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International Migration and Regional Housing Markets: Evidence from France

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Abstract

This article examines the causal relations between immigration and the characteristics of the housing market in host regions. We constructed a unique database from administrative records and used it to assess annual migration flows into France's 22 administrative regions from 1990 to 2013. We then estimated various panel VAR models, taking into account GDP per capita and the unemployment rate as the main regional economic indicators. We find that immigration has no significant effect on property prices but that higher property prices significantly reduce immigration rates. We also find no significant relationship between immigration and social housing supply.

JEL classification: E20, F22, J61.

Keywords: Immigration, Property Prices, Social Housing, Panel VAR.

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1 Introduction

The housing market is a particularly interesting angle for studying immigration in developed countries. It can reveal both the native population's fear of increased difficulty in finding a home, and immigrants' reasons for settling down in a specific locality. Thus, the interactions studied here are similar to those observed in the labor market, but the tensions are potentially greater because the housing market adjusts more slowly. In this paper, we study the relationship between immigration and the housing market at a regional level in France over the period 1990-2013. To our knowledge, this is the first study for the case of France.

There are four main facts that characterize the relationship between immigration and housing in France. First, the foreign-born population is very unevenly spread around the country. According to census data, the three main French administrative regions: Ile-de-France (the Paris region), Rhône-Alpes and Provence-Alpes-Côte d'Azur, are home to one-third of the native population but two-thirds of the foreign population. Second, local housing markets are very mixed in terms of prices: average prices in different regions can differ by as much as 100%. Third, the social housing sector, i.e. rental housing subject to access restrictions and provided at below market prices, is important and represents 44% of the overall rental housing stock in France. However, this social housing is also unevenly spread around the country; for example, 30% of all social housing is in the Paris region. Lastly, the native and foreign populations are differently distributed according to occupancy status of the housing unit: 56% of immigrants live in rented accommodation compared to 34% of the non-immigrant population. The homeownership rate is also particularly low for recently arrived immigrants in France (see Gobillon and Solignac, 2015, for an analysis over the period 1975-1999). Among tenants, the percentages of native- and foreign-born individuals living in social housing are 40% and 52% respectively (see Fougère et al., 2013 for analysis of the statistical association between social housing policy and location choice). These facts illustrate a lack of uniformity both geographically and within the housing market for the immigrant and non-immigrant populations, although they do not necessarily imply sharp segmentation.

The interaction between immigration and local housing markets is theoretically ambiguous. The inflow of immigrants into a region would increase the demand of housing in that region. The effect on prices depends on the supply and demand adjustments. In the basic stock-flow model of the housing sector, housing prices adjust to equalize the changes in demand in the short run given the already existing stock of housing (supply) (see d’Albis and Djemai, 2017). However, the supply of housing itself adjusts to these changes: the stock expands gradually with new building. If housing markets are not regulated, housing prices are expected to positively react to an inflow of immigrants in the short run, while the long run effect would depend on the responsiveness of housing supply to changes in market conditions. In practice, housing markets are often regulated and the adjustment of prices could be constrained, therefore delaying supply adjustment. An additional difficulty lies in the fact that housing conditions could influence immigrants’ choice of location. All else equal, particularly economic conditions, immigrants may choose to settle in a region where housing is more affordable in the first place.

Our aim is to establish causal links between the variables that characterize immigration and those that characterize the housing market. As highlighted above, there is endogeneous interaction between migration and housing conditions. Different approaches were considered to address this endogenous interaction. The first approach relies on natural experiment (Hunt, 1992; Verdugo, 2016) The second approach uses the instrumental variable technique. Because of the persistence in migration flow, one can rely on *internal* instruments, i.e., lagged values as instruments (for example, as in Dustmann et al., 2005). A recent method inspired from trade literature, uses *external* instruments obtained for gravity model prediction, particularly on cross-sectional data (such as in Alesina et al., 2016; Ortega and Peri, 2014). In our case, there is no recent natural experiment and a gravity-based approach on international migration would be difficult to implement at the regional level of the host country. We therefore address the endogeneity by using a vector autoregressive (VAR) approach that brings out the persistence behind the use of lagged values as instruments in the single equation approach. Given the lack of satisfactory external instruments, the VAR approach has been designed to address the reverse

causality issue by allowing a dynamic interaction between variables in the system (Sims, 1980). Moreover, by putting relevant variables in the system, it also addresses the omitted bias issue. VAR models thus have the advantage of analyzing the effects of a shock impacting one variable on other variables of interest, over time. As explained above, this dynamic analysis is a very useful tool. Following Blanchard and Katz (1992)'s groundbreaking article, this method has been used convincingly to assess regional performance. In particular, Zabel (2012) studies the causal links between migration and housing in US metropolitan statistical areas (MSAs). In short, with an appropriate identification in the VAR, we are able to examine the dynamic impact of an exogenous migration shock (i.e., not caused by host economic conditions) on French regional housing and economic variables, and vice versa.

The main difficulty for studying this issue in France is that of data availability. In particular, there were no region-by-region statistical series on international migration flows covering a sufficiently long period. As such, we had to build an original database from the administrative information system of France's local police authorities (*préfectures*), which records all residence permits issued to foreigners. As nationals of the European Economic Area and Switzerland no longer need a residence permit to settle in France, this database does not contain reliable information on European immigrants after the entry into force of the freedom of movement in Europe. Thus, we focus on the legal immigration of third-country nationals (non-EU migrants), who are subject to French immigration policy. The migration flows of non-European nationals are consequently produced on the basis of residence permit statistics collected at the regional level from 1990 to 2013.

We estimated various panel VAR models, including migration flows, housing variables (housing prices and/or social housing), and two variables, for the regions' economic situation (GDP per capita and the unemployment rate) to address the potential omitted bias issue. The impulse response functions show that a positive property price shock does cause a reduction in immigration flows and, conversely, that a positive immigration shock does not cause a property price increase. Our first result might be due to the particular

characteristics of immigration in France, which is mostly family immigration. Tensions in the property market make this kind of immigration more difficult. We have illustrated this by breaking down the migration flows by the person’s gender or country of origin. Our second result could be explained by the segmentation of the housing market. However, the impulse response functions show that our results do not change when taking into account interregional differences in the social housing supply. These results are robust when we include the European migration flows; we checked this for the period 1990 to 2003 during which the EU migration flows are measured. In addition, we show that a more abundant supply of social housing does not cause an increase in immigration. This suggests the absence of a ‘magnet effect’ of social housing, contrary to Verdugo (2016) who showed that it was apparent in the late 1970s. Our results therefore highlight the changes in the nature of immigration and attitudes towards it in France.

The remainder of this article is structured as follows: Section 2 describes the data and more specifically our immigration database; Section 3 presents the econometric methodology; Section 4 details the empirical results and compares them to relevant findings in the literature; finally, Section 5 concludes.

2 Data Description

Our database covers the 22 administrative regions of Metropolitan France, i.e., excluding the French overseas territories, annually over the period 1990-2013. The regional level is “NUTS 2” that follows the European Union’s definition of regional units. The choice of the geographic unit was constrained by the availability of housing data at the local level for France. Nevertheless, more than an administrative division, each region has some degree of political and economic autonomy. Our first contribution was to set up a region-by-region database on international migration, the housing market and economic performance. This is a new and original database on international migration flows to French regions.

More precisely, we constructed regional immigration flows using information collected

in the central foreigners register, managed by the Ministry of the Interior, that is the *Application de Gestion des Dossiers de Ressortissants Etrangers* (AGDREF) (i.e., the application for managing files of foreign nationals in France), which records all residence permits issued in the country. Data are provided by the statistical service of the Ministry *Département des Statistiques, des Etudes et de la Documentation* (DSED), to the French Institute for Demographic Studies (INED). Computerized records of residence permits began in France in 1982. The AGDREF register was created in 1993 to combine the data on residence permits with all other administrative details on permit holders. A more extensive description of AGDREF and its methodology is explained in d’Albis and Boubtane (2015).

Consequently, we take advantage of the automatic data collection of residence permits to build regional gross migration flows of non European nationals.¹ The AGDREF register records all foreigners for whom a residence permit is required to settle in France. therefore, it does not contain reliable information on nationals of the European Economic Area and Switzerland, as they no longer need a residence permit to settle in France. Note that these data do not contain information on foreigners’ departures. Foreign outflows are generally unregulated and pose more measurement problems than legal inflows. Still, the residence permit database is the best available data for France to provide annual harmonized and comparable data on international migration flows.

We use the AGDREF information on the *département* (NUTS 3) to note where the residence permit was issued and at which date the immigrant entered into France to build a regional-level dataset on immigration by gender and nationality. It is, however, not possible to decompose the flows by reasons of issue of the residence permit as this information is not available in AGDREF for the early 1990s. We consider all legal immigrants in France including irregular immigrants who have been regularized. The later are counted when they arrive in France rather than when their status changed, as we use the immigrant’s date of entry to compute flows. The migration flows take into account all adults (aged 18 or over) who received a residence permit valid for one year or more, for the first

¹More precisely, we exclude nationals from EU-27 countries and from Andorra, Iceland, Liechtenstein, Monaco, Norway, San Marin, Switzerland and Vatican for the sample period.

time, over the period. This series of regional migration flows is, to our knowledge, the first to be produced for France.

Following the literature, we use the regional immigration rate computed as the annual regional migration flow as a share of the existing working-age population in the region.

