WORKING PAPER

N° 2022-12

IMMIGRATION AND LABOUR MARKET FLOWS

ANDRI CHASSAMBOULLI, IDRISS FONTAINE,
ISMAEL GALVEZ-INIESTA

www.tepp.eu

TEPP – Theory and Evaluation of Public Policies - FR CNRS 2042

Immigration and Labour Market Flows

Andri Chassamboulli [*] University of Cyprus

Idriss Fontaine [†] Université de La Réunion

Ismael Gálvez-Iniesta [‡] Universitat de les Illes Balears

Pedro Gomes [§] Birkbeck, University of London

April 1, 2022

Abstract

We examine the size and cyclicality of job-finding and job-separation rates for immigrants and natives in France, Spain and the U.S., for the period between 2003 and 2018. We decompose cyclical fluctuations in the unemployment rate of immigrants and natives into contributions attributable to inflows and outflows to and from employment and inactivity. Most facts on the differences of the relative importance and cyclicality of transition rates between immigrants and natives are not common across the three countries, suggesting that the type of migration matters. Using a VAR model we find that an inflow of foreign workers has a weak and mostly non-significant effect on the job-finding and job-separation rates of both immigrants and natives.

JEL classification: J15 J60

Keywords: Labour market flows, immigration, job-finding rate, job-separation rate, VAR.

[*] Department of Economics, University of Cyprus, CY-1678 Nicosia, Cyprus.; Email: andricha@ucy.ac.cy.

[†] Department of Economics (CEMOI), Université de La Réunion; E-mail: idriss-fontaine@univ-reunion.fr or idriss.fontaine974@gmail.com.

[‡] Department of Applied Economics, Universitat de les Illes Balears, Ctra. Valldemossa, km 7.5, 07122 Palma de Mallorca, Spain; E-mail: i.galvez@uib.es.

[§] Department of Economics, Mathematics and Statistics, Birkbeck, University of London, Malet Street, WC1E 7HX London, United Kingdom, E-mail: p.gomes@bbk.ac.uk

1 Introduction

We document facts about unemployment rates and labour market transitions of native and immigrant workers, and examine their differences. The question of how immigrants' labour market performance in the host country compares to that of natives has received considerable research attention, described below. However, most studies focus on particular countries and specific labour market outcomes, such unemployment and non-employment rates, or the incidence of job loss. This is the first study to provide a comprehensive analysis of labour market transitions rates of immigrants and natives using micro data for three different countries: France, Spain, and the United States. These countries have different labour market institutions, types of migration and legislation. Hence, by using comparable data, we can investigate whether there are common facts about differences between immigrants and natives, or whether they are country specific and cannot be generalized.

We use data from each country's representative labour market survey to construct worker flows between employment, unemployment and inactivity, both unconditional and conditional on observable characteristics. We then use our estimates of transition rates to decompose the cyclical fluctuations in the unemployment rates of immigrants and natives into the contributions of each of the flow hazards. Finally, we employ a VAR model to analyse the dynamic responses of native and immigrant transition rates after an migration shock.

The three countries differ on how their immigrants' labour market transitions compare to those of natives in terms of size, relative importance, and cyclicality. In the U.S., the unemployment rates of immigrants and natives exhibit a remarkably similar behavior in terms of size and business cycle fluctuations. The two European countries, by contrast, share the common feature of a higher unemployment rate for immigrants than natives. This reflects in part the fact that the composition of immigrants in the U.S. is more similar to that of natives than in the two European countries. In the latter, immigrants tend to be less educated and more concentrated in part-time jobs and the construction sector. In Spain also, the share of immigrants in temporary jobs is significantly larger than that of natives. Nevertheless, we find that the differences between immigrants and natives remain even after controlling for observable characteristics. Consistent with the unemployment rate gap, in Spain and France, the conditional job-finding rate is lower and job-separation rate is higher for immigrants than natives, while in the U.S. the job-finding and separation rates of immigrants and natives are very similar. The higher incidence of temporary jobs for immigrants in Spain can explain only part of the gap in separation rates, while in France, the separation rate gap remains almost intact with the inclusion of controls.

While Spain and France share the common feature of higher unemployment rate for immigrants, they display an important difference regarding their cyclical response. The unemployment rate gap in France is remarkably stable over the period, while in Spain the gap has widen after the Great Recession in the mid-2008. Likewise, the job-finding and separation-rate gaps have widened after the onset of the Great Recession in Spain, but remained stable in France. In Spain, cyclical fluctuations have more influence upon immigrants than natives, and shape the unemployment rate gap between them. However, the same conclusion cannot be reached for France, where the gap seems to reflect factors unrelated to observable characteristics or business cycle effects. Our decomposition of the cyclical fluctuations in the unemployment rate reveals that in Spain the job-separation rate is more relevant for explaining the cyclical behavior of unemployment for immigrants than for natives. In France, on the other hand, flows within participation (from employment to unemployment and vice versa) are more relevant, suggesting that immigrants are more attached to the labour market than natives.

In the rest of the paper we estimate a Structural Vector Autoregressive (VAR) model to study the effects of an inflow of foreign workers on the job-finding and job-separation rates of both immigrants and natives. We find that an increase in immigration in the U.S. decreases only marginally both the probability of separation and the probability of finding a job for natives. In contrast, in France and Spain, we find non-significant effects of immigration on natives' labour market flows. While most facts on the differences of transition rates between immigrants and natives are not common across countries, suggesting that the type of migration matters, the results from the VAR suggest that, whatever the country, an inflow of immigrants has negligible effects on the labour market flows and stocks of immigrants and natives.

Various studies show that immigrants experience labour market disadvantages relative to natives. Bratsberg et al. (2018) show that for immigrant workers in the Norwegian private sector, the probability of job loss is significantly larger and the consequences of a job loss are more severe, than for natives. Blume et al. (2009) find that immigrants in Denmark are more likely to be nonemployed or to use self-employment as a last resort to avoid non-employment. While most studies find that immigrants are more likely to lose their job compared to natives, results concerning the job-finding probability are mixed. Botrić (2018) analyses the differences in transitions from unemployment to employment and vice versa between immigrants and natives in European economies. The relative transitions from unemployment to employment depend on the country and year, but in most countries, employed immigrants are more likely to lose a job than natives. Albert (2021), on the other hand, finds for the U.S., that immigrants and especially illegal immigrants have lower unemployment rates and the difference is driven by higher job-finding rates.

Some other studies focus on the differences in the cyclical responses of immigrants' and natives' labour market outcomes. Dustmann et al. (2010) analyse differences in the cyclical pattern of employment and wages for Germany and the UK. They find larger unemployment responses to economic shocks for immigrants relative to natives within the same skill group, but little evidence for differential wage responses to economic shocks. Bratsberg et al. (2014) focus on the Eastern-European immigrants to Norway and find a larger increase in the fraction claiming unemployment insurance benefits during the "great recession" among immigrants than natives. Likewise, the findings of Carrasco and García-Pérez (2015) for Spain suggest that immigrants are more sensitive to changes in economic conditions in terms of both unemployment and employment exit rates.

From a methodological point of view, we follow a well-established literature estimating and describing worker gross flows, that started with Blanchard et al. (1990). This methodology has been applied to understand particular aspects of the labour market. It has been used to study differences between: full-time and part-time employment (Borowczyk-Martins and Lalé, 2020), permanent and temporary jobs (Silva and Vázquez-Grenno, 2013), men and women (Baussola and Mussida, 2014) or public- and private-sector employment (Fontaine et al., 2020). It has also been used to study the role of: the participation margin (Elsby et al., 2015), labour force attachment (Gomes, 2012), on-the-job search and job-to-job transitions (Fujita, 2010), worker reallocation across occupations and industries (Carrillo-Tudela et al., 2016), conditional transition probabilities (Gomes, 2015), or types of occupations (Charlot et al., 2019). Despite looking at worker flows from different angles, all the papers in this exhaustive list have ignored the duality between immigrants and natives.

Finally, we hope to contribute and strengthen a recent literature building models with search and matching frictions to understand the labour market outcomes and effects of migrants. Examples include Chassamboulli and Palivos (2014), Chassamboulli and Peri (2015, 2020), Liu et al. (2017), Battisti et al. (2018), Chassamboulli and Liu (2020), Albert (2021). At the heart of these models lies the transition rates across labour market states, so we hope that, by describing these facts about transition rates for immigrants and natives, we can inform their design and calibration.

2 Immigration in France, Spain and U.S.

2.1 Background

Unlike most European countries, France has a large experience as immigration host country. During the early years of the 20th century, it became a destination country for immigrants, attracted by large labour market needs (Barou, 2014). Most of the immigrants came from neighbouring countries, such as Italy, Spain, Belgium, and Switzerland, and they were quickly assimilated into the national population. In the 1920s, France ranked second, just after the U.S. as the country with the highest share of immigrants, reaching 7 per cent of total population. The inflow of immigrants increased substantially after the Second World War, coming from Italy, Spain, Portugal and North Africa, notably Algeria and former protectorates of Morocco and Tunisia. Immigration from Algeria boomed after the Second World War until 1958 and the Algerian civil war. Immigration from Morocco and Tunisia took place later, during the 1970s (Algan et al., 2012). This increase in immigration is mainly explained by the post-war industrial expansion, that required a large number of low-skilled workers. With the beginning of the economic downturn in 1974, inflows slowed down and the French government made migration polities stricter, restricting immigration from its former colonies (Cooper, 2018). The share of foreign-born remained stable around 8 per cent throughout the 70s to the early 2000s. During this period, the most novel feature were the inflows from the Sub-Saharan Africa (Cameroon, Ivory Coast, Mali and Senegal). Since 2000 the share has increased, reaching 12 per cent in 2015. The most recent wave of immigration is from Eastern Europe and Turkey. Still, almost half of all immigrants entering France in 2012 were born in Europe, against one-third ten years previously.

Immigration to Spain is a recent phenomenon. Throughout most of the 20th century Spain was a country of emigration, mainly to Europe (France, Germany and Switzerland), as described in Izquierdo et al. (2015). From the late 90's, Spain experienced the largest inflow of immigrants among all developed countries. From 1998 to 2008, on average immigrant inflows made up 1.1 per cent ot the total population per year. As a consequence, the number of immigrants in the labour force went up from 0.27 millions (1.6 per cent of the labour force) in 1998, to 3.4 millions (15 per cent of the labour force) in 2010. This boom is explained by a combination of factors: the Latin American crisis (Bertoli et al., 2011), Spanish migration policies (Bertoli and Moraga, 2013), the Eastern European expansion of the European Union and the Spanish economic expansion. The Great Recession halted this trend, triggering a sudden drop of immigrant inflows from 2008 to 2016. The crisis also brought a steady increase in immigrant outflows (Prieto-Rosas et al., 2018) which lead to a drop in the share of immigrants among the labour force by 3 percentage points to 12 per cent at the end of 2016. Immigration legislation in Spain (Law 4/2000 on January of 2000) established a general principle of equality between foreign and natives. Among other aspects, it gave to foreigners the same rights as natives regarding access to the legal system and ideological/religious freedom. It also allowed full access to medical assistance and education to those registered in the Municipal Registry.

The United States, in contrast with Spain, has long been one of the main immigrant host countries. Mass immigration to the U.S., mainly from the European periphery, began in 1850 and continued through 1920. During this period the share of foreign-born in the U.S. population rose from 10 to 14 per cent. The introduction of strict immigration quotas in 1921 slowed down the immigrant inflows, but only temporarily, as the relaxation of the quotas in the 70s allowed again mass immigration into the U.S.. Since 1960 the number of foreign-born in the U.S. has doubled every 30 years. It increased from around 10 million to 20 million between 1960 and 1990 and reached 45 million in 2018.[1] While the 19th century U.S. immigrants were mainly from Europe, the majority of the current immigrant stock is from Latin America (50 per cent) and Asia (30 per cent) (Abramitzky and Boustan, 2017).

The two main channels of legal immigrant entry into the U.S. are: (i) family-unification, which allows the entry of U.S. citizens' and legal permanent residents' (LPRs) immediate family members; and (ii) employment-based entry, which allows the entry of individuals whose skills are valuable to the U.S. economy.[2] Unlike family-based entries, immigrants entering on employment visas must

[1] Pew Research Center report, August 20, 2020, "Facts on U.S. immigrants, 2018"

[2] Entry is also possible through the refugee or asylum-seeker status, but at a much smaller scale.

meet certain skill/qualification requirements. Moreover, most employment immigrants are initially admitted on temporary work permits, and may transition to permanent residence status subsequently, whereas immigrants admitted on family visas can stay and work in country indefinitely (i.e. family visas are permanent).

