage F statistics are the Angrist–Pishcke multivariate $F$-test. The number of observations is lower than in Table 1, due to missing data on the duration of residence or immigrants' origin country (EU15 or not) and missing data on sectoral output used to construct the second instrument.
Statistical significance level: * $p<0.1$, ** $p<0.05$, *** $p<0.01$.

To make sure that this instrument is not correlated with the contemporaneous distribution of natives across occupations due to sectoral labor demand shocks, in all our IV regressions, we control for the (log of) labor demand shift index $(Demand\ shift)_{cot}$. This should control for the direct effect of sectoral labor demand shift on natives' occupational employment. Thus, our identification only relies on the differential impact of these shocks on immigrants because of their historical distribution within occupations across sectors.

Table 3 presents the OLS and IV estimates for the impact of the share of veteran immigrants (more than 10 years of residence) and new immigrants (less than 10 years) in an occupation for men. For both estimation methods, OLS and IV, the share of recent immigrants exerts a small but positive impact on natives' employment rate. The impact of changes in the share of veteran immigrants is much smaller and non-statistically different from zero in IV estimates. Our instruments are less strongly correlated with changes in veteran immigrants than with changes in recent immigrants. This is an expected result as these instruments are more suitable for changes in immigrant flows.[39] In column (3) of Table 3, we omit the share of veteran immigrants and obtain similar results. Overall, a 10% increase in the share of recently arrived immigrants in an occupation increases the employment rate of natives by 0.35%.

According to our model, these results are consistent with the view that new immigrants have lower outside opportunities as compared to natives and that differences in outside option gaps disappear over time as immigrants assimilate with respect to their outside option opportunities.

### *6.2. Immigrants from the EU-15 vs. immigrants from outside the EU-15*

Disparities in outside options among immigrants can also be related to their country of origin. Immigrants from the EU in particular may have a closer outside option to natives than immigrants from outside the EU.[40] If this is the case, the positive impact on natives' employment should be greater when considering non-EU15 immigrants, who are likely to accept lower wages than EU15 immigrants, and who are thus more profitable.

[39] Endogeneity issues for veteran immigrants should be of lesser relevance. Indeed, changes in their share are unlikely to be correlated with contemporaneous labor market shocks.

[40] Immigrants from the EU15 face, on average, lower migration costs and may more freely move back and forth to their home countries. Although limited, several EU15 countries have agreements regarding the international portability of work benefits that workers are entitled to.

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Within occupations, we now distinguish between immigrants from the EU15 and immigrants from outside the EU15. Since some countries do not report this distinction, some observations are lost. Let $shim_{cotEU15}$ be the ratio of immigrants coming from EU15 countries to the 1998 population size of the labor market *co* at time *t*, and let $shim_{cotNOEU15}$ be this ratio for immigrants coming from outside the EU15. The equation to be estimated then becomes[41]

$$\ln y_{cot} = \gamma_0 + \gamma_1 * \ln(shim_{cotNOEU15}) + \gamma_2 * \ln(shim_{cotEU15}) + \alpha_o + \alpha_c + \alpha_t + \alpha_{co} + u_{cot} \tag{14}$$

Under the hypothesis that the outside option of immigrants from outside the EU15 is lower than that of EU15 immigrants, we should find $\gamma_1 > \gamma_2$.

We face the same identification issues as in Eq. (8). We use shift-share instruments based on immigrants' past labor network across occupations and countries. Each shift-share instrument is computed separately for EU15 and non-EU15 immigrants in Eq. (10).

The results are presented in Table 3 for the OLS and the IV. In both estimation methods, the share of non-EU15 immigrants exerts a positive and statistically significant impact: a 10% increase in the share of non-EU15 immigrants in an occupation increases the employment rate of natives by 0.5% in the IV estimates. Instead, immigrants coming from the EU15 exert a comparatively very small impact which becomes non-significantly different from zero once estimated with IV. Considering only changes in non-EU15 immigrants in the last column increases the strength of our instrument but does not alter the result.

### 6.3. *Heterogeneity across host countries*

Up to now, we have shown that immigrants with different durations of residence in a host country and with different origin backgrounds have a different impact on natives' employment rate. We postulate that this differential impact reflects different outside option gaps among immigrants.