Despite its importance, we know surprisingly little about the evolution of the French regional housing market over the recent decades. Local data are scarce and most available data are related to a limited number of regions over a recent period of time. The only available data on regional housing prices in France is new property prices produced by the statistical department, -*Service de l'Observation et des Statistiques* (SOeS),- of the French Ministry of the Environment, Energy and Marine Affairs (MEEM). New property prices data are annual and available for newly built houses and apartments. More precisely, SOeS publishes the average sale price per square meter of apartments in newly constructed buildings and the average sale price of individual houses. To the best of our knowledge, there is no other regional indicators for real estate prices in France that cover the period 1990-2013 (see the Appendix for alternative data sources and their limitations). Therefore, we use the average sale price per square meter of apartments in newly constructed buildings as a proxy for housing prices at regional levels. We checked that the newly built homes series were strongly correlated with the Notary-INSEE housing prices index in the three regions for which data are available since 1996. Finally, real house prices are given by the ratio of nominal housing price, from Eider database (2015), to the consumers' expenditure deflator, from the INSEE national accounts database.

Social housing stock data are only available at the NUTS 2 level from Eider database (2015) and annual publications of SOeS in its collection -“*Chiffres & Statistiques*”. The stock of social housing adjusts through new construction, sales² or demolitions. What we name below the (net) social housing supply is annual variation of the stock rather than the stock itself, following standard practices in the housing markets literature. Moreover, this

²In France, tenants of social housing have, since 1965, the possibility to buy their dwelling at a discounted price, below market value. Sold units of social housing represented 0.13% of the aggregate social housing stock in 2013.

variable is a good indicator of regional social housing dynamics³. Note that our statistics are not the same as those used by Verdugo (2016), who obtained the information from the population censuses of 1982, 1990 and 1999. We wanted annual time series, which the censuses would not have allowed.

As standard in the literature, to measure economic activity, we use the real regional GDP (chain-linked volume with 2010 as a reference year) divided by working-age population. Scaling GDP by working-age population controls for the fact that regional GDP increases with working-age population size. For simplicity, we refer to this variable as “GDP per capita”. In the same spirit, to account for tension in the regional labor market, we use regional unemployment rate computed as the annual average of the quarterly estimates of the proportion of the labor force that is seeking employment. Data on these variables are from INSEE database.

Table 1: Descriptive statistics

Region	Mig. rate (per 1,000)	Real GDP per capita (2010 prices)	Unemp. rate (in %)	Real hous. price/ m^2 (2010 prices)	Soc. hous. supply (per 1,000)
Alsace	3.08	42052	6.59	2275	0.95
Aquitaine	1.75	39961	8.80	2506	1.13
Auvergne	1.57	37289	8.06	2180	1.10
Basse-Normandie	1.14	37721	8.34	2620	0.82
Bourgogne	1.53	39542	7.99	2190	0.72
Bretagne	1.25	38439	7.59	2327	1.30
Centre	1.93	39982	7.74	2344	0.92
Champagne-Ardenne	1.68	40369	9.16	2279	0.81
Corse	2.96	34507	10.29	2375	0.84
Franche-Comté	2.19	37149	7.54	2081	0.56
Haute-Normandie	1.80	39706	9.99	2392	1.12
Ile-de-France (Paris region)	6.63	66133	7.92	3813	1.49
Languedoc-Roussillon	2.87	35059	12.47	2490	1.52
Limousin	1.67	36895	7.27	2090	0.74
Lorraine	1.88	35678	8.43	2032	0.52
Midi-Pyrénées	2.30	39463	8.65	2436	1.27
Nord-Pas-de-Calais	1.62	35312	12.03	2434	1.36
Pays de la Loire	1.43	39951	7.88	2498	1.09
Picardie	1.68	35861	9.76	2381	1.09
Poitou-Charentes	1.32	37442	8.62	2473	0.78
Provence-Alpes-Côte d’Azur	3.32	42134	11.03	3344	0.94
Rhône-Alpes	2.98	44339	7.94	2680	1.41
France Metropolitan	2.21	39772	8.82	2465	1.02

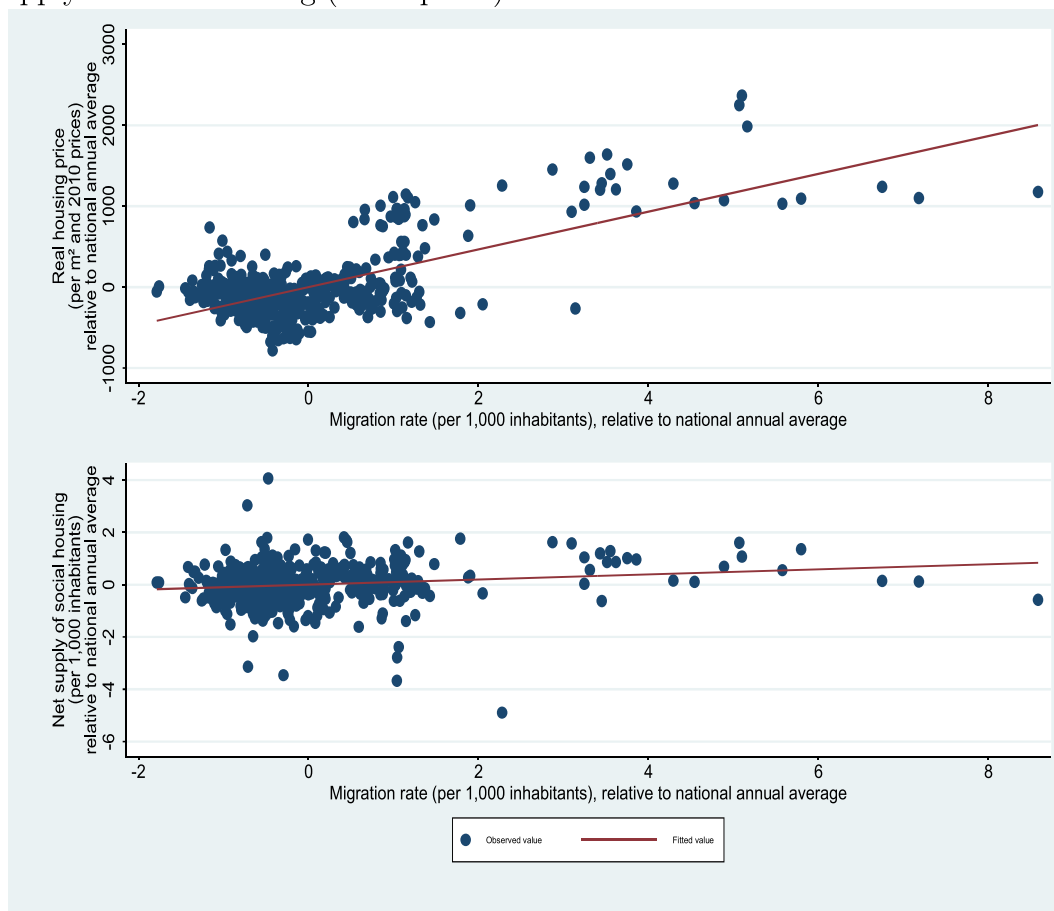
Note: Yearly averages over 1990-2013.

Source: Authors’ computations based on data from INSEE, INED (AGDREF/DSED) and SOeS.

Table 1 shows the mean values of variables considered over the period 1990-2013. It can be easily noted that the Paris region is the one where real GDP per capita, migration rate (migration as a share of mid-year working-age population) and housing

³It should be noted that subsidized dwellings are of uniform size regardless of region or year: the average number of rooms is approximately three per housing unit.

Figure 1: The relationship between migration and housing price (upper panel) and the net supply of social housing (lower panel)



Note: Each scatter plot shows observations by region and by year (1990 to 2013).
 Source: Authors' computations based on data from INED (AGDREF/DSED) and SOeS.

price are the highest, while the Languedoc-Roussillon region has the highest social housing supply per working-age population at the midpoint of the year (1.52 per 1,000 inhabitants on average over the period 1990-2013 compared to 1.49 per 1,000 inhabitants in Ile-de-France). Figures A1-A4 give the evolution over time for our variables expressed as relative to the national annual average, each graph showing changes in the immigration rate and one of the other four variables. The pattern varies widely between regions. In particular, Figure A1-A2 clearly show that the relation between immigration rate and housing variables varies each year across regions. These figures also confirm the observation from Table 1 about the distinctive characteristic of Ile-de-France with regard to immigration flow and housing prices. Here, the immigration rate ranges from 2.24 to 4 times the national average, while housing prices vary between 1.39 and 2.11 times the

national average. It also features an increase in social housing during the recent year. The distinctive feature of the Paris region with regard to immigration flow and social housing supply is consistent with its importance in terms of the corresponding stock variables. The Paris region hosts immigrants and social housing that respectively represents 45% and 27% of the national total.

Figures A3-A4 show that the effects of the 2008 crisis on both GDP per capita and the unemployment rate differ widely by region.

Figure 1 corroborates the variation in migration and housing variables across regions for a given year. It shows the positive local correlations between the immigration rate and property prices, while the correlation between the immigration rate and net social housing supply is almost zero. These correlations remain valid if the data for Ile-de-France are excluded.

3 Empirical Strategy

3.1 The model

Given the time coverage of the data, it is not possible to obtain an accurate analysis for each region. To obtain an average estimation, our empirical analysis is based on a panel VAR model with the following specification:

$$Y_{it} = A_1 Y_{it-1} + \dots + A_p Y_{it-p} + u_i + \lambda_i t + \mu_t + \varepsilon_{it}, \quad i = 1, \dots, N \text{ and } t = 1, \dots, T \quad (1)$$

where $Y_{it} = (y_{it}^1, \dots, y_{it}^K)'$ is a $(K \times 1)$ vector of endogenous variables, the A_s for $s = 1, \dots, p$ are fixed $(K \times K)$ coefficient matrices, $u_i = (u_i^1, \dots, u_i^K)'$ is a fixed $(K \times 1)$ vector of region-fixed effects, $\lambda_i t$ stands for region-specific time trend, μ_t represents the common time-specific effect, and $\varepsilon_{it} = (\varepsilon_{it}^1, \dots, \varepsilon_{it}^K)'$ is the $(K \times 1)$ vector of residuals satisfying $E(\varepsilon_{it}) = 0$, $E(\varepsilon_{it}\varepsilon_{it}') = 0 \forall i$ and t and $E(\varepsilon_{it}\varepsilon_{j\tau}') = \Omega$ for $i \neq j$ or $t \neq \tau$.