A distinct feature of the U.S. immigration system is the possibility of the so called "chain immigration effects" as it allows for networks and family ties to create future immigration opportunities and inflows into the country. A foreigner can apply for a work permit to enter the U.S. only if he has already been offered a job. Such job offers are made available to potential skilled migrants through their network of co-ethnics who are already legally employed. Employment immigrants can then generate opportunities for new immigrant entries through family ties. As the network of incumbent legal-permanent immigrants expands entry of new immigrants through either family ties or employment becomes more likely. Although there is a limit on how many permanent immigrant visas are granted each year across the various visa categories, there is no limit on the annual admission of U.S. citizens' immediate relatives. Thus, small changes in U.S. immigration laws or quotas can have substantial long-run equilibrium effects, through networks and family linkage.

Besides legal immigrants, a substantial fraction of the U.S. foreign-born are illegal/unauthorized immigrants. Over the last few decades, the number of illegal immigrants has grown rapidly from 3 million to over 11 million. They account for about 30 per cent of the total immigrant stock and about 50 per cent of the total immigrant stock in working age. Illegal immigrants are predominately unskilled, coming mainly from Latin America and especially Mexico, while Asia (mostly China and India) is the most relevant source region for high-skilled immigrants who are predominately legal, either through the family or the employment channel.

2.2 Data and samples

We use data from each country's representative labour market survey: the French Labour Force Survey, the Spanish Labour Force Survey and the U.S. Current Population Survey. The French and Spanish Labour Force Survey are both conducted quarterly through a rotating panel: in each quarter, one sixth of the sample ("oldest wave") is replaced by a new wave of entrants. The longitudinal structure of the surveys allow us to match observations from two consecutive surveys. Due to the structure of the database, we can track each individual for five consecutive quarters (one year and a half). To solve for panel attrition and non-responses, we reweighed each longitudinal sample by a method similar to the one proposed by Lundström and Särndal (1999). An exhaustive description of the method can be found in Fontaine (2016). For the Spanish data, we restrict the analysis to years after 2005, as the database began to report information about the nationality of the respondent from that year on. In contrast to the European surveys, the CPS provides information at a monthly frequency. It follows household for four consecutive months, omitting them for eight months and the interviewing them again for another four months. We follow Shimer (2012) to compute monthly transition rates between labour market states. We restrict to sample to 2003-2018 for comparison with the other two datasets. One of the advantages of using labour force survey data is that labour market states are defined according to the International Labour Organization (ILO) definition, which enables us to compare labour market characteristics of the three countries.

There is only one methodological difference between the three countries. For the U.S. and France, immigrant status is defined by country of birth. Unfortunately, the longitudinal version of the Spanish Labour Force Survey only reports respondent' citizenship and not country of birth. We therefore define immigrants as individuals with foreign citizenship, as in Dustmann et al. (2010). The cross-sectional version of the Spanish dataset does report both respondents' country of birth and citizenship. Foreign-born residents without Spanish nationality (i.e. our definition of immigrants) account for the large majority of all foreign-born residents in Spain, 77 per cent, with the remaining 23 per cent being foreign-born residents with Spanish nationality. The average unemployment rate of the two groups is 25.3 per cent and 20.9 cent, respectively. We should keep in mind that some of the differences in the labour market flows of immigrants in Spain relative to France and the U.S. might be driven by this definition discrepancy.

A legitimate question, particularly for the U.S., is about the true nature of our migration variable. Does it capture legal or illegal immigration? The unit sample of these surveys is the house and not the household. Consequently, to be selected in the final sample individuals should be enrolled in the municipality registers, regardless of the legal status. Given this, we think that it is more likely that the share of immigrants captures both legal and illegal immigration.

2.3 Descriptive Statistics

Table 1 displays descriptive statistics of the composition of immigrant and native populations in France, Spain and the U.S.. Over the period, immigrants account for a large share of the sample: 9 per cent in France, 11.7 per cent in Spain, and 15.4 per cent in the U.S.. The immigrant composition exhibits some common features in the three countries. They are more concentrated among the lowest educational group and the share of employment in the construction sector is higher than for natives. Immigrants also face a higher rate of temporary contracts in the European countries, 4 and 16 percentage points higher than for natives, in France and Spain, respectively.

However, different migration historical roots and nationality composition (Figure 1) also create relevant differences. While immigrants are significantly younger than natives in Spain, the opposite is observed in U.S. and France. In the two European countries immigrants work more in part-time jobs, about 18 per cent, and this fraction decreases to 13.7 per cent in the U.S.. While labour force participation rates of immigrants and natives are almost identical in the U.S., the participation rate of immigrants is 5 percentage points lower in France and 7 percentage points higher in Spain, when compared to natives. Also, education differences are more pronounced in the two European countries. For example, the share of tertiary educated workers is substantially lower for immigrants than natives in Spain and France, while in the U.S. the share of those with tertiary education is similar. To summarise, the compositions of the immigrant populations are heterogeneous across countries.

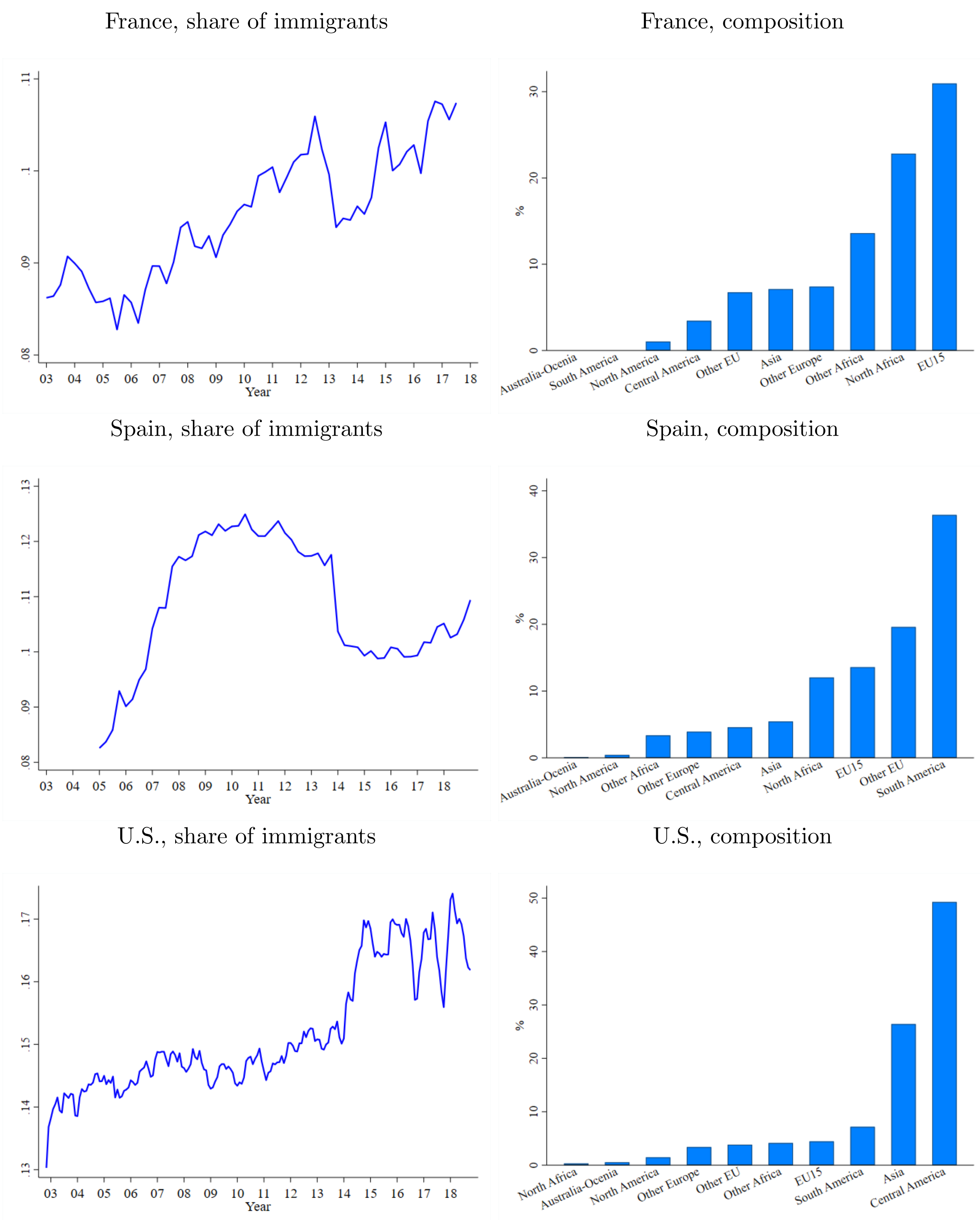
Figure 2 displays the evolution of the unemployment rates of natives and immigrants. For both Spain and France, for the period considered, the unemployment is higher for immigrants than for natives. In contrast, the U.S. exhibits few differences in the behaviour of the unemployment rate by nationality. Despite sharing the common feature of exhibiting a higher unemployment rate for immigrants, Spain and France display an important distinction, mainly the cyclical pattern of their unemployment rates. In particular, while in France the unemployment-gap between immigrants and natives is stable over the period (around 8 percentage points), in Spain it skyrocketed after the Great Recession that took place in the mid-2008. In particular, with the crisis, the unemployment rate gap raised from 6 percentage points in 2008 to more than 13 percentage points in 2013.

Table 1: Descriptive Statistics

	France		Spain		U.S.	
Immigrant share	9.00%		11.67%		15.40%	
	Natives	Immigrants	Natives	Immigrants	Natives	Immigrants
Male	52.11	55.09	56.08	53.35	52.25	59.41
Age						
16-19	2.03	0.61	1.16	1.80	4.37	1.63
20-24	8.41	4.15	6.10	7.63	11.06	7.32
25-29	11.88	9.24	11.04	14.54	11.80	11.58
30-34	12.50	12.54	13.43	19.87	11.03	13.75
35-39	13.16	14.87	14.68	18.95	10.83	14.44
40-44	13.79	15.32	14.63	14.91	11.38	14.16
45-49	13.52	15.43	13.64	10.30	11.90	13.08
50-54	12.70	13.76	11.75	6.76	11.54	10.80
55-59	9.47	10.05	8.97	3.55	9.77	8.14
60-64	2.54	4.02	4.60	1.69	6.33	5.09
Education						
High-school or less	48.16	57.93	39.83	42.28	36.57	51.57
Secondary	19.18	15.19	22.39	34.56	31.14	17.79
Tertiary	32.65	26.89	37.78	23.16	32.29	30.64
Sector						
Agriculture	3.04	1.46	3.98	6.86	2.07	2.60
Construction	5.52	10.04	7.99	13.24	7.03	11.17
Industry	13.71	9.26	14.66	9.33	10.56	12.29
Services	77.73	79.24	73.37	70.57	80.38	73.95
Part-time	15.82	17.65	12.38	18.40	17.15	13.65
Temporary rate	11.48	14.44	22.80	39.74		
Unemployment rate	8.54	16.50	14.86	23.65	6.39	6.13
LF participation rate	71.14	66.77	70.60	77.53	74.33	74.43

Source: French Labour Force Survey (2003-2018), Spanish Labour Force Survey (2005-2018) and CPS (2003-2018). Based on individuals aged 16-65.

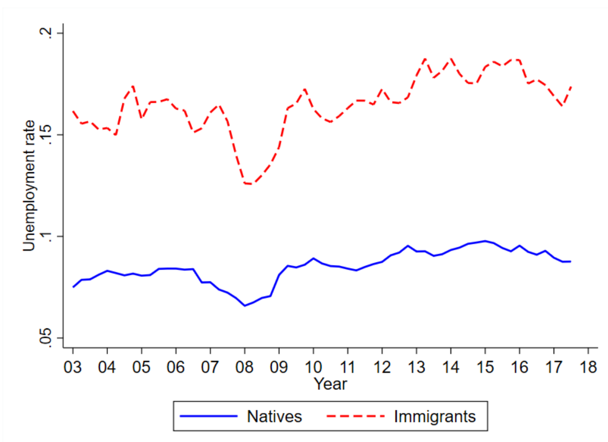
Figure 1: Share of Immigrants and Composition by Nationality



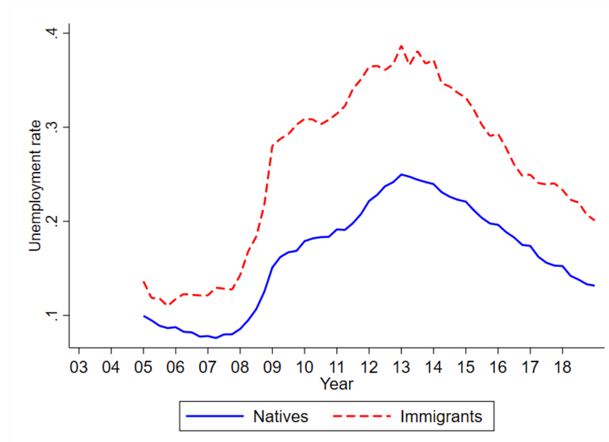
Source: French Labour Force Survey (2003-2018), Spanish Labour Force Survey (2005-2018) and CPS (2003-2018). Based on workers aged 16-65.