We now explore heterogeneity across destination countries with respect to institutional characteristics that may affect the outside option gaps of immigrants with respect to natives. In the model, the outside gap essentially reflects the differential eligibility to unemployment benefits. This is, in particular, related to the fact that immigrants have a lower host country labor market experience and that most countries ask for a minimum work experience in order to become eligible to unemployment benefits. If this is the case, conditional on being unemployed, immigrants' take-up rate of unemployment benefits should be lower than that of natives. We then use the ratio of natives' to immigrants' take-up rates for similar natives and immigrants as a "proxy" for the outside option gap between immigrants and natives. Giuletti et al. (2011) have estimated the conditional unemployment benefit take-up rate ratio for each country in our sample using data from the EU-SILC survey for the year 1999.[42] Using their results, we estimate the following relation:

$$\ln y_{cot} = \beta_0 + \beta_1 * \ln shim_{cot} + \beta_2 * TUR_c * \ln shim_{cot} + \alpha_c + \alpha_o + \alpha_t + \alpha_{co} + \alpha_{ot} + u_{cot} \tag{15}$$

where the unemployment benefit take-up rate ratio, *TUR*, is computed as

$$\frac{\text{Natives' conditional take} - \text{up rate}}{\text{Non} - \text{EU15 immigrants' conditional take} - \text{up rate}}$$

We are only interested in the differential effect of immigrants on natives' employment rate depending on the level of the unemployment benefit take-up rate ratio. We let the direct effect of divergent institutions across countries be absorbed by the country fixed effect included in our regressions. We interpret a higher ratio as a higher outside option of natives relative to immigrants. We then focus on the interacted term between the TUR and the share of immigrants in a particular labor market *co* in year *t*. In this way, we capture the differential impact of immigrants on natives' employment rate across destination countries due to differences in conditional take-up rates between natives and immigrants across countries.

We also create the dummy variable *DummyTUR* that distinguishes countries where the unemployment benefit take-up rate ratio is above 1 (high outside option gap countries) from those where the ratio is below one (low outside option gap countries), and we run similar regressions.[43]

The results are reported in Table 4. Columns (1) and (3) provide the OLS estimates, while in columns (2) and (4) we correct for the endogeneity bias using instrumental variable estimation. As an additional instrument; we simply use the take-up ratio crossed with our previous shift-share instrument. In all cases, we control for country-by-occupation and occupation-by-year fixed effects.[44]

The results provided by columns (1) and (2) show that the positive impact of the share of immigrants on natives' employment rate increases with the conditional TUR. Moreover, as shown in columns (3) and (4), the impact of immigrants on natives' employment is positive and statistically significant only for countries where the conditional immigrants' take-up rate is below that of natives (high outside option gap countries).

[41] We do not have sufficient variation in the data to add a full set of occupation-by-year and country-by-year fixed effects.

[42] The authors estimate a probit model in which the dependent variable is the probability of being an unemployment benefit recipient, conditional on being unemployed. The explanatory variables are gender, age, education and dummies for the country of residence. The authors note that, while immigrants have a higher unconditional probability of receiving unemployment benefits, this is no longer the case once conditioned on unemployment status and socioeconomic characteristics. Instead, if anything, immigrants are less likely to be unemployment benefit recipients.

[43] High outside option gap countries are Germany, Belgium, Portugal, France, Sweden, Ireland, Spain, and the Netherlands. Low outside option gap countries comprise Italy, Denmark, Luxembourg, Norway, Austria and Finland. Data for Greece is not available.

[44] Country-by-year fixed effects are not introduced since we would not have enough variation to separately estimate all the fixed effects (as well as their interactions) and the impact of immigrants across host countries.