Estimating the model in panel form requires the assumption of regional homogeneity in the relationship between variables. To allow such a homogenous assumption for the 22

French regions, we mitigate the regional heterogeneity by including region-fixed effects (u_i) and region-specific time trends (λ_it). We also account for cross-region contemporaneous interdependence by including year-specific effects (μ_t), as in Blanchard and Katz (1992) and Beetsma et al. (2008).⁴

Ordinary least-squares (OLS) estimation with fixed effects does not yield consistent estimates when the number of individuals (N) is large relative to the time series dimension (T). In our study $N = 22$ and $T = 24$, we therefore use the bias-corrected fixed-effects estimator of Hahn and Kuersteiner (2002). This estimator is suitable when T and N are of comparable size i.e., when $0 < N/T < \infty$ (as here), and may be understood as an implementable version of Kiviet's (1995) bias-corrected fixed-effects estimator. Particularly, it can be applied to $VAR(p)$ models with higher order $p > 1$ using the fact that, by imposing blockwise zero and identity restrictions on the VAR coefficients, any $VAR(p)$ process can take a $VAR(1)$ form (Hahn and Kuersteiner, 2002; Lütkepohl, 2005, p. 15).⁵ Moreover, Monte Carlo simulations performed by Hahn and Kuersteiner (2002) show that the efficiency of the bias-corrected estimator measured by the root mean squared error (RMSE) often dominates that of the GMM estimator.

We estimated two models separately. The first one is a four-dimensional VAR model in which the vector of endogenous variables Y_{it} is:

$$\text{Model 1 } Y_{it} = (M_{it}, HP_{it}, GDP_{it}, U_{it})',$$

where M_{it} is the logarithm of the migration rate; HP_{it} is the logarithm of property prices in region i and in year t ; GDP_{it} is the logarithm of GDP per capita; and U_{it} is the logarithm of the unemployment rate. The regional economic variables (GDP per capita and unemployment) are included in the system particularly to address a potential omitted variable bias. The second model is the same as the first except that we add the variable related to social housing supply. This five-dimensional VAR model considers the

⁴Notice that, according to the Frisch-Waugh theorem, including time-specific effects in regression is equivalent to use de-meaned variables (the deviation of variables from the corresponding national annual average), as done in Blanchard and Katz (1992).

⁵See Juessen and Linnemann (2012) for an example of application of this bias-correction in a panel VAR framework.

following vector:

$$\text{Model 2 } Y_{it} = (M_{it}, SH_{it}, HP_{it}, GDP_{it}, U_{it})'$$

where SH_{it} is the logarithm of the net supply of social housing per capita.⁶

The choice of the number of lags in the estimated models was made using AIC (Akaike information criterion) and BIC (Bayesian information criterion) tests. This leads us to select one lag.

After having estimated the VAR coefficients, we now aim to establish causal relationship between variables.

3.2 Identification strategy

In VAR models, one can assess causality with Choleski decomposition. To this end, we order variables from the “least endogenous” to the “most endogenous”. In this decomposition, variables ordered first in the VAR system are allowed to have a contemporaneous impact on the other variables, while variables ordered later in the VAR system can affect those ordered first only with a lag. In other words, we make assumptions about the contemporary impacts of the shocks specifying which variables may be influenced in period t by a change in another variable in the same period t , while no restriction is placed on the variables for dates after t . Precisely, a structural shock, or innovation, on one variable can impact at time t this very variable and the other variables ordered afterwards, and from $t + 1$, all the variables of the system.

For our two models, we consider the following Choleski orderings : $(M_{it}, HP_{it}, GDP_{it}, U_{it})$ for Model 1, and $(M_{it}, SH_{it}, HP_{it}, GDP_{it}, U_{it})$ for Model 2. The choice of these orderings can be justified as follows. In the first model, we put the immigration rate before property price based on the assumption that exogenous shocks to immigration can impact contemporary property prices, while changes in housing prices can at best impact immigration with a lag. This is because the decision to migrate to France is taken before arrival, and

⁶To handle negative values on net supply of social housing (variation in social housing stock) in log transformations, as in practice, we use $\log(1+\text{net supply of social housing per capita})$. After computing the impulse responses, they are transformed to be interpreted as percentage change in the response variable due to a shock representing a one percentage increase in the impulse variable (see Section 4)

the administrative immigration procedure is quite long (a visa has to be applied for and approved). The procedure is even longer for immigrants coming under family reunification provisions because the person bringing in his/her family has to apply beforehand for eligibility. It is also very long for those coming to study in France, as they must first apply to a French university. People coming to study or for family reasons account for half of all migration flows into France (d’Albis and Boubtane, 2015). Second, we put property prices before GDP per capita because of nominal rigidities, as in Iacoviello (2005), and put GDP per capita before unemployment rate, which is often done in macroeconomic models.

We turn to the ordering of the second model, which includes the social housing variable. We have placed social housing after migration for the same reason as for property prices. It is placed before property prices in the decomposition because the time lag between the decision to build and the delivery of the building is much longer for social housing than for private homes. This is because the administrative procedure is more cumbersome and the average building size is larger. Of course, we cannot exclude the possibility that immigrants may anticipate the relative economic situations in different regions and that this may affect their choice of one region over another, but we do not think this is a realistic argument, if only because regional economic data are hard to come by. However, we did recalculate our impulse response functions for alternative ordering and found the results unchanged.⁷

3.3 Stationarity properties of series

To choose the appropriate VAR model (VAR in level or in first difference), we first consider the stationarity properties of the series. To this end, we use the second generation panel unit root test developed by Pesaran (2007) that accounts for cross-sectional dependence. This methodology, with the null hypothesis of the presence of a unit root in all series, is based on augmenting the usual augmented Dickey-Fuller (ADF) regression with the lagged cross-sectional mean and its first difference to capture the cross-sectional

⁷Details are available from the authors upon request.

dependence.

Table 2: Panel unit root test

Variables	CIPS	P-value
Log(real GDP per capita)	-2.713**	0.021
Log(unemployment rate)	-2.277***	0.006
Log(migration rate)	-2.714**	0.021
Log(housing price)	-2.734**	0.016
Log(1+net supply of soc. hous. per capita)	-3.741***	0.000

Notes: The statistic test is the CIPS truncated version of Pesaran (2007). The test has the null hypothesis of the presence of a unit root. For all series, one lag is introduced to allow for serial correlation in the errors. For the unemployment rate, only the intercept is included, for the other series, intercept and trend are included. **, *** denote the significance at 5%, 1% levels, respectively.

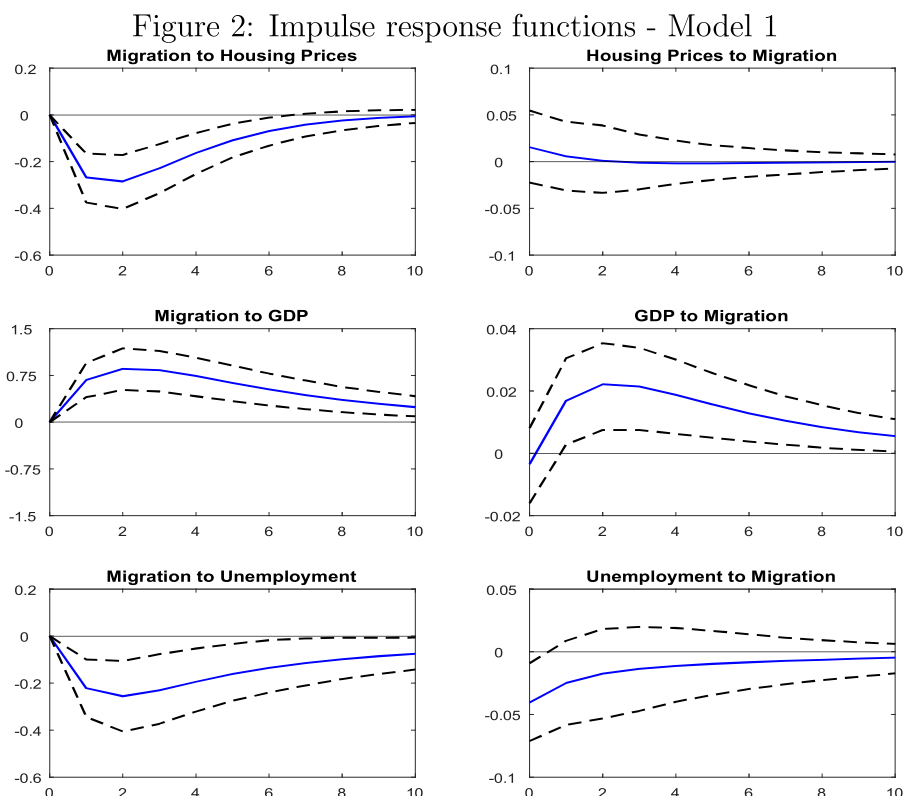
The results of the panel unit root test of Pesaran (2007) are reported in Table 2. These results show that any series considered does not have a unit root. In particular, (the logarithm of) real GDP per capita follows a trend stationary process. This finding is in line with that of Carrion-i-Silvestre et al. (2005) who obtained evidence that points to the trend-stationarity of GDP per capita in a set of 15 OECD countries (including France) from 1870 to 1994, once cross-sectional dependence and breaks in the series are considered. The result that (the logarithm of) the unemployment rate is mean-stationary supports the natural rate hypothesis in French regions. The migration rate, housing prices and social housing net supply per capita (in logarithms) are characterized by a trend stationary process. The trend-stationarity property of housing prices was also found by Kuethe and Pede (2011) for US state-level quarterly data over the period 1988-2007.