Figure 2: Unemployment Rate of Natives and Immigrants

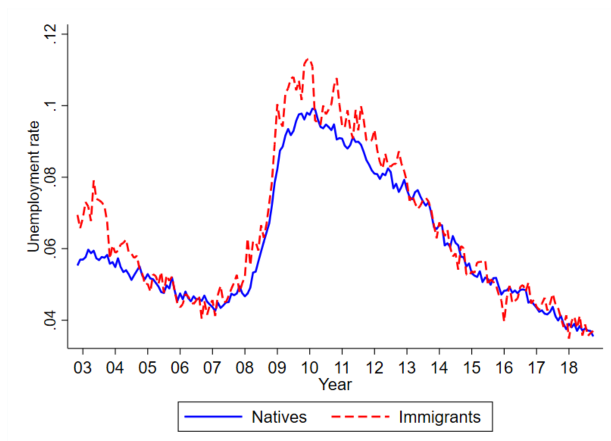
France



Spain



U.S.



Source: French Labour Force Survey (2003-2018), Spanish Labour Force Survey (2005-2018) and CPS (2003-2018). All series are seasonally-adjusted.

3 Worker Flows

3.1 Constructing Worker Flows

In the main analysis, we consider three labour market states, namely employment (E), unemployment (U) and inactivity (I). Exploiting the longitudinal nature of the surveys, we match individuals belonging to two consecutive periods. We then compute individuals' transitions and aggregate them to calculate the gross worker flows and transition rates in each period. We then rely on a three-state Markov model of labour market adjustments. We denote the vectors of stocks as $S_t = (E, U, I)_t^{'}$ which evolves according to :

$$S_t = P_t S_{t-1} \tag{1}$$

where P_t denotes a 3×3 matrix, whose elements $P_{i,j}$ capture the probability of transition from labour status $i \in (E, U, I)$ to labour status $j \in (E, U, I)$. As noted by many papers of the worker flow literature, time series of transition rates are not directly useful because: i) they present seasonal variation, ii) due to margin errors they are not consistent with the exact labour market states (Elsby et al., 2015) and iii) time-aggregation problem can bias the measurement (Shimer, 2012). Consequently, we proceed to three adjustments of the transition rates. We first seasonally adjust gross flows using x13. Along the lines of Elsby et al. (2015), we then compute transition probabilities that are consistent with the observed changes in stocks. Finally, as gross flows provide transition probabilities observed at discrete points of time, in order to correct these measures for possible multiple transitions occurring within a period, we correct gross flows for time-aggregation bias (Shimer, 2012). Details about the exact procedures we employ are left as appendix materials (see Appendix A).

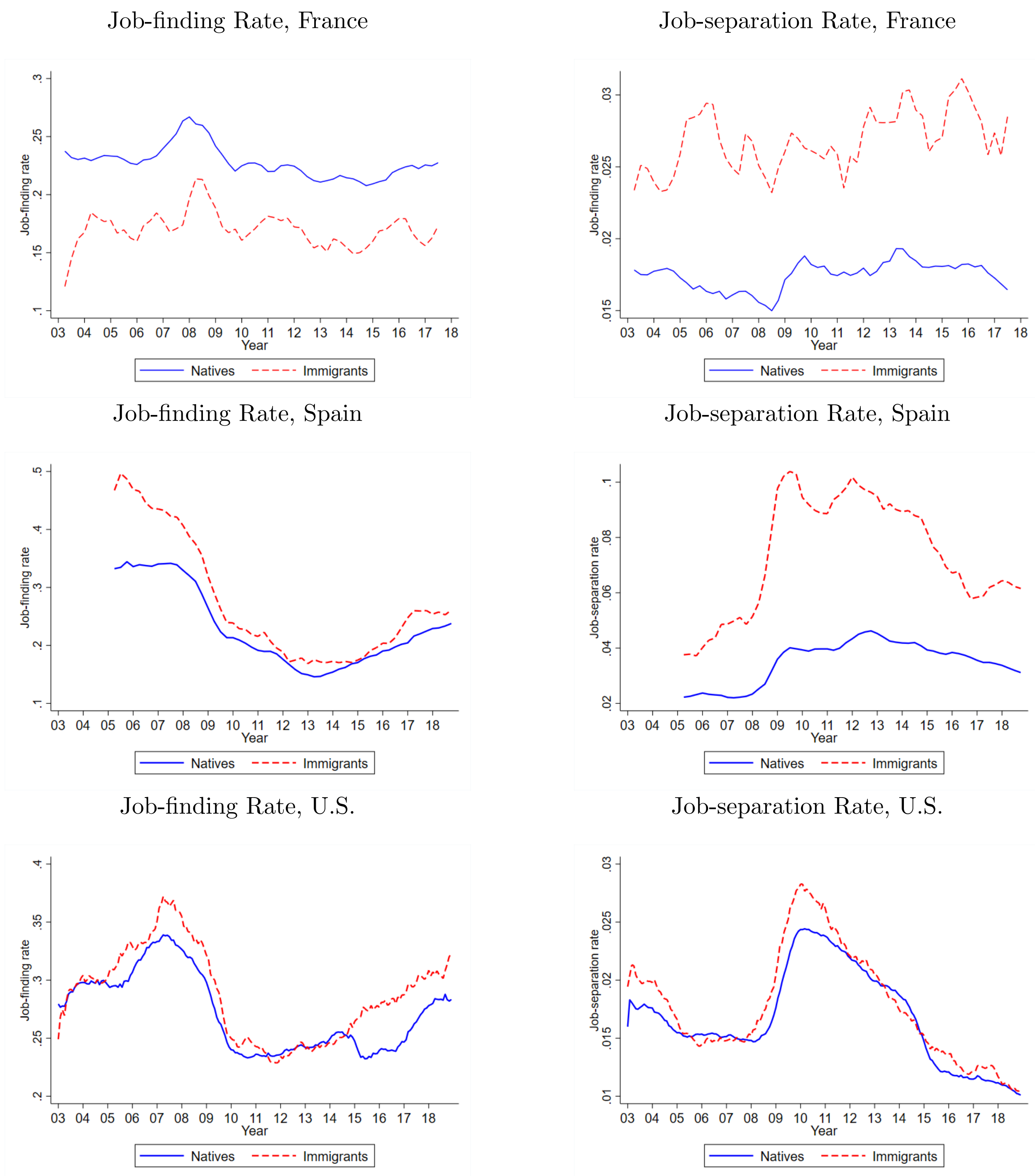
3.2 Unconditional Worker Flows

Figure 3 displays the evolution of immigrants and natives' job-finding and job-separation rates for each country. In France, immigrants' higher unemployment rate is explained by both a lower job-finding rate and a higher job-separation rate. Worker flows exhibit similar cyclical patterns, for both immigrants and natives: the job-finding rate was roughly 30 per cent lower in 2012-13 than in its peak in 2008. The relative increase in the job-separation rates was also similar for natives and immigrants. In contrast, Spain shows an interesting pattern: before the crisis (2005 to mid 2008), the job-finding rate was higher for immigrants than for natives (40-45 per cent vs 32 per cent, respectively), but after the crisis both rates converged quickly to a low level of around 20 per cent. In other words, the decline in the probability of finding a job associated to the Great Recession was bigger for immigrants than for natives. Similarly to the case of France, in Spain, immigrants also exhibit a higher job-separation rate through all the period. However, in contrast with France, the gap widened significantly after the crisis: while for natives the probability of losing a job was multiplied by 1.5 between 2005 and 2012 (from 2.2 per cent to 4 per cent in 2010-12), for immigrants it more than doubled.

In the U.S., the differences in the job-finding and job-separation rates between immigrants and natives are marginal, which explains their similar unemployment rates. Immigrants have a slightly higher job-separation rate, especially in the initial periods (2003-2005) and during the crisis (2008-2013). The absence of unemployment rate gap is explained by a slightly higher job-finding rate for immigrants, as confirmed in the left panel. Like in Spain, the job-separation rate in the U.S. rose more for immigrants than for natives during the Great Recession. Interestingly, and in contrast to the other two countries, we also find differences in timing: immigrants experienced an earlier increase in their job-losing probability, suggestive evidence that immigrants were among the group of workers first hit by the economic downturn.

These figures should be viewed with caution, as differences may be just due to differences in experience, sector composition, or type of jobs (temporary, part-time). The case of Spain is particularity illustrative, as immigrants are relatively more concentrated in temporary jobs and in the construction sector, a higher job-separation rate for immigrants than for natives is far from being surprising. Similarly, the large gap in both rates for France might be also due, in part, to composition effects. In the next section we ask how much of the gap remains after controlling for workers' observables.

Figure 3: Unconditional Labour Market Transitions by Nationality



Note: For Spain and France, transitions are seasonally adjusted using a 4-quarters moving average, constructed from the Spanish Labour Force Survey-Flows and the French Labour Force Survey. For U.S., transitions are seasonally adjusted using a 12-months moving average

3.3 Conditional Worker Flows

To understand the role of composition in explaining the differences in transition rates, we estimate the probabilities of finding a job and exiting from a job for immigrants and natives, using the following linear probability model:

$$UE_{i,t} = \alpha_1 + \alpha_1^{m} imm_i + \alpha_1^{my} imm_i \times year_t + \boldsymbol{\delta} X_{i,t}^1 + \epsilon_{i,t}^1 \tag{2}$$

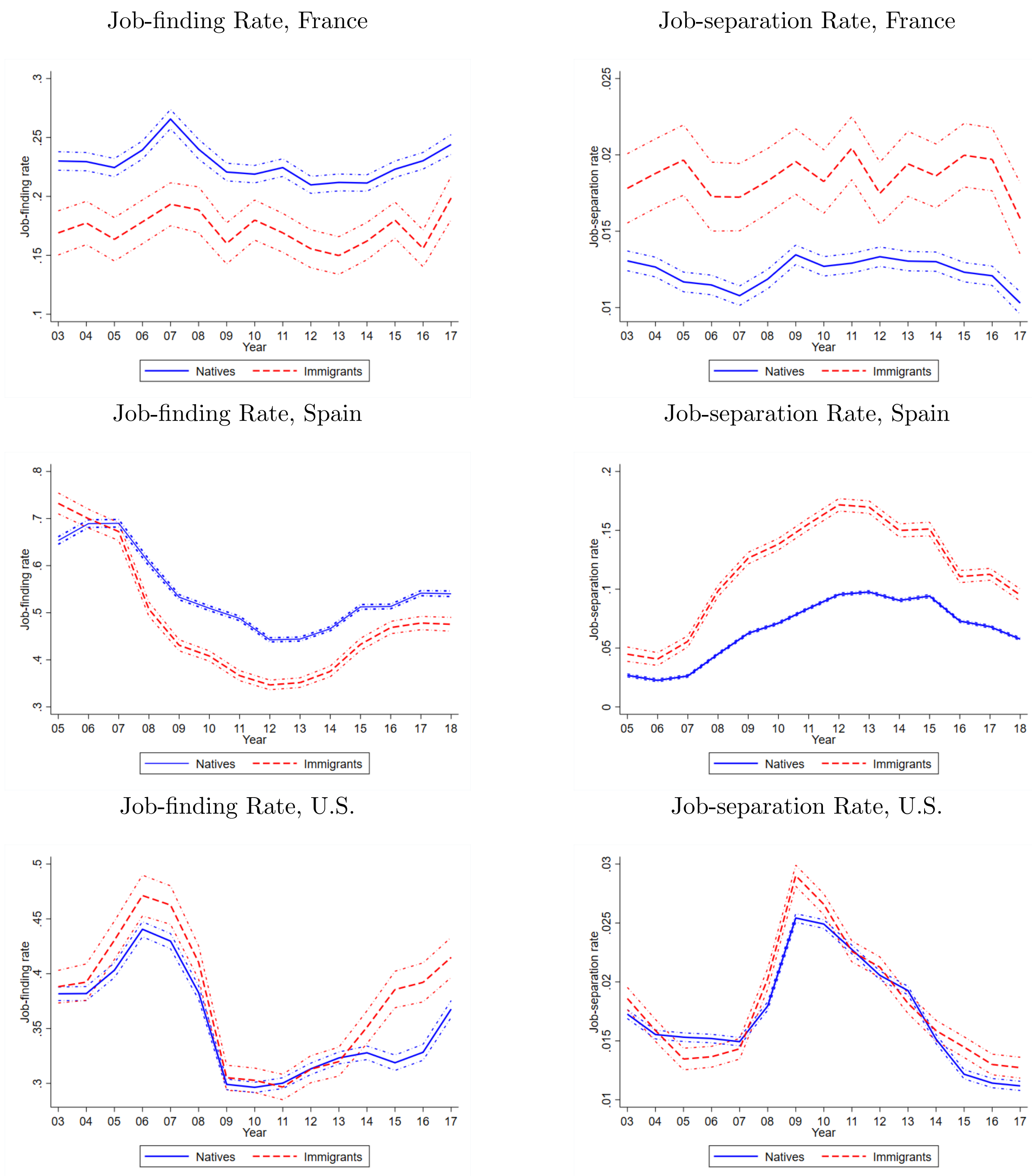
$$EU_{i,t} = \alpha_2 + \alpha_2^{m} imm_i + \alpha_2^{my} imm_i \times year + \boldsymbol{\delta} X_{i,t}^2 + \epsilon_{i,t}^2 \tag{3}$$

where $UE_{i,t}$ ($EU_{i,t}$) is a dummy variable defined only for the unemployed (employed) which takes value 1 if a job is found (lost) at period t (quarter for France and Spain, month for the U.S.) and 0 otherwise; imm_i is a dummy variable that takes the value 1 if the worker is an immigrant and 0 otherwise; $year_t$ denotes year dummies; $\boldsymbol{X}_{i,t}^1$ is a vector of control variables including dummies for education, potential experience, marital status, age, gender, region (state) of residence. It also includes year dummies; $\boldsymbol{X}_{i,t}^2$ includes all variables in $\boldsymbol{X}_{i,t}^1$ and it further adds as controls the occupation, sector of activity, type of contract (permanent or temporary) and type of job (full-time or part-time); $\epsilon_{i,t}$ is the idiosyncratic error term.[3] Using these estimates we can compare the evolution of the predicted job-finding and job-separation rate of a native worker and an immigrant worker with the average characteristics in the economy.