**Table 4**
Relevance of divergent institutions across countries.

| Estimation method | Dependent variable: log(employment rate of natives) | | | |
|---|---|---|---|---|
| | OLS | IV | OLS | IV |
| | (1) | (2) | (3) | (4) |
| Independent variables | | | | |
| $\ln(\text{immigrants/workforce } 98)_{cot}$ | $-0.003$ | $-0.015$ | 0.013 | 0.009 |
| | (0.011) | (0.013) | (0.012) | (0.012) |
| $\text{TUR} \cdot \ln(\text{immigrants/workforce } 98)_{cot}$ | 0.019* | 0.033*** | | |
| | (0.010) | (0.009) | | |
| $\text{DummyTUR} \cdot \ln(\text{immigrants/workforce } 98)_{cot}$ | | | 0.009 | 0.014** |
| | | | (0.006) | (0.007) |
| First-Stage $F$ test of excluded instrument | | | | |
| $\ln(\text{immigrants/workforce } 98)_{cot}$ | | 8.55 | | 65.98 |
| $\text{TUR} \cdot \ln(\text{immigrants/workforce } 98)_{cot}$ | | 16.42 | | |
| $\text{DummyTUR} \cdot \ln(\text{immigrants/workforce } 98)_{cot}$ | | | | 54.83 |
| Fixed Effects | | | | |
| country, year and occupation | yes | yes | yes | yes |
| year by occupation | yes | yes | yes | yes |
| country by occupation | yes | yes | yes | yes |
| Observations | 609 | 609 | 609 | 609 |

Units of observation are country-occupation, *co*, in each year *t* from 1998 to 2004. The dependent variable is the log of the natives' employment rate in the country-occupation cell in a given year. The explanatory variables of interest are the log of the share of immigrants in the workforce of a country-occupation cell and its cross term with the unemployment benefit take-up rate ratio between natives and immigrants (columns (1) and (2)), or the same share cross with a dummy equal to 1 for host countries for which this ratio is above one (columns (3) and (4)). The size of the workforce cell has been set to its 1998 value. Under all estimates, we report heteroskedasticity robust standard errors clustered at the country-occupation level in parentheses**. Instrumental variables are the shift-share instruments and their cross terms with the TUR variable.
Statistical significance level: * $p<0.1$, ** $p<0.05$, *** $p<0.01$.

**Table 5**
Baseline parameter values.

| | | | | | | | |
|---|---|---|---|---|---|---|---|
| Job productivity | $h$ | = | 1 | Recruiting cost | $c$ | = | 0.3 $h$ |
| Interest rate | $r$ | = | 0.04 | Exogenous separation rate | $s$ | = | 0.12 |
| Matching elasticity | $\alpha$ | = | 0.5 | Bargaining power | $\eta$ | = | 0.5 |
| Matching efficiency | $m_0$ | = | 0.41 | Outside option natives | $b$ | = | 0.6 |

These results support the view that disparities in outside options, mediated by labor market institutions, are a potentially relevant channel along which immigrants exert their positive externality on natives' employment rate.

## 7. Numerical simulations

Is our model able to reproduce the empirically estimated impact of immigrants on natives' employment rate? We answer this question by numerically simulating our theoretical labor market. More precisely, we simulate the impact of a 1% increase in the share of immigrants on natives' employment rate in a particular labor market for different values of the gap between the outside option of natives and immigrants. For every value of the outside option gap, we provide the elasticity of natives' employment rate. We compare these simulated results with the empirically estimated effect.

The numerical values of the parameters are summarized in Table 5.[45] According to our estimation of Eq. (8), a 1% increase in the immigrants' share within an occupation fosters a rise in the employment rate of natives around 0.05% ($se=0.017$).

[45] The discount factor, the recruiting cost and the bargaining power are taken from Mortensen and Pissarides (1994). The elasticity of the matching process with respect to job seekers ($\alpha$) is taken from Petrongolo and Pissarides (2001). To set the value of the exogenous job destruction rate we use the estimations provided by Davis et al. (1996) for the US and by the OECD (1996) for other Western countries. For simplicity, productivity is normalized to one and, for natives, the outside option of employment is set to 0.6 (see Blanchard and Wolfers (2000)). The scale parameter of the matching function $m_0$ is chosen so that the average unemployment rate of immigrants and natives is around 10–11% for $b_I/b_N<0.6$.