4 Results

As all variables in the system are found to have mean or trend-stationarity, we can set a VAR model taking all the variables considered in levels while controlling for regional heterogeneity (using region-specific effects and region-specific time trends) and regional interdependence (using year-specific effects).

We will examine at the two models in turn. The first model examines the relationship between the immigration rate, property prices, unemployment rate and GDP per capita, while the second adds social housing supply per capita. Figures 2 and 4 show the impulse

response functions (IRFs) obtained from the estimation of the two models. The estimated impulse responses are expressed as percentage change in the response variable due to a shock representing a one percentage increase in the impulse variable. The 90% confidence intervals are generated by Monte Carlo with 5,000 repetitions.

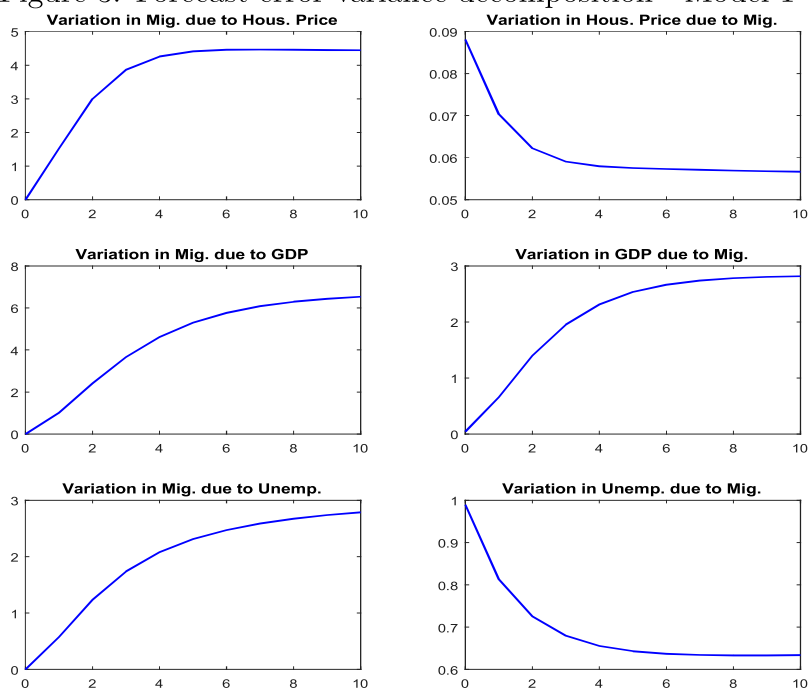


Notes: The solid line shows the estimated percentage change in the response variable due to a shock representing one percentage increase in the impulse variable. The dashed lines indicate the 90% confidence intervals that are generated by Monte Carlo with 5,000 repetitions.

4.1 Immigration and housing prices

The relationship between the immigration rate and property prices is particularly interesting. Immigration reacts significantly and negatively to property prices whereas property prices do not react significantly to immigration rates. First, this finding shows that the positive correlation seen in Figure 1 is not robust when one controls for the two economic variables (GDP per capita and unemployment) and when the endogeneity between the variables is taken into account. This significant relationship is a negative one, and it highlights the lasting effect of property prices on immigration rates. As seen from Figure 2, the effect is significant for six years, from the year of the shock to the

Figure 3: Forecast error variance decomposition - Model 1



Note: The forecast error variance decomposition reports the percentage of the fluctuation in one variable that is attributable to the shock to another variable, accumulated over time.

fifth year afterwards. In response to a shock that increases the property price by 1%, the migration rate decreases by 0.27% after 1 year and by 0.28% (at the peak, after 2 years).

The importance of property price in explaining migration variation is corroborated by the forecast error variance decomposition (Figure 3) that indicates the percentage of the variation in one variable that is caused by the shock to another variable, accumulated over time. Figure 3 shows that, over the 10 years, property prices explain approximately 4.5% of the variation in regional migration, while the contribution of migration to property price volatility is nearly zero.

As Ile-de-France (Paris region) is an important region in France with regard to immigrant population (45%), as a first robustness analysis, we control for regional heterogeneity by estimating Model 1 without Ile-de-France. The corresponding impulse responses are reported in Figure A5. As shown by this Figure, excluding the Paris region does not alter our findings.

Because of data limitations, we were not able to include the regional migration of Europeans over the sample period. For another robustness check, we nevertheless compute

these flows for the period 1990-2003 and run a new model. Our results are robust to the introduction of European migration. Most notably, property prices do not respond significantly to either EU migration or non-EU migration over the period 1990 to 2003.⁸

Discussion

Comparing estimates across studies is problematic because of differences in the countries considered, differences in methodological approaches, various levels of the data aggregation and the periods cover. That being said, some of our results are in line with those of Akbaria and Aydedea (2012) who found that recent immigrants have no impact on average housing prices in Canada. These authors used panel data based on 258 Census Divisions across Canada for three census years from 1996 to 2006 and estimated the effect of immigration on average housing prices. Moreover, Stillman and Maré (2008) also found no evidence that the inflow of foreign-born immigrants is positively related to regional housing prices in New Zealand. They examined how international migration affects rents and sales prices in different local labor market areas in New Zealand. While the return of the native-born expatriates is associated with the increase of local housing prices, their results do not provide evidence that foreign-born immigrants increased these prices. More recently, employing a VAR approach on Norwegian quarterly data over the period 1990-2014, Furlanetto and Robstad (2016) find that international migration has no impact on housing prices. Furthermore, the influence of housing markets on migration decision was examined by Zabel (2012), who estimated a panel VAR model of 277 US metropolitan areas over the period 1990-2006. He showed that the local housing market is an essential determinant of migration responses to labor demand and supply shocks.

However, our results differ from those of other studies. In particular, Saiz (2007) found that a 1% increase in the immigration rate causes a rise in housing values of approximately 1% in the United States. He used, as we do, administrative data on immigrants admitted legally to the US but a different methodology. He estimated the impact of immigration on rents and housing prices using the instrumental variables approach on yearly data at the MSA level over the period 1983-1997. This positive impact of immigration on housing

⁸IRFs are available upon request.

prices can also be found in other studies using the same instrumental variables approach: Degen and Fischer (2009) for Switzerland and Gonzalez and Ortega (2013) for Spain.

It should be noted that some studies using more disaggregated data at the local level (e.g., Saiz and Wachter (2011) for the US and Sá (2015) for the UK) find that immigration has a negative effect on housing prices. As a large proportion of non-European migrants arriving in France are from a French speaking country, our results are consistent with the finding of Fischer (2012), who shows that immigrants coming from a common language country have statistically insignificant impact on Swiss house prices. For research conducted on France, Sotura (2013) examined the resale property market in Paris between 1993 and 2008. Using an exhaustive database of property transactions over the period, she shows that foreign buyers pay more for their homes than do French buyers but that the effect on prices is negligible. She also highlights that it is primarily foreign buyers living abroad who push prices up, while foreign buyers resident in France have little impact on prices.

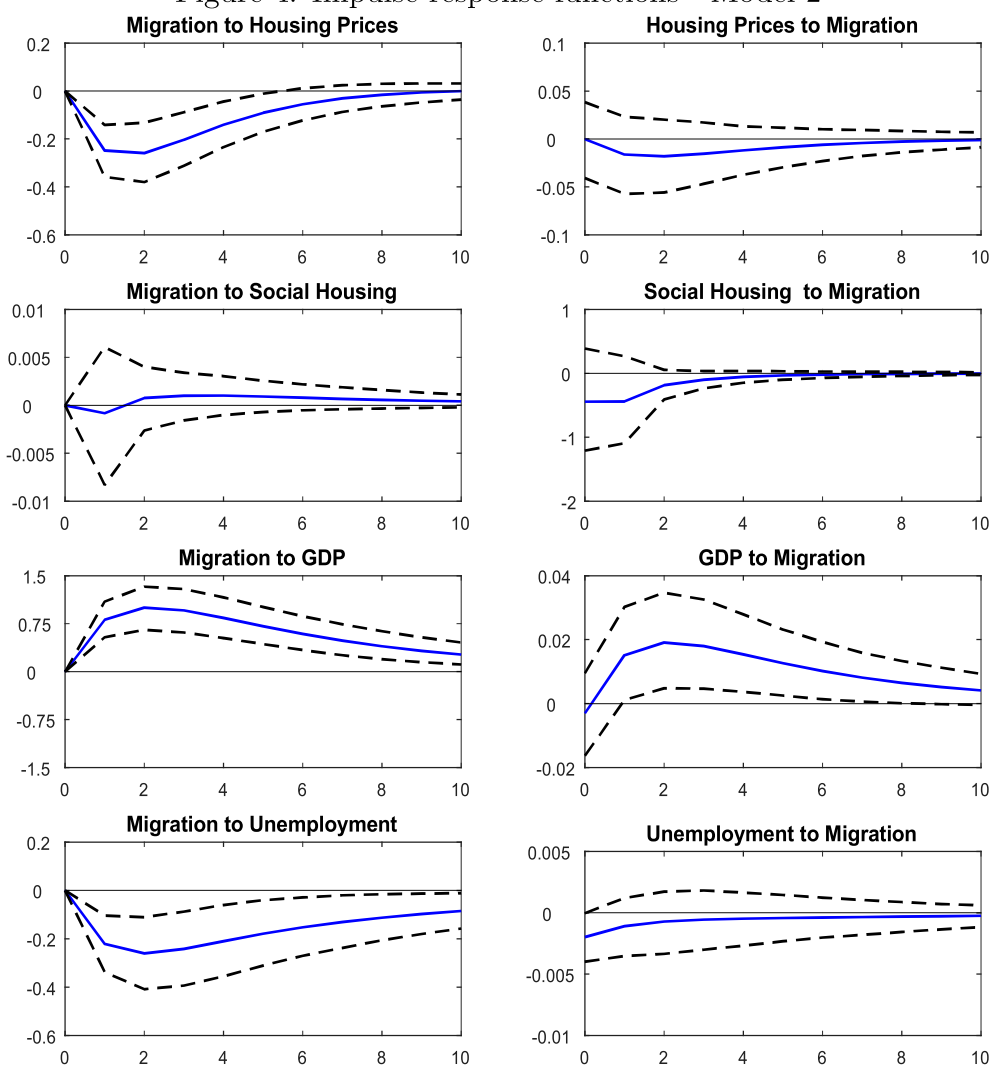
The impact of housing prices on immigration can be explained by the particular characteristics of migration. In France, international migration is mainly for family reasons, and immigrants have to meet adequate housing requirements to bring their families. As shown in d'Albis and Boubtane (2015), family migration is the largest category of migration, representing more than 50% of the flows. Immigrants may be spouses of French nationals or families of foreign residents in France arriving under the administrative procedure for family reunification. To be eligible for this procedure, a foreign resident must have a certain level of resources and a home sufficiently large for the family when they arrive. These conditions are probably harder to meet in regions where property prices are high, and this would reduce migration flows into these regions. To test this intuition, we estimated a new model, which differentiates between male and female immigrants. The immigration flows of men and of women are considered separately, while the rest of the model is unchanged. In the Choleski decomposition, we placed the male migration rate before the female rate because a majority of female immigrants come for family reasons (d'Albis et al., 2016; Mazuy et al., 2016). The impulse response functions we obtained can