Figure 4 plots the predicted job-finding (left panel) and job-separation rates (right panel) for immigrants and natives. In France, the gap in the job-finding rate is practically unaffected by the inclusion of observable characteristics of workers and jobs. In contrast, difference in the job-separation rates of immigrants and natives is smaller than in the unconditional series, suggesting that composition plays a role. The same finding is observed in Spain: the job-separation rate gap is reduced with the inclusion of observables. This result is not surprising given that Spain and France have dual labour markets, where temporary and permanent jobs coexist. And, as we saw in Subsection 2.3, immigrants have higher shares of temporary jobs, with higher job-separation rates.

[3] We interact the nationality with year dummies instead of quarter dummies to reduce the noise in the estimation due to the small number of unemployment to employment transitions in a given quarter. Regarding control variables, in the Spanish data we also control for the tenure on the job. In the U.S. we do not control for the type of contract because of lack of information.

Figure 4: Conditional labour market transitions by nationality



Note: Residuals (evaluated at the average at means of other covariates) obtained from the estimation of equations (2) and (3) using a linear probability model. All regressions include controls for education, potential experience, marital status, age, gender, region of residence, sector of activity, occupation, type of job (full-time or part-time) and year dummies. For France and Spain we additionally control by the type of contract (temporary or permanent) and tenure. The dashes lines report the 95 percent confidence interval on the prediction. Source: Spanish LFS-Flows (2005-2018), French LFS (2003-2018) and CPS (2003-2018).

Still, our results suggest that among comparable workers, immigrants are roughly 1.5-2 times more likely to separate from a job than natives, both in Spain and France.

Interestingly, the pre-crisis job-finding rate gap between immigrants and native vanishes once we take into account composition effects. In both countries the conditional job-finding rate was higher among natives than immigrants. This is the opposite to what we saw in Figure 3, where we found no differences across the two rates. Figures are almost unaffected in the U.S., especially regarding the job-finding rate. But we do find that the pre-crisis unconditional job-separation rate gap is closed.

There are still large differences regarding cyclical patterns. The pattern can be clearly seen both in Spain and the U.S. in both flows: the drop (increase) in the unconditional job-finding (separation) rate is higher for immigrants than for natives. In Appendix B we provide additional evidence for the cyclicality of the labour market flows.

3.4 What Drives Unemployment?

With estimates of transition rates in hand, our goal is to decompose cyclical fluctuations in unemployment rate into contributions attributable to each of the flow hazards. To do so, we run the dynamic decomposition of Elsby et al. (2015) which does not assume that labour market states are at their steady-state levels. The relaxation of this hypothesis is of first importance given the relative *sclerotic* nature of labour market dynamics in European countries (Elsby et al., 2013). The output of the variance decomposition is a set of β values that can be expressed as the share of unemployment rate variance that is accounted for by the transition rate from $i \in (E, U, I)$ to $j \neq i$:

$$\beta_u^{ij} = \frac{\text{Cov}\left(\Delta u_{t-1,t}, \Delta \tilde{u}_{t-1,t}^{ij}\right)}{\text{Var}(\Delta u_{t-1,t})} \tag{4}$$

where, Δ is the first-difference operator and $\tilde{u}_{t-1,t}^{ij}$ the first difference in a counterfactual unemployment rate obtained when only the transition rate from $i \in (E, U, I)$ to $j \neq i$ fluctuates. To compute $\tilde{u}_{t-1,t}^{ij}$, we proceed as follows. First, we compute labour market stock changes that are driven by contemporaneous but also *past* changes in transition rates. This recursive formulation of stock variations is at the heart of the non-steady state decomposition. Second, we express the variance of any given labour market stock as the sum of its covariance with any counterfactual obtained in the

Table 2: Elsby et al. (2015) Non-Steady State Decomposition

	France		**Spain**		**U.S.**	
	Natives	Immigrants	Natives	Immigrants	Natives	Immigrants
$E \rightarrow U$	0.22	0.28	0.22	0.34	0.32	0.31
$E \rightarrow I$	-0.02	0.07	-0.01	-0.02	0.05	0.01
$U \rightarrow E$	0.34	0.49	0.50	0.44	0.20	0.23
$U \rightarrow I$	0.20	0.08	0.18	0.11	0.19	0.21
$I \rightarrow E$	0.01	-0.00	0.04	0.03	-0.01	0.00
$I \rightarrow U$	0.26	0.09	0.07	0.10	0.25	0.25
Relative contribution (sum to 100)						
Job-finding rate vs. Job-separation rate						
	61-39	63-37	69-31	57-43	61-39	58-42
Flows to non-participation vs. Flows within participation						
	44-56	23-77	28-72	22-78	48-52	47-53

Note: the gross flows series are previously seasonally adjusted using the X13 Census programme and the transition probabilities are corrected for time aggregation bias using the methodology applied by Shimer (2012).Series are "smoothed" with a 3-order moving average for France, and Spain, and a 9-order moving average for the U.S.

previous step. In what concerns the unemployment rate, an additional step is needed as the Elsby et al. (2015) method provides a variance decomposition of labour market stocks. We so employ a first order linear approximation to express changes in the unemployment rate Δu_t. Appendix C provides more details about the variance decomposition.

Table 2 displays the results of the decomposition for both natives and immigrants in each country. Cyclical fluctuations of unemployment over time are primarily driven by job-finding rates, for both immigrants and natives. In the U.S. and France, there are few differences between the two groups. Out of the total contribution of flows in and out of employment, 61-63 per cent are attributed to the job-finding rate in France and 61-58 per cent in the U.S, for natives and immigrants respectively. Larger differences are found in Spain, where the job-separation rate is more relevant for explaining the cyclical behaviour of unemployment for immigrants than for natives (43 against 31 per cent).

The decomposition exercise also assesses the relative importance of flows in and out of the labour force. The bottom panel of Table 2 provides the relative split of the contribution of flows to non-participation (EI + UI) and flows within participation (EU+UE). Flows within participation account for most of the variation in unemployment for all countries, especially in Spain. We also find that the relative importance of within participation flows is higher for immigrants than for natives across all countries, suggesting a higher attachment to the labour market among immigrants. In U.S. and Spain, the relative split is closer for immigrants and natives (48-52 and 47-53 in the U.S. and 28-72 and 22-78 for Spain), but a large difference arises in France, where flows to non-participation are disproportionably more relevant for natives than for immigrants (44 against 23 per cent).

4 VAR Model

In the final section, we use a VAR model to investigate the dynamic responses of native and immigrant workers transition rates after an immigration shock, in the tradition of the literature on worker flows (Fujita, 2011; Canova et al., 2012; Hairault and Zhutova, 2018; Fontaine, 2019), as well as the literature on the dynamic effect of immigration shocks in the economy (d'Albis et al., 2016; d'Albis et al., 2019; Furlanetto and Ørjan Robstad, 2019).

4.1 Empirical Strategy

The reduced-form VAR We start our baseline econometric strategy by the estimation of the following reduced-form VAR model:

$$Y_t = \Gamma_c + \sum_{k=1}^{K} \Gamma_k Y_{t-k} + \nu_t \tag{5}$$

where Y_t is an $N \times 1$ vector containing our N endogenous variables, Γ_c is an $N \times 1$ vector of constants, Γ_k for $k = 1, \dots, K$ are the $N \times N$ matrix of coefficients, K is the total number of lags included in the VAR and ν_t the $N \times 1$ matrix of reduced-form residuals. Denoting by L the lag operator and rearranging equation (5) we can write $\Psi(L)Y_t = \nu_t$. Assuming that $\Psi(L)$ is invertible, the VAR has a Wold moving-average representation:

$$Y_t = \Psi(L)^{-1}\nu_t = C(L)\nu_t \tag{6}$$

with $C(L)$ a matrix of polynomials in the operator L. In our baseline exercise, we include five endogenous variables (so $N = 5$) in the following specific order so that:

$$Y_t = \begin{pmatrix} \Delta \log \omega \\ \Delta \log \lambda_{nat}^{EU} \\ \Delta \log \lambda_{nat}^{UE} \\ \Delta \log \lambda_{immi}^{EU} \\ \Delta \log \lambda_{immi}^{UE} \end{pmatrix} \tag{7}$$

with ω being the share of immigrants in the working age population as measured from the Labour Force Surveys of the three countries. Each endogenous variable enters the VAR in first differences of its logarithm. Such a specification choice follows the application of the standard Augmented Dickey-Fuller test indicating that most variables are non-stationary in (log-)levels, but are stationary when expressed in first differences of their logarithms. Recall that the time series are quarterly and cover the 2003Q1-2019Q4 period for France and the U.S. and the 2005Q1-2019Q1 period for Spain.[4]

Along the lines of Balleer (2012), Balleer and van Rens (2013) and Canova et al. (2012) our reduced-form VAR is estimated within a Bayesian framework and we employ the Minnesota prior. The prior incorporates a fixed residual variance governing the tightness on own lags, other lags as well as the decay of the lags. The Minnesota prior is flexible enough to enable us including a large number of lags. In our baseline specification, we use four lags.[5]

Identifying structural immigration shock In reduced-form VAR models, the residuals do not have meaningful economic interpretation because its variance-covariance Σ is not diagonal. The main purpose of the identification strategy is to find a mapping that allows for the retrieving of structural shocks ε_t from the reduced-form residual ν_t. Under standard assumptions,[6] reduced-form residuals and structural shocks are related by the following relationship:

$$\nu_t = D\varepsilon_t \tag{8}$$

[4] In the case of the U.S., time series recovered from the CPS are monthly. Consequently, we apply an arithmetic average to retrieve the same frequency as for France and Spain.

[5] Our main results do not depend on that choice.

[6] It is assumed that structural shocks are mutually independent. Furthermore, we adopt the standard normalization that $E(\varepsilon_t \varepsilon_t^{'}) = I_N$.

with $\varepsilon \sim N(0, I_N)$, where I_N is an $(N \times N)$ identity matrix and where D is a non-singular matrix. To construct a structural immigration shock, we follow d'Albis et al. (2016), d'Albis et al. (2019) and d'Albis et al. (2021) and the matrix D is computed as:

$$\Sigma = E(\nu_t \nu_t^{'}) = DE(\varepsilon_t \varepsilon_t^{'})D^{'} = DD^{'} \tag{9}$$

where D is the unique lower-triangular Choleski factor of Σ. Given the structure of the matrix D the identifying scheme relies on the following assumption: variables ordered first in Y_t can impact other following variables contemporaneously, while variables ordered last impact variables entering before them in Y_t with a period of lag. In our specific case, this implies that a shock on the share of immigrants, which is the first variable in Y_t, can affect the other four labour market variables contemporaneously. Moreover, each labour market flow variable entering Y_t after ω, has a zero immediate impact on the share of immigrants and affects it only with a one-period delay.