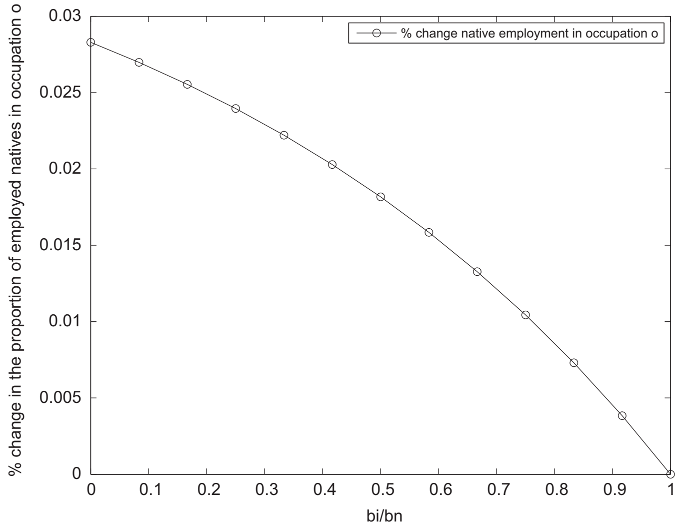


**Fig. 4.** Percentage change in the proportion of employed natives in occupation *o* when the proportion of immigrants in that occupation increases by 1%.

When distinguishing among immigrants by years of residence in the host country, the rise in natives' employment rate equals 0.035% ($se = 0.016$), which is explained by recently arrived immigrants. When immigrants are differentiated by country of origin, the impact rises to 0.053% ($se = 0.015$), all due to non-EU15 immigrants. Finally, when considering institutional heterogeneity across countries, the impact equals 0.014% ($se = 0.007$).

Fig. 4 displays the main results. As observed, for a sufficiently low, though plausible, outside option gap of immigrants as compared to natives ($b_I/b_N < 0.7$), the simulated elasticity of natives' employment rate to an immigrant labor supply shock falls within the confidence interval of our empirically estimated effects (IV estimates). More precisely, when the immigrant outside option represents between 50 and 60% of that of natives, the impact on natives' employment rate equals 0.014%, which is our median interval point estimates when distinguishing countries by the generosity of their institutions. While we lack a precise measure of this gap, our reading of the literature on immigrant-native wage discrimination suggests that this gap is a reasonable value (see for example the paper by Nanos and Schluter (2014)).

Overall, we view results from these simulations as further evidence that disparities in outside options between natives and immigrants provide a plausible mechanism through which immigrants may improve employment opportunities for natives.

## 8. Conclusion

The increasing contribution of immigrants to the labor force is among the most important contemporaneous labor supply shocks facing European labor markets. To date, most of the literature has discussed the labor market consequences of such shocks using a standard neoclassical labor-supply labor-demand framework. However, this approach does not allow to introduce important differences in non-productive assets between immigrants and natives.

We have shown in this paper that, once introduced into a frictional labor market, differences in host-country-specific assets between immigrants and natives can reverse the conclusions reached by the standard model: in the short run, immigrants improve the employment prospects of competing native workers. Thus, instead of crowding out natives, immigrants may instead crowd in natives in sectors and occupations to which they contribute.

The employment creation effect has been found more important for new immigrants, for immigrants from non-EU15 countries, and for countries that display large differences in the unemployment benefit take-up rate between similar immigrants and natives. Overall, these results highlight that immigrants may lack host-country-specific assets, which explains their positive impact on natives' employment.

Some implications are worth pursuing further. First, regarding the design of an optimal immigration policy. On the one hand, recent research indicates that skilled immigrants may crowd out natives in skilled jobs (see Borjas, 2009). On the other hand, it has been argued that unskilled immigrants may improve incentives for natives to acquire human capital by raising the skill premium (see Hunt, 2011). By contrast, our conclusions suggest that host countries with a more selective immigration policy could improve the employment rate of skilled workers and, at the same time, boost incentives for

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natives to acquire human capital. A welfare analysis of such a policy is a natural extension of the model proposed in this paper.

Second, on the empirical side, we highlight the importance of distinguishing immigrants according to their origin country or duration of residence. More generally, our more realistic approach to the functioning of the labor market stresses the importance of considering any heterogeneity between immigrants and natives that would affect their relative bargaining position with respect to employers.