be seen in Figure A8 in the Appendix. First, it is clear that the main results (significant negative effect of property prices on immigration and non-significant effect of immigration on property prices) are robust. Second, we see that the magnitude of the reactions to a housing price shock differ across genders, the reaction of the female migration rate being larger than that of the male migration rate (-0.26 and -0.33 at the peak, respectively, for male and female migrations). This suggests that the housing conditions that immigrants must provide to bring their families seem to explain the impact of property prices on immigration rates. Further evidence can be provided by decomposing the flow of immigrants according to their nationality. We have estimated a new model that considers separately the flow of immigrants who are nationals of a high-income country and the flow of those who are nationals of a low-income country. Immigration for family reasons is much more likely in the latter group (Mazuy et al., 2016). The results are reported in Figure A9 in the Appendix. We immediately see that immigrants from low-income countries react to housing prices, whereas immigrants from high-income countries barely react to them.

While the link is hard to establish empirically because of the limitations of the data, the idiosyncrasies of the French housing market could explain at least in part the absence of the impact of immigration on housing prices. First, according to OECD estimates, there is no aggregate shortage of housing in France: the stock of housing per inhabitant is, in particular, higher than that of the US (Andrews et al., 2011). Second, France is one of the OECD countries where the housing market reacts least to shocks. This can be explained by several reasons. First, the information on prices is poor, with data on prices usually coming with a delay and only at a rather aggregate level. Second, France's housing market is highly segmented with 76% of the housing stock as either owner-occupied (58%) or rented for social purposes (19%). The social housing sector accounts for 44% of the French rental market on average over the period 1990-2013, with significantly lower rents than that of the private sector (Trevien, 2013 evaluated the rent gap between comparable private and social dwellings in France at 261 euros per month). This social housing sector share is 3.3 times higher in France compared to the

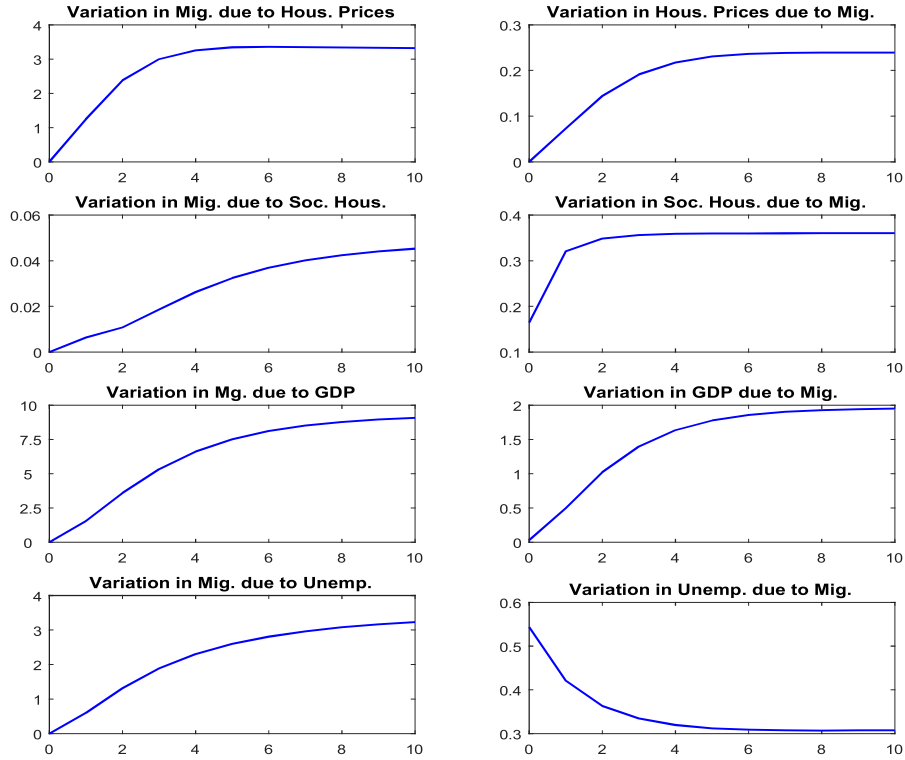
US (Andrews et al., 2011). However, the private rental sector is also highly regulated with stringent rent control and stricter tenant-landlord regulations. Finally, compared to the US, transaction costs in France are particularly high at, approximately 14% of the property value, almost 10 percentage points higher than that in the US. The segmentation of the market could explain the absence of immigration effect on property prices but, as we shall now show, this does not implies an interaction between social housing and immigration.

Figure 4: Impulse response functions - Model 2



Notes: The solid line shows the estimated percentage change in the response variable due to a shock representing one percentage increase in the impulse variable. The dashed lines indicate the 90% confidence intervals that are generated by Monte Carlo with 5,000 repetitions.

Figure 5: Forecast error variance decomposition - Model 2



Note: The forecast error variance decomposition reports the percentage of the fluctuation in one variable that is attributable to the shock to another variable, accumulated over time.

4.2 Social housing

Figure 4 shows the impulse response functions for the second model, which includes the supply of social housing. These functions give two main results. First, the estimated relationships between immigration, property prices and economic variables are unchanged. In particular, the effect of property prices on immigration rates is still significant for the period running from the year of the shock to the fifth year afterwards. Second, there is no significant relationship between net social housing supply and immigration. In particular, we do not find that a more abundant supply of social housing in a region attracts immigrants to that region. This is confirmed by the forecast error variance decomposition reported in Figure 5. Over the 10 years, the contribution of social housing (migration) to variation in regional migration (social housing) is insignificant.

Two robustness checks can be provided using Model 2. First, we control for regional heterogeneity. As Paris region is also important in terms of social housing (27%), we exclude it and estimate again Model 2. As shown in Figure A6, our finding remains

unchanged. Second, we account for the fact that household composition may differ across regions. We consider the system described in Model 2 by replacing the variable social housing supply per working-age population by the social housing supply per household. The impulse response functions in Figure A7 show that our results are unchanged when we consider this alternative measure of social housing.

Discussion

In our case, the social magnet hypothesis (Borjas, 1999) does not seem to be supported. This hypothesis has given rise to quite an abundant literature. For France, a recent study by Verdugo (2016) seems to show the opposite effect to the one we found. He considers that a 9% increase in social housing supply in a given urban area increases the probability that an immigrant will settle there by 40%. Apart from questions regarding data (Verdugo used the censuses), spatial coverage (he only studied major towns) and methodology (he used a change in the social housing allocation rules), the main difference between our study and Verdugo's is the period considered. Verdugo's analysis covers the period from 1975 to 1982, whereas we analyzed the period from 1990 to 2013. Both the pattern of migration and the social housing stock underwent profound changes in France in those two decades. Today, the social housing magnet theory does not seem to be empirically supported; on the contrary, there seems to be discrimination against non-European foreigners in the allocation of social housing. Bonnal et al. (2102) show that non-European foreigners on housing lists have to wait significantly longer than others before being allocated a home, despite the legal obligation to base social housing allocations solely on welfare criteria.

4.3 Regional economic performances

The impulse response functions reproduced in Figures 2 and 4 also indicate a significant response of migration to regional economic performances. Specifically, migration responds positively to GDP per capita and negatively to unemployment rate. The response of the migration rate to GDP per capita is positive and significant for at least 10 years after

the shock, whereas the response to unemployment is negative and significant over the same period. Conversely, immigration has a positive impact on GDP per capita and no significant impact on the unemployment rate. The positive impact on GDP per capita is significant from one year to six years after the shock. As shown in Figures 3 and 5, the variance decomposition analysis bears out the importance of regional economic performances in explaining variation in migration. Over the 10 years, the share of the fluctuation in regional migration that is attributable to regional GDP per capita and unemployment rate are respectively 6.44–8.96% and 2.74–3.17%.