From our point of view, such an identification scheme is reasonable as the decision to migrate to another country is a lengthy process (d'Albis et al., 2019). Moreover, it is also known that the migration decision is based, not only on the current economic performances of the host country, but also on its economic conditions for many years that preceded the individual decision. As we have quarterly data, it is unlikely that variations in labour market variables influence within a quarter the share of immigrants in the host country. Arguably, imposing a zero response of the share of immigrants to variation in labour market variables is supported by the current evidence of the literature (d'Albis et al., 2016; Smith and Thoenissen, 2019; d'Albis et al., 2019).

Our migration variable enters the VAR as a first difference of its logarithm. Consequently, this variable captures the percent net change in the share of immigrants as estimated by the Labour Force Surveys of the three countries. To the best of our knowledge, we are the first to use such a migration measure in the context of a VAR framework. Most papers employing VAR to estimate the economic impact of migration rely on net migration, namely the difference between inflows and outflows of people in a given host country during a given time period. However, the latter variable is generally only available at an annual frequency (d'Albis et al., 2019, 2021).[7]

[7] A notable exception is Furlanetto and Ørjan Robstad (2019) who work with Norwegian quarterly data.

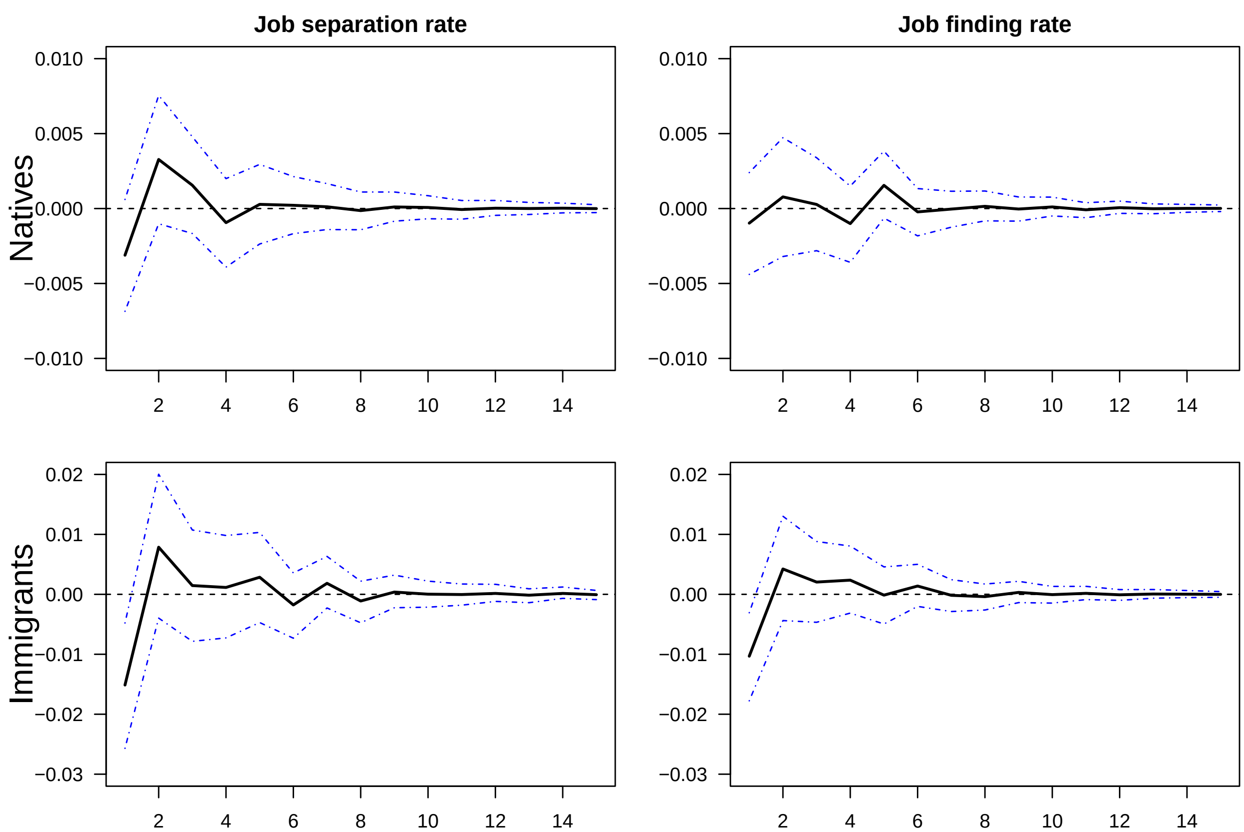
4.2 Main Empirical Results

We start by estimating a simpler VAR model, with only three endogenous variables: immigrants share and both natives' and immigrants' unemployment rates. The identification scheme is as explained before: a shock to the unemployment rate does not have a contemporaneous impact on the share of immigrants in the country. The size of the shock is set to a one per cent. Results are displayed in Figure 8 to 10, in the Appendix. In France, a 1 per cent increase in the immigrant share has a negative and significant impact on both natives' and immigrants' unemployment rate, which goes down by around 0.2 and 0.5 per cent respectively in the quarter which the migration shock takes place. The effect vanishes after 2-5 quarters. For the U.S., the impact on natives' unemployment rate is close to zero, while in Spain, the migration shocks leads to a contemporaneous increase in both natives' and immigrants' unemployment rates, but the effect is not significant for natives. In what follows, we estimate the baseline VAR model with the transition flows, instead of a stock variable (unemployment rate), as endogenous variables. With this approach, we aim to split the overall employment impact into the impact on the probability of finding a job and the impact on the probability of losing a job.

Impulse response analysis The estimates of the impulse responses for France, Spain and the U.S. are shown in Figures 5 to 7. In the U.S. (Figure 7), a migration shock leads to a significant decrease in natives' job-separation rates, with the effect lasting around 2-3 quarters. Quantitatively, the results suggest that a 1 per cent increase in the migrants share leads to a 1.5 per cent decrease in native job-separation rates after 2 quarters, and 0.3 after 3 quarters. The migration shock also leads to a decrease in natives' job-finding rate, of a similar magnitude (around 1 per cent in its peak, after 2 quarters). These two opposite effects explain why we find close-to-zero effects on the overall unemployment rate when we estimate the simple VAR model, and emphasize the importance of splitting the two effects. The migration shock generates higher turnover, increasing both ins and outs of employment. Looking at immigrants' variables, we observe that the job-separation rate declines by 1.8 per cent 2 quarters after the shock while at other horizons the effects are not significant. Similarly, our model indicates a weak and non-significant response for the job-finding rate of immigrants. Again, estimated responses of our five-variables VAR model are in line with what is obtained in the VAR with only the unemployment rates.

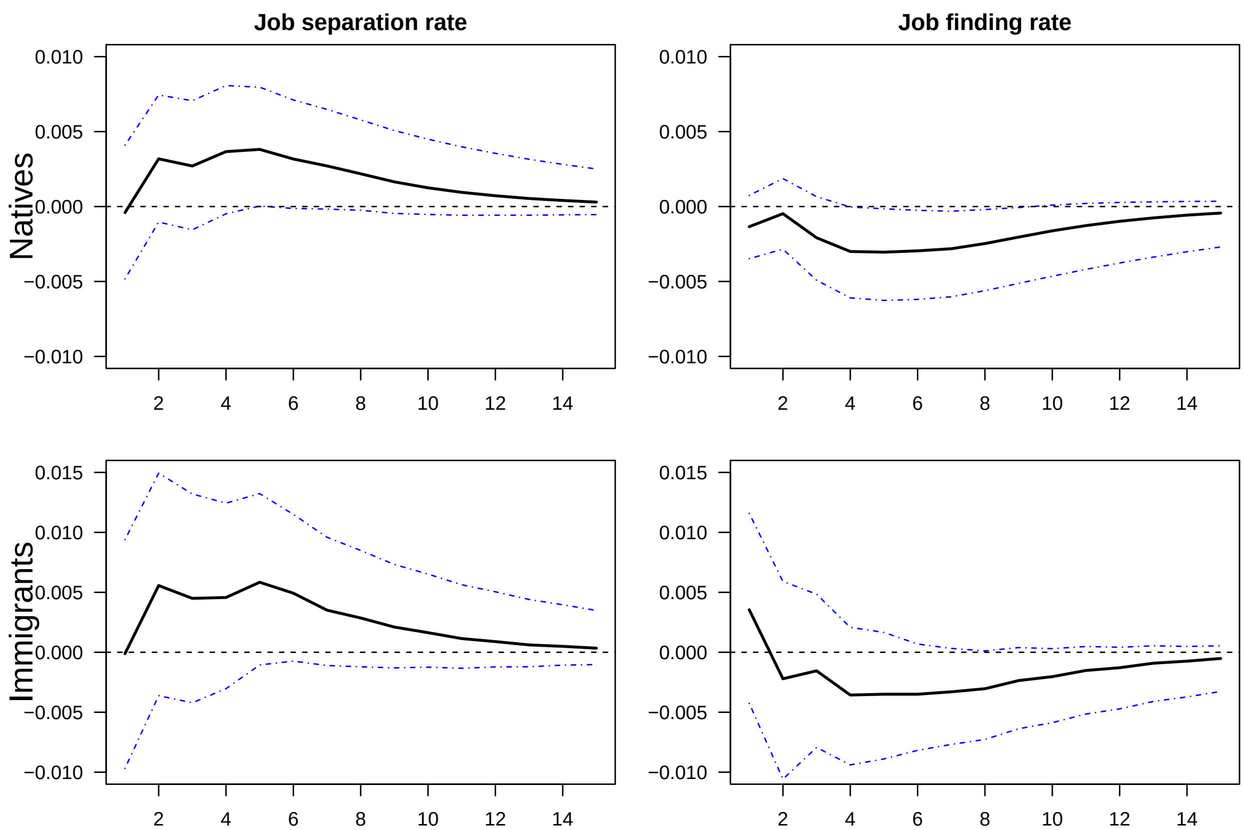
In Spain and France, the response of natives' labour market flows to immigration is very moderate: in both countries, the impulse response functions suggest a negative contemporaneous effect on natives' job-finding and job-separation rates, but the effect is non-significant. In the case of Spain, estimates show a higher persistence of the effect of the migration shock on job-finding rates, as the point estimate remains below zero for more than 10 quarters after the shock. But, as we just mentioned, point estimates are not significant. The bottom panel of Figure 5 shows that immigrants' flows do react to the migration shock: job-separation and job-finding rates drop by 1.5 and 1 per cent, respectively, at the quarter of the migration shock in France. By contrast, the estimated responses for immigrants are non-significant in Spain even if the point estimates suggest an increase in the job-separation rate and a decrease in the job-finding rate. Overall, our VAR model for the three countries depicts a common stylized fact: an inflows of foreign workers has a weak and mostly non-significant effect on the job-finding and job-separation rates of both immigrants and natives.

Figure 5: Impulse Response Function After an Immigration Shock in France



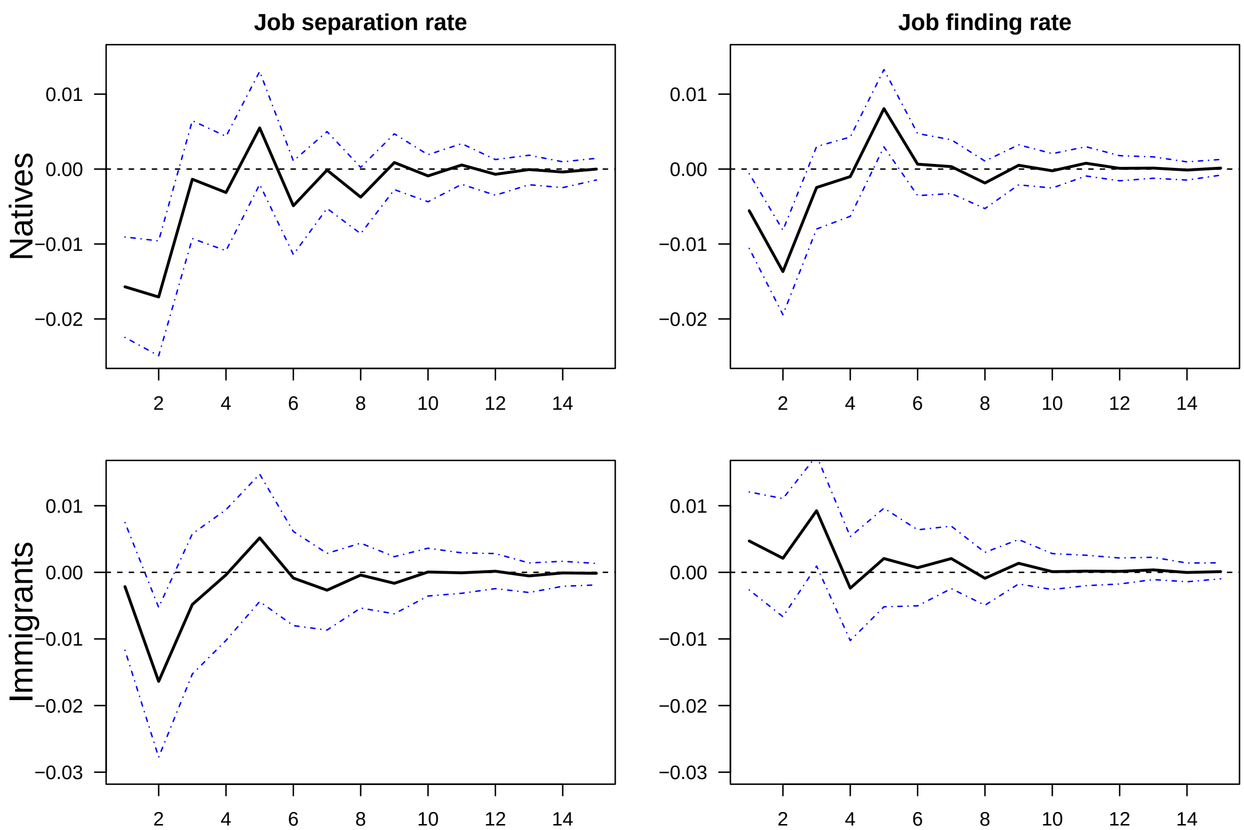
Notes: IRFs to a one percentage point increase in the immigration share are plotted. Confidence intervals are the classical Bayesian bands. Source: French Labour Force Survey (2003-2018).