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## Appendix A. Supplementary data

Supplementary data associated with this paper can be found in the online version at http://dx.doi.org/10.1016/j.euroecorev.2015.10.001.

## Appendix B. Estimation results

Table A1

**Table A1**
OLS and 2SLS estimates of the effect of immigration on natives' employment rate within country-occupation (males).

| Dependent variable: | log(employment rate of natives in an country-occupation cell) | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| Estimation method: | OLS | | | | 2SLS | | | |
| | | | | | IV | IV | IV | IV |
| Independent Variable | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
| ln(immigrants/workforce 98)oct | 0.016*** | 0.012*** | 0.016*** | 0.016*** | 0.027** | 0.018** | 0.052*** | 0.047*** |
| | (0.004) | (0.003) | (0.005) | (0.004) | (0.012) | (0.008) | (0.019) | (0.017) |
| Labor demand shift index | | | | 0.136*** | | | | 0.134*** |
| | | | | (0.039) | | | | (0.041) |
| Additional Fixed Effects | | | | | | | | |
| Country by occupation | YES | YES | YES | YES | YES | YES | YES | YES |
| Occupation by year | NO | YES | YES | YES | NO | YES | YES | YES |
| Country by year | NO | NO | YES | YES | NO | NO | YES | YES |
| First-Stage F on excluded instrument (Kleibergen-Paap rank Wald F statistic) | | | | | 18.80 | 17.86 | 11.41 | 10.54 |
| Observations | 654 | 654 | 654 | 654 | 654 | 654 | 654 | 654 |

*Note*: Units of observation are country, occupation and year. Country, year and occupation fixed effects are included in all regressions. The labor demand shift is based on the sector distribution of occupation with the weights of sectors based on national sectoral employment. The instrumental variable is the shift-share instrument based on historical settlement patterns of immigrants (see text for details). Standard errors clustered at country-occupation levels are reported in parentheses. Regressions are weighted by the total number of natives in a country-occupation cell in 1998, using an alternative weight does not affect the result.
Statistical significance level: ** $p < 0.05$, *** $p < 0.0$.

## Appendix C. A two-sector model

### C.1. The matching process

We consider a labor market represented by occupation $o$. This occupation covers two sectors: A and B. We assume that productivity in sector A is higher, so wages earned by people employed in sector A are also higher. We assume that unemployed people in sector A have a per period probability $\lambda$ of being depreciated to sector B. We allow workers employed in sector B to do on-the-job search in sector A where wages are higher. Since we are considering a single occupation, these flows between sectors within an occupation are perfectly reasonable.

Let us denote as $t=A,B$ the two existing sectors, $j=N,I$ native and immigrant workers, $v^t$ the number of vacancies in sector $t$, $u_j^t$ the number of job seekers, $n_j^t$ the number of employed individuals and $eo_j$ the on-the-job search effort. The matching functions can thus be written as: $M^A=m^A(v^A,u_N^A+u_I^A+eo_N\cdot n_N^B+eo_I\cdot n_I^B)$ and $M^B=m^B(v^B,u_N^B+u_I^B+\lambda u_N^A+\lambda u_I^A)$. We assume a standard homogeneous matching function of the form $M^A=m_0(v^A)^{1/2}(u_N^A+u_I^A+eo_N\cdot n_N^B+eo_I\cdot n_I^B)^{1/2}$ and $M^B=m_0(v^B)^{1/2}(u_N^B+u_I^B+\lambda u_N^A+\lambda u_I^A)^{1/2}$.