Discussion

The results regarding the relationship between migration and regional economic performances confirm our earlier findings in d’Albis et al. (2016). By estimating a VAR model on monthly data for France (at country level) over the period 1994-2008, we showed that immigration flows significantly respond to France’s macroeconomic performance (positively to the country’s GDP per capita and negatively to its unemployment rate) and that immigration itself increases France’s GDP per capita. These results are thus robust not only to a change in the coverage (regional instead of national) and the database (taking into account the date of arrival instead of the date of issue of the residence permit) but also to variables characterizing the housing sector. From this standpoint, France does not seem to differ much from other countries where studies have been conducted with a similar methodology, i.e., in countries with immigration data of sufficiently long-period coverage. Using annual data between 1930 and 2002 at country level for Australia, Canada, and the United States, Morley (2006) found some evidence of long-term causality between immigration and macroeconomic conditions. With Norwegian quarterly data over the period 1990-2014, Furlanetto and Robstad (2016) find that migration has a positive impact on GDP and wages. Moreover, Withers and Pope (1985) for Australia and Islam (2007) for Canada found interactions between migration and unemployment that are similar to ours. In contrast, for the US, Kiguchi and Mountford (2013) found a temporary negative impact of unanticipated shocks to the labor force while Weiske (2016) found no signif-

icant impact of immigration on per capita variables. Employing panel VAR approach on regional data from Spain from 1999 through 2007, Amuedo-Dorantes and De la Rica (2010) find results indicating that immigration flow reduces regional employment rate disparities only temporarily.

Evidences based on cross-country panel data estimations are mixed, however. Ortega and Peri (2009) estimated a gravity model using data on 14 OECD countries over the period 1980-2005 and found that immigration had no effect on GDP per capita. Dolado et al. (1994) and Boubtane et al. (2016) estimated an augmented Solow model on cross-country OECD panel data and found that the relative magnitude of the capital dilution with that of the increase in human capital depends on the period of consideration. Our results reinforce the recent studies of Alesina et al. (2016), Ager and Brückner (2013) and Ortega and Peri (2014) who found that immigration promotes total factor productivity by increasing diversity in productive skills.

5 Conclusion

A simple analysis of French regional data over the period 1990-2013 exhibits a positive association between immigration rates of non-Europeans and property prices. In this paper, we estimated a panel VAR model and showed that higher immigration flows do not cause an increase in property prices, while higher prices do cause a reduction of immigration flows. Contrary to what would be suggested by descriptive statistics, the relationship is negative. These results differ from some studies conducted for other countries, possibly because of the particular nature of immigration in France, which is mostly motivated by family reasons. Administrative conditions that are necessary to obtain a residence permit for a family notably rely on housing conditions. Those requirements are obviously more difficult to fulfill in regions where housing markets are tense, which consequently reduces immigration. In d'Albis et al. (2016), we showed that family immigration enhanced France's economic performance, as measured by GDP per capita. This suggests that administrative barriers towards family reunification that are linked

to housing conditions are counterproductive. Social housing, which amounts to 44% of rented housing in France, could possibly play a role. We nevertheless showed that our results were not modified when social housing was taken into account. A larger supply of social housing does not affect the impact of property prices on immigration. Moreover, we showed that social housing does not impact location decisions of non-European immigrants. Note that we were not able to discuss the regional implications in France of the freedom of movement of European-citizens, because of data limitations. A potential extension of this work would be to focus on the Paris region, which accounts for 27% of the country's stock of social housing and 45% of the country's stock of immigrants. Housing market differences across the Paris region are tremendous and therefore could be useful to further understand the relationship between immigration and housing markets.

Data appendix

In France, the principal sources of migration data are population censuses and the residence permits database. Although French decennial population censuses organized by the French National Institute of Statistics and Economic Studies (INSEE) provide a rich dataset on migration stock for census years, we require annual data on immigration for the purposes of this research. It should be noted that the residence permits database from AGDREF was exploited first by the French Institute for Demographic Studies (INED), which has produced immigration flows statistics at the national level since 1994 (see Thierry, 2001 for the presentation of the database). The French Ministry of Interior publishes annual statistics (whatever their duration of validity) on residence permit issued since 2007.

The data on new property prices, used in this paper, came from a survey on the sale of new homes, the *Enquête sur la Commercialisation des Logements Neufs* (ECLN) which is a part of the annual program of public service statistical surveys in France. Responding to the ECLN is mandatory, and property management companies have to pay an administrative fine if they do not respond or if they transmit inaccurate information. It should be noted that the statistical department of the French Ministry of the Environment, Energy and Marine Affairs (SOeS) publishes data on new propriety prices at both the NUTS 2 regional level and the national level (Eider database, 2015).

To the best of our knowledge, there are no available annual regional data on resale housing or on rents in France over the period 1990-2013. The resale housing price in France, the Notary-INSEE housing price index, has been published for three regions (Ile de France, Provence-Alpes-Côte d'Azur and Rhône-Alpes) since 1996 and for Nord-Pas-de-Calais since 2007. The third version of the Notary-INSEE index is based on hedonic adjusted prices, from data on transaction. It was not until 1983 that the INSEE calculated a price index for the existing property in the city of Paris. The series was based on transactions registered in Paris notarial offices. To cover the rest of France, two agreements were signed in 1998 and 1999 between INSEE and the Higher Notary Council (*Conseil Supérieur du Notariat*) to produce local housing price indexes for the

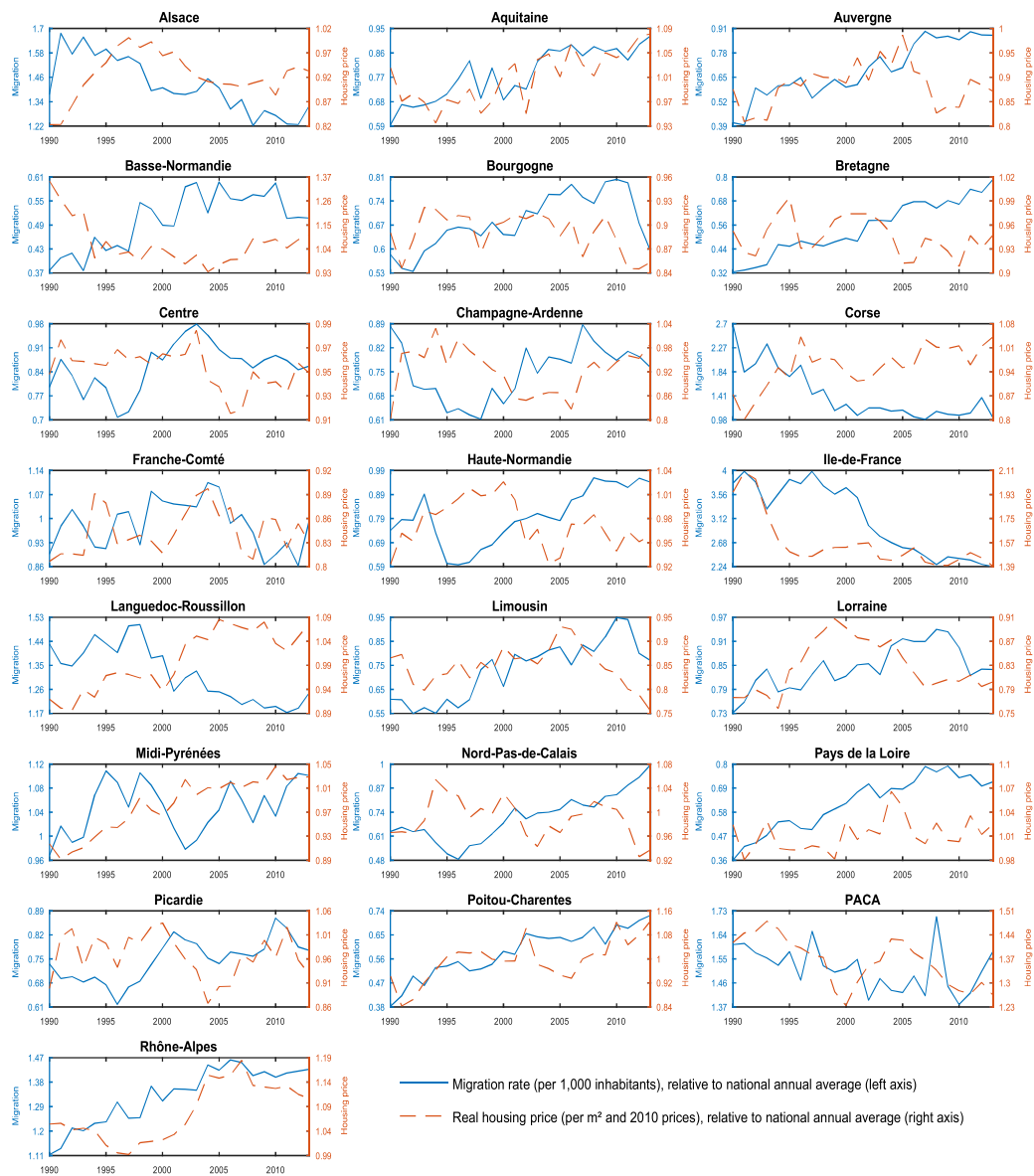
whole French territory (see David et al. (2002)). It should be noted that the data on transactions from the notary database is the main source used by housing market analysts in France. However, the geographic and temporal coverage of these data is very limited. Precisely, transactions data of Paris notarial offices (*Base d'Informations Economiques Notariales* - BIEN) cover all the departments of Ile-de-France (the Paris region) since 1996 only. The data coverage of the Higher Notary Council database (PREVAL) is even more reduced particularly in the 1990s. Clarenc et al. (2014) estimate that the notary database covered only 56% of provincial France in 2010. Moreover, the collection and transmission of transaction data by notarial offices is voluntary and data gathering was not part of the public service mission of the notaries until 2011.