Figure 6: Impulse Response Function After an Immigration Shock in Spain



Notes: IRFs to a one percentage point increase in the immigration share are plotted. Confidence intervals are the classical Bayesian bands. Source: Spanish Labour Force Survey (2005-2018).

Figure 7: Impulse Response Function After an Immigration Shock in U.S.



Notes: IRFs to a one percentage point increase in the immigration share are plotted. Confidence intervals are the classical Bayesian bands. Source: CPS (2003-2018).

Forecast Error Variance Decomposition Table 3 shows the percentage contribution of migration shocks to the forecast error variance at any given quarter after the migration shock. At the quarter of the shock, the shock's contribution to the volatility of natives' job-separation rates ranges from 1 per cent in Spain to roughly 8 per cent in the U.S. Over 4 quarters, those contributions are 5 per cent and 15 per cent, respectively. Results for the shock's contribution to the variation of the job-finding rate are quantitatively similar.

The contribution of migration shocks to the volatility of natives' flows increases more in Spain than in France over time. At 4 quarters-horizon, the contribution of the migration shock to fluctuations of natives' job-separation rates ranges from 5 per cent in Spain and France to more than 14 per cent in the U.S. The contribution of migration shocks to the fluctuations in the job-finding rate is quantitatively similar: higher in the U.S. (12 per cent in a 4-quarter horizon) than in France and Spain (3.7 and 4.9 per cent, respectively).

Table 3: Percentage of Fluctuations Attributable to the Immigration Shock

	Native		Immigrant	
Quarter	Job-separation rate	Job-finding rate	Job-separation rate	Job-finding rate
		France		
1	1.21	0.77	4.03	3.28
2	3.02	2.03	4.83	4.39
4	4.70	3.68	5.58	5.92
8	5.48	5.02	6.67	6.54
16	5.64	5.16	6.74	6.65
		Spain		
1	0.84	1.33	0.85	0.92
2	2.77	2.15	2.62	1.78
4	5.34	4.96	4.57	3.22
8	8.37	10.79	6.28	4.88
16	9.55	13.30	6.87	5.60
		U.S.		
1	7.88	2.22	3.06	1.03
2	14.92	11.71	8.88	2.16
4	14.09	12.03	10.96	5.18
8	15.20	14.29	12.43	6.34
16	15.30	14.47	12.97	6.64

Note: Figures represent the median of the posterior distribution. Quarter 1 stands for the year of the shock. The percentage of fluctuations attributable to a given shock is the forecast error variance of the corresponding variable explained this shock.

5 Conclusion

Immigrants come to a foreign country mainly to find a job and improve their lifes. It is therefore important to understand their labour market outcomes and how they compare to those of natives. We provide a comprehensive analysis of labour market transition rates of immigrants and natives for France, Spain, and the United States. We document differences in labour market flows between immigrants and natives, and show that most facts are not common across these three countries and cannot be accounted by differences in the composition of their immigrant populations in terms of education, age and sector. While various previous studies highlight several labour market disadvantages that immigrants experience in certain countries such as, greater unemployment risk, or greater vulnerability to economic downturns for immigrants, we find that such conclusions cannot be generalized across the three countries we examine.

In the U.S., labour market transitions of immigrants and natives are remarkably similar in terms of size, relative importance and cyclicality. In France and Spain, on the other hand, immigrants face a disadvantage relative to natives in terms of both job finding and separation rates, which explains why the former have higher unemployment rates. Our results, however, indicate a greater vulnerability of immigrants to business cycles shocks in Spain, but not in France. In the latter, the gaps in job finding, separation and unemployment rates between natives and immigrants remain stable over the period.

The specific channels of migrant entry and legislation in each of these countries and their distinct labour market institutions are important for shaping these differences. Immigrant entry opportunities into the U.S. strongly depend on networks as the two main entry channels are family unification and employment-based entry, which requires finding an employer prior to entry. U.S. immigrants are thus more concentrated in terms of origin and background and may have stronger ties and connections with the local community. Spain is characterized for the high duality in their labour markets, where temporary and permanent jobs coexists, while immigrants in France are more attached to the labour force. The latter may be related to the fact that a relative large portion of immigrants in France come from Africa, implying less frequent return migrations.

In the final part of the paper, we undertake another empirical approach to provide an answer to the question about the effect of a migration shock on labour market outcomes. By relying on a

VAR model for the identification of a migration shock, we find that a shock increasing the share of immigrants has a weak and mostly non-significant effect on the job-finding and job-separation rates of natives as well as immigrant workers. Furthermore, the identified migration shock has a low contribution to the variation of transition rates of both natives and immigrants. Although our approach is silent on the effect on wages, our result indicate that inflows of immigrants do not deteriorate labour market outcomes of the host countries.

While immigration has long been a topic of interest, the large inflows of immigrants into many developed countries in the last decades keep immigration at the top of the current and future research agenda. Recent studies build macroeconomic models based on the flow approach to the labour market to analyze immigrants' labour market outcomes and effects. We hope the facts about transition rates of immigrants and natives that we document can contribute to the development of this literature.

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Appendices

A Adjustments applied before running variance decomposition

This section presents adjustments applied to transition rates, namely the margin-error correction and the temporal aggregation correction. Then the variance decomposition method of Elsby *et al.* (2015) is detailed. Based on the raw microdata, we first compute labour market stocks and gross worker flows for each time period t. We then adjust the resulting time series for seasonality using the X-13ARIMA-SEATS Seasonal Adjustment Program of the Census Bureau. After the series are adjusted we compute their transition rates $\tilde{p}_t^{ij}$ with $i \in \{E, U, I\}$, $j \in \{E, U, I\}$ and $i \neq j$. In particular, let us denote gross worker flows by two consecutive capital letters: the first one is the origin of the flow, the second one its destination. Transition rates $\tilde{p}_t^{ij}$ are the number of individuals who move from state i to state j between $t-1$ and t divided by the number of individuals in state i in period $t-1$. For instance, the job separation rate to unemployment is: $\tilde{p}_t^{EU} = \frac{EU_t}{P_{t-1}}$. At this stage, it should be observed that we perform all these adjustments independently for the native and non-native groups.

Adjustment for margin error

The sample design of the *Labour Force Surveys*, but also the adjustment for seasonality, imply that obtained transition rates do not lead to the exact measures of changes of labour market stocks. To deal with this issue, we apply for each time period what the worker flow literature calls the "margin-error" adjustment. This adjustment restricts the estimates of transition rates to be consistent with the observed evolution of the corresponding labour market stocks. In general, this adjustment has only a marginal incidence on the level and the cyclicality of transition rates. We now describe in detail the method.

We denote the vector of labour market stock observed at each period t as follows:

$$S_t = \begin{pmatrix} E_t \\ U_t \\ I_t \end{pmatrix} \tag{10}$$

After, normalizing labour market stocks by the working-age population, the system has only two dimensions and the corresponding vector of stocks is denoted as s_t. Denoting by Δ the first-difference operator, the stock evolution between period t and $t-1$ is:

$$\begin{aligned}\Delta s_t &= \begin{pmatrix} -E_{t-1} & -E_{t-1} & U_{t-1} & 0 & I_{t-1} & 0 \\ E_{t-1} & 0 & -U_{t-1} & -U_{t-1} & 0 & I_{t-1} \end{pmatrix} \begin{pmatrix} p_t^{EU} \\ p_t^{EN} \\ p_t^{UE} \\ p_t^{UI} \\ p_t^{IE} \\ p_t^{IU} \end{pmatrix} \\ \Delta s_t &= X_{t-1}\mathbf{p}_t \end{aligned} \tag{11}$$

where p^{ij} (with $i \neq j$) are stock-consistent transition rates. However, from the data we do not observe the matrix of transition rates $\mathbf{p}_t$ but solely the non-adjusted one $\tilde{\mathbf{p}}_t$. To retrieve the former with only information on the later we minimize, as in Elsby et al. (2015), the weighted sum of squares of margin-error adjustments under the constraint (18):

$$\text{minimize}(\mathbf{p}_t - \tilde{\mathbf{p}}_t)^{'}\mathbf{W}_t(\mathbf{p}_t - \tilde{\mathbf{p}}_t), \text{ subject to } \Delta S_t = X_{t-1}\mathbf{p}_t \tag{12}$$

where $\mathbf{W}_t$ is a matrix proportional to the covariance matrix of $\tilde{\mathbf{p}}_t$ (also called the weighting matrix).[8] Denoting by μ the vector of Lagrange multipliers associated to (12), we derive that

$$\begin{bmatrix} \mathbf{p}_t \\ \mu \end{bmatrix} = \begin{bmatrix} \mathbf{W}_t & X_{t-1}' \\ -X_{t-1} & 0 \end{bmatrix} \begin{bmatrix} \mathbf{W}_t\tilde{\mathbf{p}}_t \\ \Delta S_t \end{bmatrix} \tag{13}$$

Since all elements of the right hand side of (13) are observed, it is quite straightforward to get stock-consistent transition rates.

Adjustment for time aggregation bias

The last adjustment we perform is to deal with the fact that discrete transition rates are subject to time aggregation bias. Indeed, the *Labour Force Surveys* we use in this paper allow us to record individual labour market positions at a quarterly frequency (monthly in the U.S.). This discrete time representation of labour market dynamics could miss some transitions since all "infra-period" multiple movements are not observed. The problem is that, within a quarter an individual can make multiple transitions and the matching of observations belonging to two consecutive surveys will catch at most one. To deal with this issue, we follow Elsby et al. (2015) and we exploit the relationship governing the "eigenvalue-eigenvector" decomposition of the between the discrete-time and the continuous-time representation of the Markov-chain.

Let P_t denote the square matrix of order 4 of discrete time transition rates and H_t its continuous time counterpart. For every time period, we use the eigen-decomposition of P_t such that: $P_t =$

[8] See the appendix of Elsby et al. (2015) to see the exact form of the weighting matrix $\mathbf{W}_t$.

$V_t D_t V_t^{-1}$ where D_t is a diagonal matrix whose elements are the eigenvalues of P_t and V_t the matrix of associated eigenvectors. If the diagonal elements of D_t are distinct, real and non-negative (which is always the case in our samples) there is a unique relationship between the eigenvalues of P_t and H_t. More specifically, if the eigenvalues of H_t are all distinct, we can write H_t such that: $H_t = V_t C_t V_t^{-1}$ where C_t is the log value of D_t. With knowledge of P_t, V_t, and D_t it is straightforward to get C_t, H_t and the underlying hazard rates h_t^{ij}. Last, with estimates of h_t^{ij} in hand, we infer values of time-aggregation adjusted transition probabilities λ_t^{ij} by applying $\lambda_t^{ij} = 1 - \exp(-h_t^{ij})$.