Labor market opportunities are described by the market tightness variables $\theta^A=v^A/(u_N^A+u_I^A+eo_N\cdot n_N^B+eo_I\cdot n_I^B)$ and $\theta^B=v^B/(u_N^B+u_I^B+\lambda u_N^A+\lambda u_I^A)$. The probability of filling an empty vacancy equals $q(\theta^t)=M^t/v^t$. The probability of finding a job is given by $p(\theta^A)=M^A/(u_N^A+u_I^A+eo_N\cdot n_N^B+eo_I\cdot n_I^B)$ and $p(\theta^B)=M^B/(u_N^B+u_I^B+\lambda u_N^A+\lambda u_I^A)$. In sector A, a vacancy is filled by a native worker with probability $q(\theta^A)\frac{u_N^A+eo_N\cdot n_N^B}{u_N^A+u_I^A+eo_N\cdot n_N^B+eo_I\cdot n_I^B}$ and by an immigrant with probability $q(\theta^A)\frac{u_I^A+eo_I\cdot n_I^B}{u_N^A+u_I^A+eo_N\cdot n_N^B+eo_I\cdot n_I^B}$. In sector B, the probability equals $q(\theta^B)\frac{u_N^B}{u_N^B+u_I^B+\lambda u_N^A+\lambda u_I^A}$ for natives and $q(\theta^B)\frac{u_I^B}{u_N^B+u_I^B+\lambda u_N^A+\lambda u_I^A}$ for immigrants.

### C.2. The agents' behavior

#### C.2.1. Workers

Employed workers coming from unemployment are paid $w_j^t$, whereas workers in sector A who were previously employed in sector B earn $w_j^{AB}$, for $j=N,I$. Jobs are destroyed at the exogenous probability $s$. Workers employed in sector B have a probability $eo_j\cdot p(\theta^A)$ of finding a job in sector A, but they bear a disutility cost linked to the search effort equal to $\tau(eo_j)=\phi_0\cdot eo_j^{\phi_1}$, where $\phi_1>1$ so that $\tau'(eo_j)>0$ and $\tau''(eo_j)>0$.

The asset values of employment in sectors A and B and in sector A but for someone coming from B, are respectively given by

$$rE_j^A=w_j^A+s(U_j^A-E_j^A) \tag{16}$$

$$rE_j^B=w_j^B-\tau(eo_j)+s(U_j^B-E_j^B)+eo_j\cdot p(\theta^A)(E_j^{AB}-E_j^B) \tag{17}$$

$$rE_j^{AB}=w_j^{AB}+s(U_j^A-E_j^{AB}) \tag{18}$$

where $U_j^t$ stands for the asset value of unemployment.

The asset values of unemployment are written as follows:

$$rU_j^A=b_j+p(\theta^A)(E_j^A-U_j^A)+\lambda(U_j^B-U_j^A) \tag{19}$$

$$rU_j^B=b_j+p(\theta^B)(E_j^B-U_j^B) \tag{20}$$

where $b_N>b_I$.

An individual employed in sector B searches on the job until all possible rents are exhausted, that is, until the marginal cost of an additional unit of search effort equals the marginal expected benefit from the on-the-job search:

$$\tau'(eo_j)=p(\theta^A)(E_j^{AB}-E_j^B) \tag{21}$$

Because $\tau''(eo_j)>0$, we deduce that an increase in $p(\theta^A)$ should drive up the on-the-job-search effort. Intuitively, if employment opportunities improve in sector A, while wages in both sectors remain unchanged, individuals will search more intensively in sector A.

#### C.2.2. Firms

The value of an empty vacancy in sector B is given by

$$rV^B=-\gamma+q(\theta^B)(\overline{J^B}-V^B) \tag{22}$$

where $\overline{J^B}$ represents the average value of a filled vacancy. The value of a filled vacancy is defined by the instantaneous profit $h^B-w_j^B$ associated with the job (productivity minus the wage) plus the expected loss if the vacancy becomes empty, either because of an exogenous job destruction shock or because the worker finds a position in sector A:

$$rJ_N^B=h^B-w_N^B+s(V^B-J_N^B)+eo_N\cdot p(\theta^A)(V^B-J_N^B) \tag{23}$$

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$$rJ_I^B = h^B - w_I^B + s(V^B - J_I^B) + eo_I \cdot p(\theta^A)(V^B - J_I^B) \tag{24}$$

The average value of a filled vacancy in sector B results from the weighted average $\overline{J^B} = \omega_1^B J_I^B + (1-\omega_1^B) J_N^B$, where $\omega_1^B = \frac{u_I^B}{(u_N^B + u_I^B + \lambda u_N^A + \lambda u_I^A)}$.