Data on social rental housing stock is from the statistical department of the French Ministry of the Environment, Energy and Marine Affairs. Social rental housing units are managed by the social housing providers (*bailleurs sociaux et sociétés d'économie mixte*), which are non-profit organizations that provide housing at affordable rents (*HLM - habitation à loyer modéré*). In France, social housing is allocated to eligible tenants through local administrative procedures. There is a queuing system in each administrative locality with consideration given to some priority-rated households (particularly the vulnerable households who have waited a long time for a social housing). To our knowledge, data on the local demand for social housing are not collected by the SOeS, and no time series data are available on the evolution of the demand for social housing in France even at the national level. It should be noted that INSEE collects information on social housing demand in a survey on housing conditions named *Enquête Logement*. However, the survey is only conducted every 4 to 5 years, and it is representative at the national level. An exception is the last wave (2013 survey), which was conducted on additional samples for some localities (Ile-de France, Nord-Pas-de-Calais, oversea territories and sensitive urban zones).

Social housing stock data used in this paper are based on two surveys, *le répertoire du parc locatif des bailleurs sociaux* (RPLS) until 2010 and then *l'enquête sur le parc locatif social* (EPLS). The survey is conducted annually among social housing providers whose

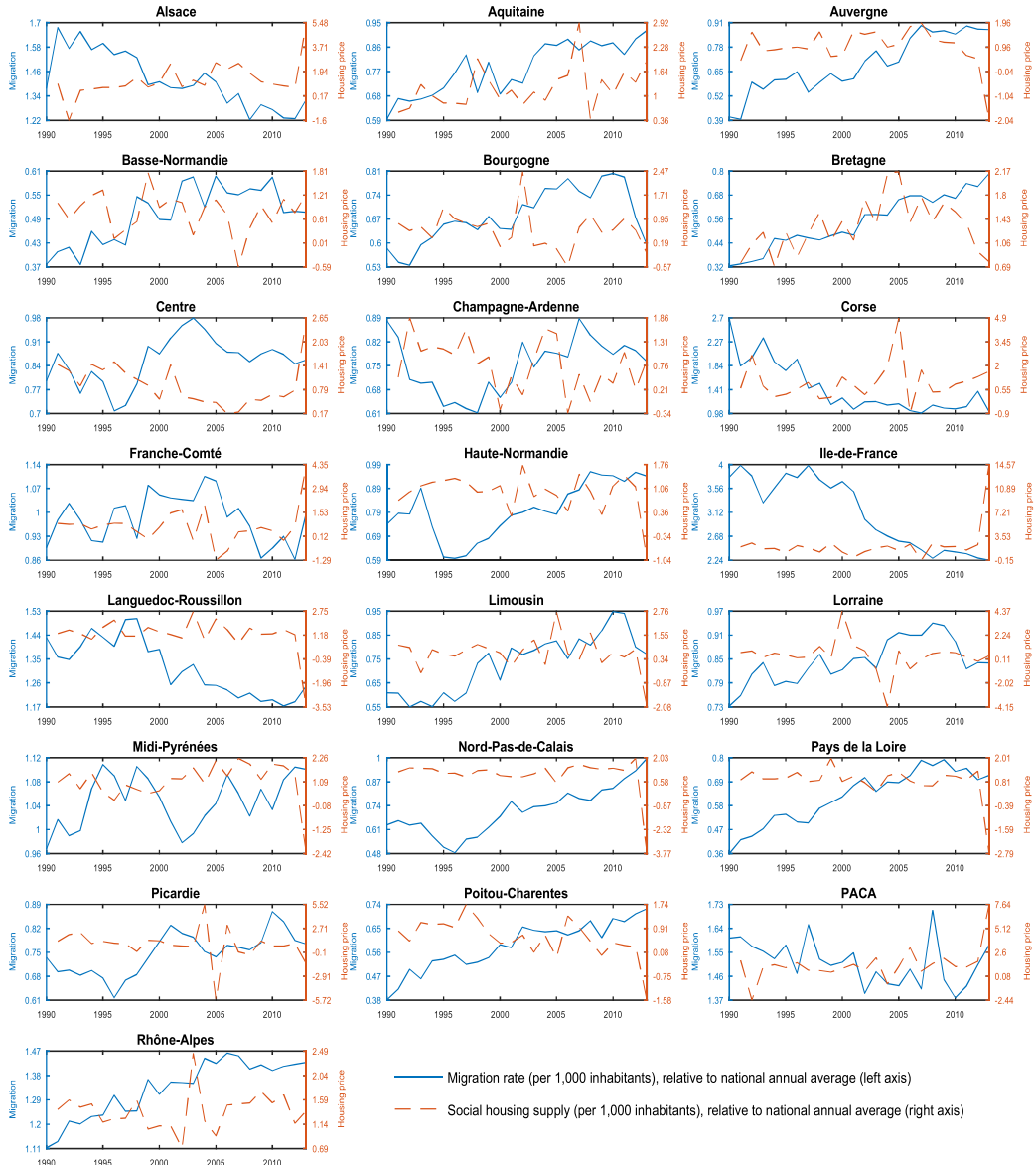
participation is mandatory.

Figure A1: Migration rate and housing price



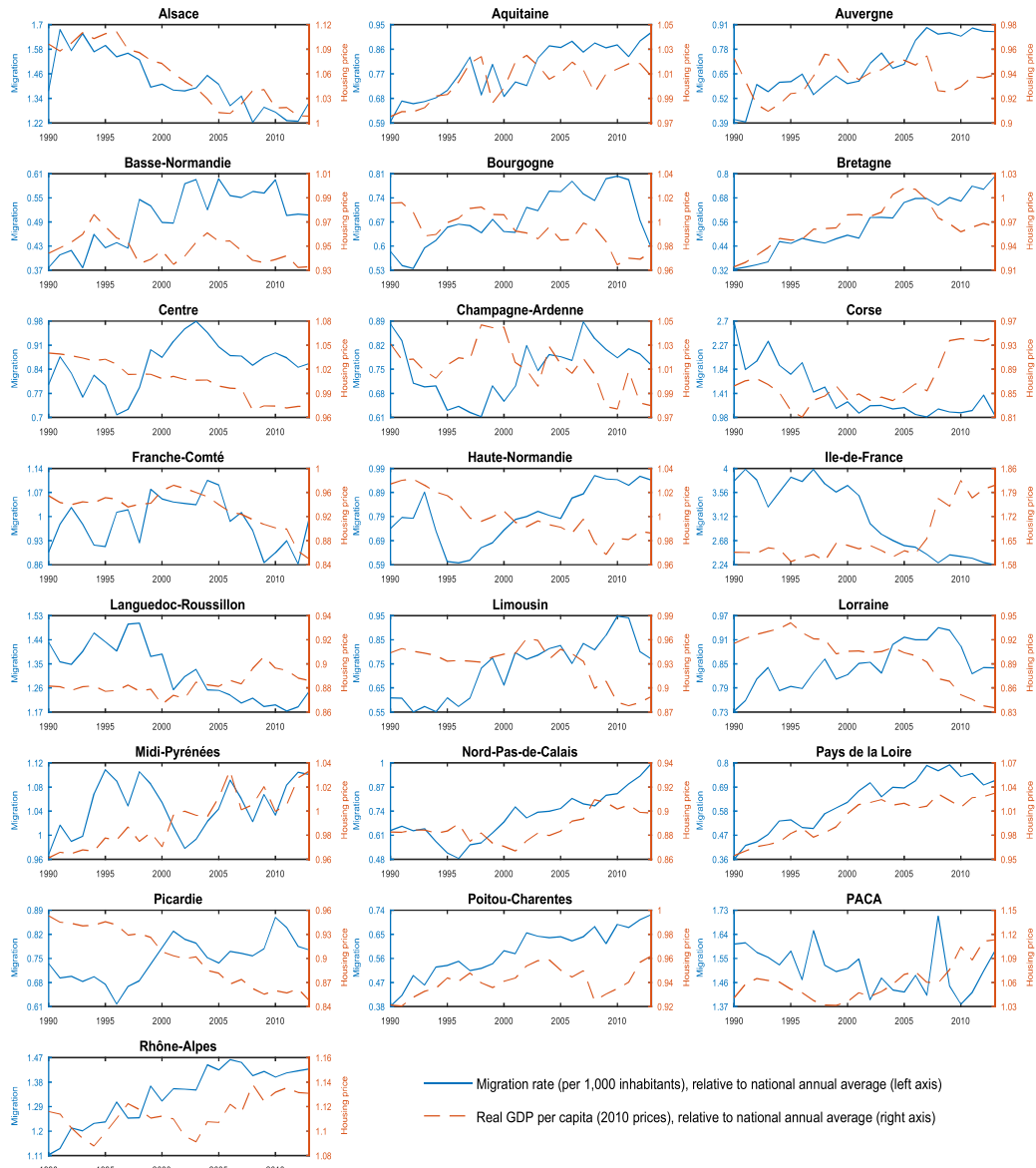
Source: Authors' computations based on data from INED (AGDREF/DSED) and SOEs.