B The cyclicality of labour market flows

Section 3.3 suggests that immigrants' flows were more sensitive to the outset of the Great Recession, even when controlling for composition effects. We now estimate a linear probability model to quantify the differential impact of the crisis on the employment transitions of immigrants and natives:

$$UE_{i,t} = \beta_1 + \beta_1^m imm_i + \beta_1^c crisis_t + \beta_1^{mc} imm_i * crisis_t + \boldsymbol{\delta}_1 \boldsymbol{X}_{i,t}^1 + \varepsilon_{i,t}^1 \tag{14}$$

$$EU_{i,t} = \beta_2 + \beta_2^m imm_i + \beta_2^c crisis_t + \beta_2^{mc} imm_i * crisis_t + \boldsymbol{\delta}_2 \boldsymbol{X}_{i,t}^2 + \varepsilon_{i,t}^2 \tag{15}$$

where the dummies $UE_{i,t}, EU_{i,t}$ and imm_i are defined as above; $crisis_t$ is a dummy capturing the Great Recession, which equals 1 in periods with a negative quarterly growth rate of real GDP[9] and 0 otherwise; the vector $\boldsymbol{X}_{i,t}^1$ and $\boldsymbol{X}_{i,t}^2$ includes the same control variables as above; and $\varepsilon_{i,t}$ is the idiosyncratic error term. The coefficients of interest are $\beta_1^{mc}, \beta_2^{mc}$, which are associated with the interaction term of the variables imm_i and $crisis_t$. Their signs and magnitudes will be used to quantify the differential impact of the crisis on the probability of finding (losing) a job between immigrant and native workers.

The results of the estimation of Equation (14) and (15) are displayed in Table 4 and 5, respectively. The baseline estimation (which includes all controls) for the probability of finding a job can be found in column (3). The sign and significance of β_1^{mc} indicates that the drop in the probability of finding a job during the crisis was significantly stronger for immigrants than natives in Spain, but not in the U.S. or France. For Spain, the estimation suggests that, ceteris paribus, the crisis is associated with a 5 p.p. decrease in natives' job-finding rates (β_1^c), while for immigrants it dropped by 16.1 p.p. ($\beta_1^c + \beta_1^{mc}$), implying that that the crisis hit 3 times as high the probability of finding a job for immigrants than for natives. Given that non-crisis immigrants' predicted job-finding probability was 51.1 per cent (Table 6 in Appendix), our estimates suggest that the decrease in this probability was seizable, around 31 per cent.

For the U.S. we find that the impact of crisis for natives and immigrants is not significantly different from zero: the crisis is associated with a 3.6 p.p. drop in the job-finding rate of both workers. In France we do not find any cyclical heterogeneity in the impact of the crisis between natives and immigrants. In contrast, our estimates shows that job-finding rates are lower for the latter group during non-crisis times: among comparable workers, immigrants' job-finding rates are 4.3 p.p. lower (β_1^m) that natives'. In relative terms, this means that their chances of finding a job are around 18 per cent lower than natives (see Table 6 in the Appendix for the predicted values). This is opposite to the estimates for Spain, where β_1^m is positive and significant (at 1 per cent level), implying that immigrants experience higher job-finding rates in non-crisis times. We set out estimates of Equation 15 in Table 5. For Spain and the U.S. we find that the increase in the

[9] 2008Q3-2013Q2 for Spain, 2008Q2-2009Q2 for France and 2008m6-2009m6 for the U.S.. The results are robust to changes in the definition of the dummy. Results are also very similar when defining the dummy crisis as taking value 1 for all periods after 2008 and 0 otherwise.

Table 4: Estimation results: UE

	France			Spain			U.S.		
	(1)	(2)	(3)	(1)	(2)	(3)	(1)	(2)	(3)
β_1^m	-0.056***	-0.062***	-0.043***	-0.079***	0.022***	0.013**	0.019***	0.0168***	-0.007*
	(0.003)	(0.004)	(0.004)	(0.002)	(0.006)	(0.006)	(0.002)	(0.004)	(0.004)
β_1^{mc}		-0.001	-0.002		-0.122***	-0.111***		-0.001	0.002
		(0.010)	(0.010)		(0.007)	(0.007)		(0.008)	(0.008)
β_1^c		-0.022***	-0.022*		-0.049***	-0.050***		-0.036***	0.036***
		(0.007)	(0.006)		(0.009)	(0.009)		(0.006)	(0.006)
Year FE	YES	YES	YES	YES	YES	YES	YES	YES	YES
Controls	NO	YES	YES	NO	YES	YES	NO	YES	YES
Observations	202244	202244	201960	509655	509655	502629	409861	409861	409861

Notes: Regression of a dummy variable for the transition from unemployment to employment (UE, in column (1)) and from employment to unemployment (EU, in column (2)) on dummies for the migration status, crisis and the interaction term of the last two. Both regressions include controls for education, potential experience, marital status, age, gender, region of residence,occupation, sector of activity, type of job and year dummies. The regression for EU additionally include as controls the type of contract (permanent or temporary), type of job (full-time or part-time), and tenure. Significance levels: $^{*}p < 0.05$, $^{**}p < 0.01$, $^{***}p < 0.001$. *Spanish LFS-Flows (2005-2018), French LFS (2003-2018) and CPS (2003-2018).*

probability of losing a job during the crisis was higher for immigrants than natives, with estimates of 0.003 and 0.043, respectively. These numbers are read as follows: among comparable workers, during the crisis the increase in job-separation probability was 0.3 p.phigher for immigrants than natives in U.S. and 4.3 p.p in Spain. These numbers indicate large sensitivity to recessions for the immigrant population. For example, in the U.S. during non-crisis times, an immigrant with average job/worker characteristics have a 0.69 per cent probability of losing a job in non-crisis months. Our estimates therefore suggest that immigrants' job-losing probabilities almost doubled in the crisis, while for natives' probability rise by roughly 20 per cent (1.18 to 1.42). For Spain, as expected from previous sections, magnitudes are even higher: ceteris paribus, the crisis is associated with a 1.1 p.p. increase in natives' job-separation rate (β_2^c), while for immigrants it raised by 5.4 p.p. ($\beta_2^c + \beta_2^{mc}$).

We find a similar sensitivity of job-separation rates for immigrants and natives in France, as estimates suggest that $\beta_2^{[mc}$ is not statistically different from zero. Again, we also find that in France job-separation rates are higher for immigrants than natives in non-crisis times. Remarkably, comparing our estimates in column (2) and (3) highlight the importance of accounting for composition effects, as otherwise we would be overestimating both β_2^m and β_2^{mc}. For example, in Spain and France, adding job/worker characteristics cuts by three the non-crisis job-separation gap between immigrants and natives.

Table 5: Estimation results: EU

	France			Spain			U.S.		
	(1)	(2)	(3)	(1)	(2)	(3)	(1)	(2)	(3)
β_2^m	0.010***	0.009***	0.003***	0.065***	0.027***	0.011***	0.0010***	0.000	-0.005***
	(0.000)	(0.000)	(0.001)	(0.001)	(0.002)	(0.002)	(0.000)	(0.000)	(0.000)
β_2^{mc}		-0.000	0.000		0.065***	0.043***		0.003***	0.003***
		(0.001)	(0.001)		(0.002)	(0.002)		(0.001)	(0.001)
β_2^c		0.002***	0.001*		0.015***	0.011***		0.002***	0.002***
		(0.001)	(0.001)		(0.001)	(0.001)		(0.000)	(0.000)
Year FE	YES	YES	YES	YES	YES	YES	YES	YES	YES
Controls	NO	NO	YES	NO	NO	YES	NO	NO	YES
Observations	1935024	1935024	1862061	2680904	2680904	2101557	8739247	8739247	8739247

Notes: Regression of a dummy variable for the transition from unemployment to employment (UE, in column (1)) and from employment to unemployment (EU, in column (2)) on dummies for the migration status, crisis and the interaction term of the last two. Both regressions include controls for education, potential experience, marital status, age, gender, region of residence,occupation, sector of activity, type of job and year dummies. The regression for EU additionally include as controls the type of contract (permanent or temporary), type of job (full-time or part-time), and tenure. Significance levels: $^*p < 0.05$, $^{**}p < 0.01$, $^{***}p < 0.001$. *Source: Spanish LFS-Flows (2005-2018), French LFS (2003-2018) and CPS (2003-2018).*

Table 6: Adjusted predictions and marginal effect

		Probability finding a job (UE)			Probability exiting from a job (EU)		
		Crisis 0	Crisis 1	Marginal Effect	Crisis 0	Crisis 1	Marginal Effect
France	Natives	22.59	20.44	-2.16^{***}	1.26	1.54	0.09^{*}
	Immigrants	18.60	16.27	-2.33^{**}	1.15	1.65	0.14
Spain	Natives	53.65	48.68	-4.98^{***}	6.21	7.31	1.10^{***}
	Immigrants	51.06	34.97	-16.09^{***}	8.16	13.59	5.44^{***}
U.S.	Natives	35.61	31.98	-3.63^{***}	1.18	1.42	0.24^{***}
	Immigrants	34.76	31.38	-3.34^{***}	0.69	1.25	0.55^{***}

*Notes: Adjusted predicted probabilities and marginal effects computed by the linear probability model of Equations (**??**) and (**??**), estimated with all the control variables. The predicted values are evaluated at the mean of the covariates. Significance level:* $^*p < 0.1$, $^{**}p < 0.05$, $^{***}p < 0.01$. *Source: Spanish LFS-Flows (2005-2018), French LFS (2003-2018) and CPS (2003-2018).*

C Labour market stock variance decomposition

This appendix section presents the variance decomposition used in the paper. Let us first recall the relationship between labour market stocks and the associated transition rates.

$$\begin{pmatrix} E \\ U \\ I \end{pmatrix}_t = \begin{pmatrix} 1-p^{EU}-p^{UI} & p^{UE} & p^{UI} \\ p^{EU} & 1-p^{UE}p^{UI} & p^{IU} \\ p^{EI} & p^{UI} & 1-p^{IE}-p^{IU} \end{pmatrix}_t \begin{pmatrix} E \\ U \\ I \end{pmatrix}_{t-1} \tag{16}$$

Normalizing the working-age population to 1 (such that $P_t + G_t + U_t + I_t = 1$), (16) simplifies to:

$$\underbrace{\begin{pmatrix} E \\ U \end{pmatrix}_t}_{s_t} = \underbrace{\begin{pmatrix} 1-p^{EU}-p^{EI}-p^{IE} & p^{UE}-p^{IE} \\ p^{EU}-p^{IU} & 1-p^{UE}-p^{UI}-p^{IU} \end{pmatrix}_t}_{\tilde{P}_t} \underbrace{\begin{pmatrix} E \\ U \end{pmatrix}_{t-1}}_{s_{t-1}} + \underbrace{\begin{pmatrix} p^{IE} \\ p^{IU} \end{pmatrix}_t}_{q_t} \tag{17}$$

The steady-state of the latter system is given by: $\bar{s}_t = (I - \tilde{P}_t)^{-1} q_t$. The evolution of labour market stock can be written as:[10]

$$\Delta s_t = A_t \Delta \bar{s}_t + B_t \Delta s_{t-1} \tag{18}$$

where $A_t = (I - \tilde{P}_t)$ and $B_t = (I - \tilde{P}_t)\tilde{P}_{t-1}(I - \tilde{P}_{t-1})^{-1}$. The first term in (18) captures changes in labour market stock driven by the contemporaneous changes in transition rates that shift the equilibrium steady-state $\bar{s}_t$. The second term captures remaining changes in current labour market stock that are due to past changes in transition rates. Iterating (18) backwards, it is possible to write the present change in labour market stock as a distributed lag function of changes in steady-state values and some initial value for the first observed value:

$$\Delta s_t = \sum_{k=0}^{t-1} C_{k,t} \Delta \bar{s}_{t-k} + D_t \Delta s_0 \tag{19}$$

where $C_{k,t} = \left(\prod_{n=0}^{s-1} B_{t-n}\right) A_{t-k}$, $D_t = \prod_{k=0}^{t-1} B_{t-k}$ and Δs_0 denotes changes in labour market stock observed in the first period of data. Such a representation of the system shows that fluctuations in current labour market stock s_t are governed by changes in the underlying hazard rates h_t^{ij} that affect transition probabilities p_t^{ij} (the elements of A_t and B_t) and the steady state the system is converging at each time period, $\bar{s}_t$. Consequently, to have a mapping between changes in labour market stocks and changes in hazard rates, we take a first-order approximation of the change in steady-state labour market stocks around the lagged value of the flow hazard rates:

$$\Delta \bar{s}_t \approx \sum_{i \neq j} \frac{\partial \bar{s}_t}{\partial h_t^{ij}} \Delta h_t^{ij} \tag{20}$$

With estimates of transition rates and hazard rates in hand, the computation of $\Delta \bar{s}_t$ can be readily obtained by differentiating the continuous-time analogue of the reduced-state Markov chain (17).

[10] See Elsby et al. (2015) appendix for details.