In sector A, the vacancy may be filled by a native worker (unemployed or coming from sector B) or by an immigrant (unemployed or coming from sector B). The decision concerning the number of vacancies to open is then also based on the average expected profit. We denote $V^A$ the value of an empty vacancy and $J_N^A$, $J_N^{AB}$, $J_I^{AB}$ and $J_I^A$ the values of a position filled, respectively, by a native worker previously unemployed, a native worker previously employed in sector B, an immigrant worker previously employed in sector B, and an immigrant worker previously unemployed. These values are given by

$$rV^A = -\gamma + q(\theta^A)(\overline{J^A} - V^A) \tag{25}$$

$$= -\gamma + q\left(\theta^A\right)\left(\frac{u_N^A}{u_N^A + u_I^A + eo_I \cdot n^B + eo_I \cdot n_I^B} J_N^A + \frac{eo_N \cdot n_N^B}{u_N^A + u_I^A + eo_N \cdot n_N^B + eo_I \cdot n_I^B} J_N^{AB} + \frac{eo_I \cdot n_I^B}{u_N^A + u_I^A + eo_N \cdot n_N^B + eo_I \cdot n_I^B} J_I^{AB} \right.$$
$$\left. + \frac{u_I^A}{u_N^z + u_I^z + eo_N \cdot n_N^B + eo_I \cdot n_I^B} J_I^A - V^B \right)$$

where

$$rJ_N^A = h^A - w_N^A + s(V^A - J_N^A) \tag{26}$$

$$rJ_N^{AB} = h^A - w_N^{AB} + s(V^A - J_N^{AB}) \tag{27}$$

$$rJ_I^{AB} = h^A - w_I^{AB} + s(V^A - J_I^{AB}) \tag{28}$$

$$rJ_I^A = h^A - w_I^A + s(V^A - J_I^A) \tag{29}$$

where $h^A$ corresponds to the productivity of the job and $w_N^A$, $w_N^{AB}$, $w_I^{AB}$ and $w_I^A$ stand, respectively, for the wage of a native previously unemployed, for the wage of a native coming from sector B, for the wage of an immigrant coming from sector B, and for the wage of an immigrant previously unemployed. We denote as $\omega_1^A = \frac{u_I^A}{u_N^A + u_I^A + eo \cdot n_N^B + eo \cdot n_I^B}$ the proportion of immigrant job-seekers in sector A who were previously unemployed, $\omega_2^A = \frac{eo \cdot n_I^B}{u_N^A + u_I^A + eo \cdot n_N^B + eo \cdot n_I^B}$ the proportion of immigrants in sector A who were previously employed in sector B, the share of native job seekers coming from sector B equals $\omega_3^A = \frac{eo \cdot n_N^B}{u_N^A + u_I^A + eo \cdot n_N^B + eo \cdot n_I^B}$ and the proportion of native job seekers previously unemployed is given by $\left(1 - \omega_1^A - \omega_2^A - \omega_3^A\right) = \frac{u_N^A}{u_N^A + u_I^A + eo \cdot n_N^B + eo \cdot n_I^B}$.

Firms open vacancies until no more profit can be obtained so that, at equilibrium, the free-entry condition $V^t = 0$ applies, i.e.

$$\frac{\gamma}{q(\theta^A)} = \overline{J^A} \quad \text{and} \quad \frac{\gamma}{q(\theta^B)} = \overline{J^B} \tag{30}$$

where

$$\overline{J^A} = \frac{h^A - \omega_1 w_I^A - \omega_2 w_N^{AB} - \omega_3 w_I^{AB} - (1 - \omega_1 - \omega_2 - \omega_3) w_N^A}{r+s} \tag{31}$$

$$\overline{J^B} = \omega_1^B J_I^B + (1 - \omega_1^B) J_N^B = \omega_1^B \frac{h^B - w_I^B}{r + s + eo_I \cdot p(\theta^A)} + (1 - \omega_1^B) \frac{h^B - w_N^B}{r + s + eo_N \cdot p(\theta^A)} \tag{32}$$

### C.3. Wages

We consider a wage determination process in the style of Hall and Milgrom (2008) so that the outcome of the symmetric alternating-offers game is

$$w_N^t = \eta h^t + (1-\eta) b_N \tag{33}$$

$$w_I^t = \eta h^t + (1-\eta) b_I^t \tag{34}$$

$$w_N^{AB} = \eta h^A + (1-\eta) w_N^B \tag{35}$$

$$w_I^{AB} = \eta h^A + (1-\eta) w_I^B \tag{36}$$

where $\eta$ can be interpreted as the bargaining power of each party and is set to 1/2.