The latter is given by:

$$\dot{s}_t = \underbrace{\begin{pmatrix} -h^{EU} - h^{EI} - h^{IE} & h^{UE} - h^{IE} \\ h^{EU} - h^{IU} & -h^{UE} - h^{UI} - h^{IU} \end{pmatrix}_t}_{\tilde{F}_t} s_t + \underbrace{\begin{pmatrix} h^{IE} \\ h^{IU} \end{pmatrix}_t}_{g_t} \tag{21}$$

The continuous-time expression of the system's steady state is so $\bar{s}_t = -\tilde{F}^{-1} g_t$ and matrix algebra allows us to compute elements of equation (20) analytically.

Using the observed values of Δh_t^{ij} in equation (20), we are now able to obtain time series of counterfactual changes in labour market stocks driven by current and past change in hazard rates. All these elements in hand, combined with the linearity of equation (20), yield to the following decomposition of variance:

$$var(\Delta s_t) \approx \sum_{i \neq j} cov\left(\Delta s_t, \sum_{k=0}^{t-1} C_{k,t} \frac{\partial \bar{s}_{t-k}}{\partial h_{t-k}^{ij}} \Delta h_{t-k}^{ij}\right) \tag{22}$$

Given expression (22), one can compute the share of variance of changes in any given labour market stock accounted for by variations in any hazard rate. As an example, if one were interested by the contribution of changes in the job separation rate (h^{EU}) to changes in unemployment, one could compute:

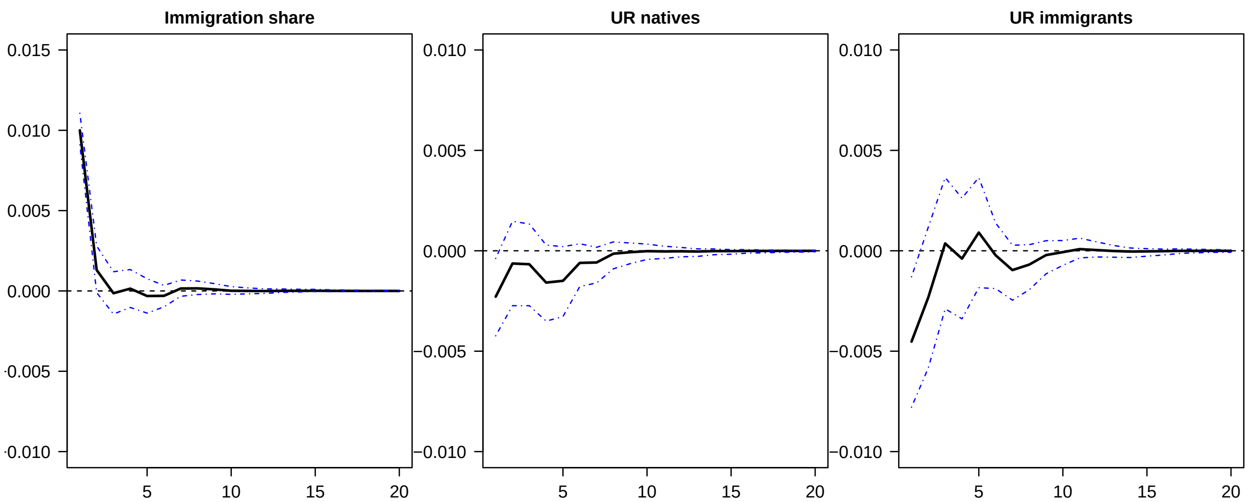
$$\beta_U^{EU} = \frac{cov\left(\Delta U_t, \sum_{k=0}^{t-1} C_{k,t} \frac{\partial \bar{s}_{t-k}}{\partial h_{t-k}^{EU}} \Delta h_{t-k}^{EU}\right)}{var(\Delta U_t)} \tag{23}$$

However, we are more interested in decomposing the variance of the unemployment rate $u_t = \frac{U_t}{E_t + U_t}$ rather than the one of the stock of unemployment. We so use a first order linear approximation to express the changes in u_t in terms of E_t and U_t:

$$\Delta u_t = (1 - u_{t-1}) \frac{\Delta U_t}{E_{t-1} + U_{t-1}} - u_{t-1} \frac{\Delta E_t}{E_{t-1} + U_{t-1}} \tag{24}$$

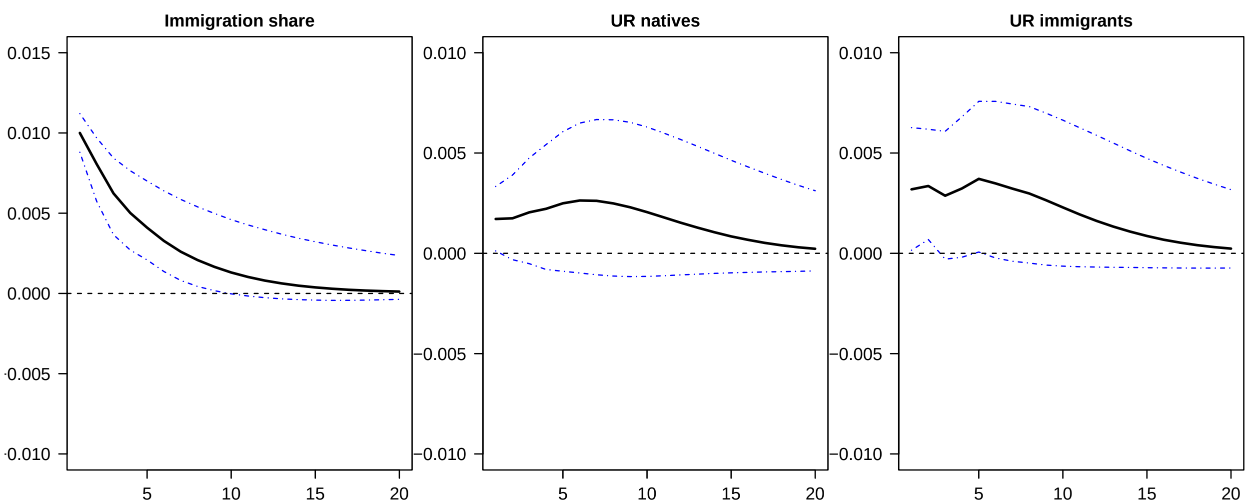
D VAR impulse response functions with unemployment

Figure 8: Impulse response function after an immigration shock in France



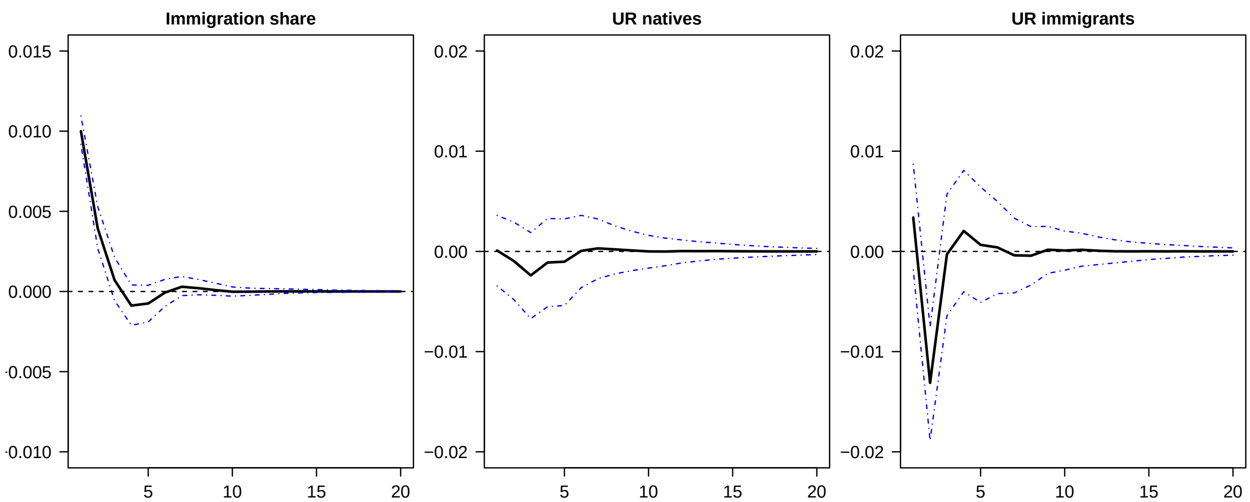
Notes: IRFs to a one percentage point increase in the immigration share are plotted. Confidence intervals are the classical Bayesian bands. Source: French Labour Force Survey (2003-2018).

Figure 9: Impulse response function after an immigration shock in Spain



Notes: IRFs to a one percentage point increase in the immigration share are plotted. Confidence intervals are the classical Bayesian bands. Source: Spanish Labour Force Survey (2005-2018).

Figure 10: Impulse response function after an immigration shock in the U.S.



Notes: IRFs to a one percentage point increase in the immigration share are plotted. Confidence intervals are the classical Bayesian bands. Source: CPS (2003-2018).

E Forecast Error Variance Decomposition after an immigration shock

Table 7: Forecast Error Variance Decomposition after an immigration shock - France

	M_L	M_A	M_U	EU_L^N	EU_A^N	EU_U^N	UE_L^N	UE_A^N	UE_U^N	EU_L^M	EU_A^M	EU_U^M	UE_L^M	UE_A^M	UE_U^M
1	100.0	100.0	100.0	0.11	1.21	4.80	0.07	0.77	3.31	0.71	4.03	10.04	0.48	3.28	8.84
2	74.59	83.21	90.33	0.75	3.02	8.04	0.50	2.03	5.49	1.43	4.83	11.16	1.27	4.39	10.06
4	65.80	74.45	82.10	1.92	4.70	9.58	1.51	3.68	7.58	2.37	5.58	11.01	2.52	5.92	11.46
8	61.12	70.13	78.28	2.63	5.48	10.25	2.49	5.02	9.18	3.28	6.67	11.99	3.21	6.54	11.88
16	59.92	69.25	77.63	2.77	5.64	10.43	2.60	5.16	9.39	3.35	6.74	12.08	3.30	6.65	11.99

Note: M denotes the immigrant share, while EU^j and UE^j denotes, respectively, the job-separation and job-finding rate for natives ($j = N$) and immigrants ($j = M$). The subscript M denotes the median of the posterior distribution, while L and U denotes the 16th and 84th percentiles of the posterior distributions.

Table 8: Forecast Error Variance Decomposition after an immigration shock - Spain

	M_L	M_A	M_U	EU_L^N	EU_A^N	EU_U^N	UE_L^N	UE_A^N	UE_U^N	EU_L^M	EU_A^M	EU_U^M	UE_L^M	UE_A^M	UE_U^M
1	100.0	100.0	100.0	0.08	0.84	3.61	0.12	1.33	5.34	0.08	0.85	3.65	0.08	0.92	3.97
2	90.88	95.32	98.02	0.72	2.77	7.04	0.52	2.15	6.11	0.67	2.62	6.70	0.43	1.78	5.12
4	79.56	87.66	93.18	1.84	5.34	12.51	1.58	4.96	13.10	1.77	4.57	9.66	1.29	3.22	6.91
8	68.62	80.33	88.45	3.11	8.37	19.31	3.41	10.79	26.27	2.67	6.28	13.21	2.15	4.88	9.88
16	63.63	77.70	87.03	3.51	9.55	22.92	4.22	13.30	32.45	2.92	6.87	15.30	2.44	5.60	12.13

Note: M denotes the immigrant share, while EU^j and UE^j denotes, respectively, the job-separation and job-finding rate for natives ($j = N$) and immigrants ($j = M$). The subscript M denotes the median of the posterior distribution, while L and U denotes the 16th and 84th percentiles of the posterior distributions.

Table 9: Forecast Error Variance Decomposition after an immigration shock - the U.S

	M_L	M_A	M_U	EU_L^N	EU_A^N	EU_U^N	UE_L^N	UE_A^N	UE_U^N	EU_L^M	EU_A^M	EU_U^M	UE_L^M	UE_A^M	UE_U^M
1	100.0	100.0	100.0	2.61	7.88	15.36	0.25	2.22	7.13	0.06	0.71	3.06	0.10	1.03	4.22
2	72.99	81.82	89.23	8.26	14.92	23.02	5.92	11.71	19.22	1.03	3.74	8.88	0.55	2.16	5.60
4	61.77	71.53	80.28	8.27	14.09	21.71	6.73	12.03	18.91	2.44	5.65	10.96	2.24	5.18	9.93
8	55.95	66.02	75.30	9.37	15.20	22.84	8.80	14.29	21.34	3.64	7.07	12.43	3.24	6.34	11.12
16	54.23	64.81	74.41	9.36	15.30	23.17	8.89	14.47	21.75	3.84	7.38	12.97	3.44	6.64	11.63

Note: M denotes the immigrant share, while EU^j and UE^j denotes, respectively, the job-separation and job-finding rate for natives ($j = N$) and immigrants ($j = M$). The subscript A denotes the median of the posterior distribution, while L and U denotes the 16th and 84th percentiles of the posterior distributions.

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