### C.4. Employment opportunities

Employment opportunities are measured by the labor market tightness which is determined by the free-entry condition (30). Combining this equation with (31) and (32) yields

$$\frac{\gamma}{q(\theta^A)}=\overline{J}^A=\frac{h^A-\overline{w^A}}{r+s} \quad \text{and} \quad \frac{\gamma}{q(\theta^B)}=\overline{J^B}=\omega_1^B\frac{h^B-w_I^B}{r+s+eo_I\cdot p(\theta^A)}+(1-\omega_1^B)\frac{h^B-w_N^B}{r+s+eo_N\cdot p(\theta^A)} \tag{37}$$

Since $\frac{\gamma}{q(\theta^t)}=\frac{\gamma}{m_0}(\theta^t)^{1/2}$ for $t=A,B$, we find

$$\theta^A=\left(\frac{m_0(h^A-\overline{w^A})}{\gamma(r+s)}\right)^2 \quad \text{and} \quad \theta^B=\left(\frac{m_0}{\gamma}\omega_1^B\frac{h^B-w_I^B}{r+s+eo_I\cdot p(\theta^A)}+\frac{m_0}{\gamma}(1-\omega_1^B)\frac{h^B-w_N^B}{r+s+eo_N\cdot p(\theta^A)}\right)^2 \tag{38}$$

where $\overline{w^A}=\omega_1^Aw_I^A+\omega_2^Aw_I^{AB}+\omega_3^Aw_N^{AB}+(1-\omega_1^A-\omega_2^A-\omega_3^A)w_N^A$.

### C.5. The unemployment rates

At the steady state, outflows from one sector must equal inflows. Moreover, inside every sector, entries to unemployment must equal exits. Outflows from sector A (inflows toward sector B) equal $\lambda(u_I^A+u_N^A)$, while inflows to A (outflows from B) correspond to $(n_I^Beo_Ip(\theta^A)+n_N^Beo_Np(\theta^A))$. At equilibrium

$$\lambda(u_I^A+u_N^A)=(n_I^Beo_Ip(\theta^A)+n_N^Beo_Np(\theta^A))$$

Inside sector A, outflows from unemployment equal $\lambda u^A+p(\theta^A)u^A$ where $u^A=(u_I^A+u_N^A)$. Inflows to unemployment equal $sn^A$, where $P^A=n^A+u^A$. Equalizing inflows and outflows from unemployment in sector A gives us the aggregate unemployment in sector A:

$$u^A=\frac{sP^A}{\lambda+p(\theta^A)+s}$$

Distinguishing between immigrants and natives, we can compute the unemployment and employment rates in a similar manner:

$$u_N^A=\frac{sP_N^A}{\lambda+p(\theta^A)+s} \quad u_I^A=\frac{sP_I^A}{\lambda+p(\theta^A)+s}$$
$$n_N^A=\frac{p(\theta^A)eo_Nn_N^B+p(\theta^A)P_N^A}{s+p(\theta^A)} \quad n_I^A=\frac{p(\theta^A)eo_In_I^B+p(\theta^A)P_I^A}{s+p(\theta^A)}$$

Applying the same reasoning to sector B

$$u^B=\frac{sP^B+\lambda u^A}{p(\theta^B)+s}, \quad u_N^B=\frac{sP_N^B+\lambda u_N^A}{p(\theta^B)+s} \quad \text{and} \quad u_I^B=\frac{sP_I^B+\lambda u_I^A}{p(\theta^B)+s}$$
$$n_N^B=\frac{p(\theta^B)P_N^B}{s+p(\theta^A)eo_N+p(\theta^B)} \quad n_I^B=\frac{p(\theta^B)P_I^B}{s+p(\theta^A)eo_I+p(\theta^B)}$$

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