Figure A2: Migration rate and net supply of social housing



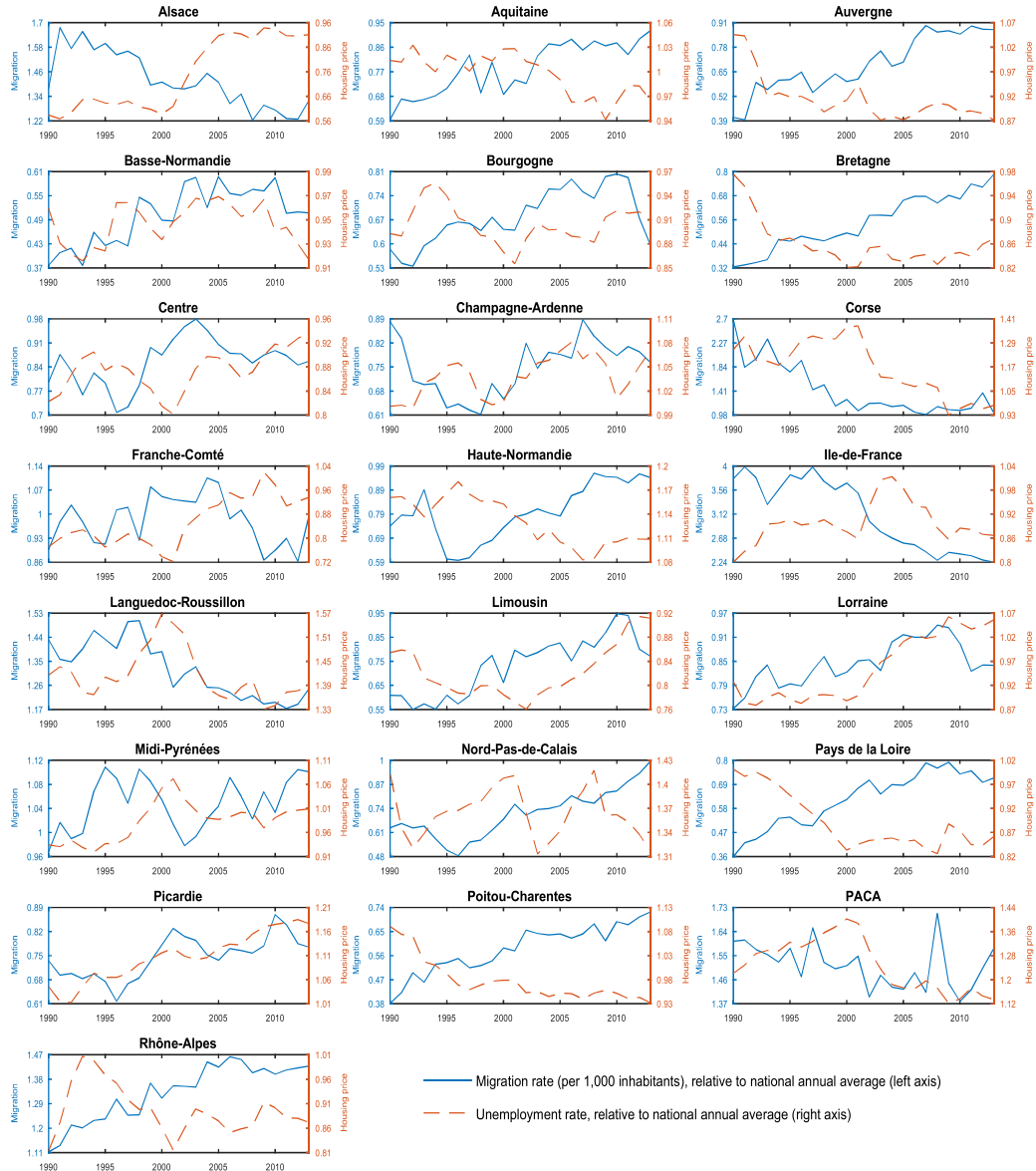
Source: Authors' computations based on data from INED (AGDREF/DSED) and SOeS.

Figure A3: Migration rate and real GDP per capita



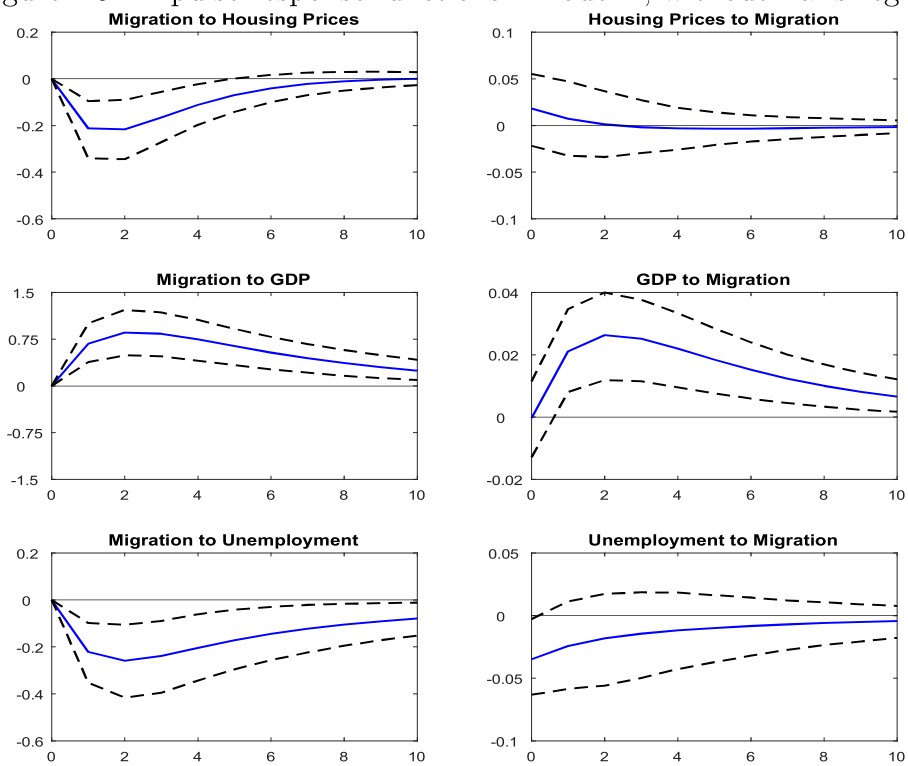
Source: Authors' computations based on data from INSEE and INED (AGDREF/DSED).

Figure A4: Migration rate and unemployment rate



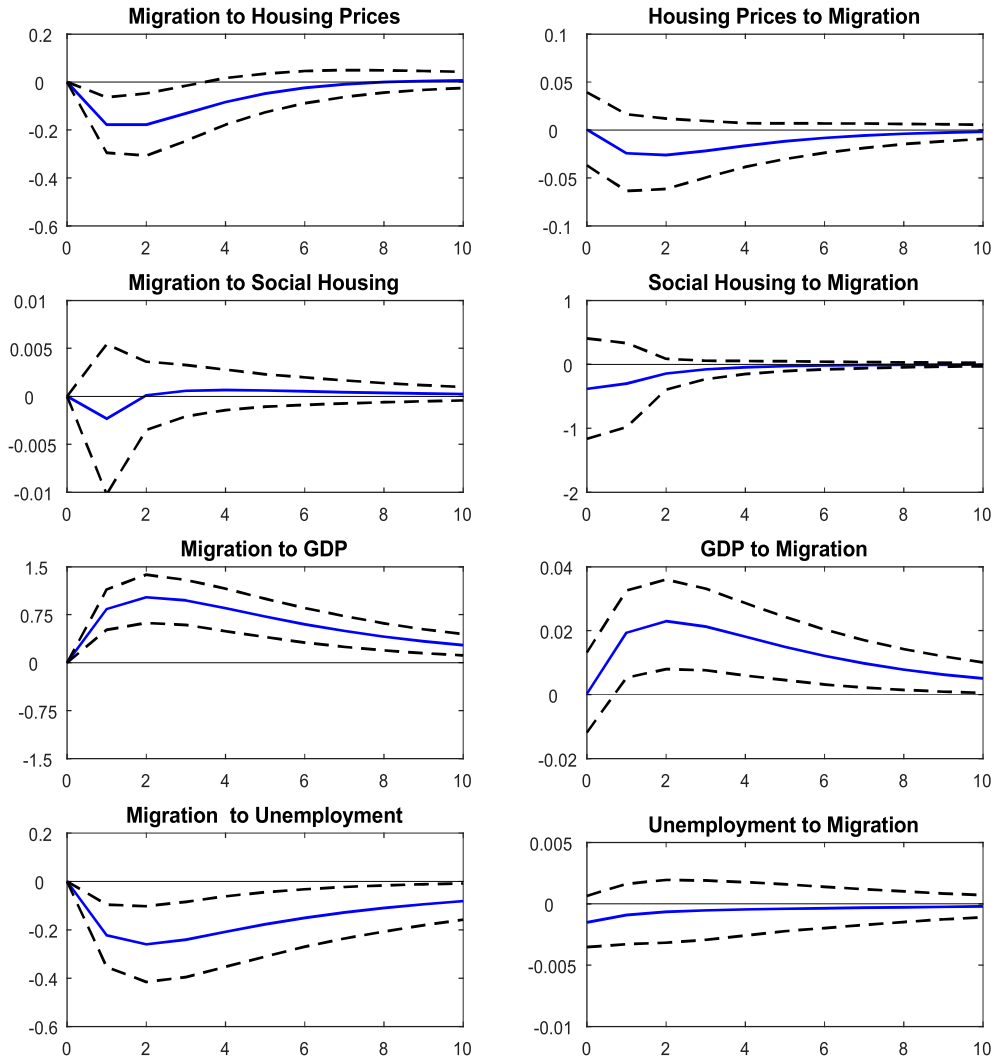
Source: Authors' computations based on data from INSEE and INED (AGDREF/DSED).

Figure A5: Impulse response functions - Model 1, without Paris region



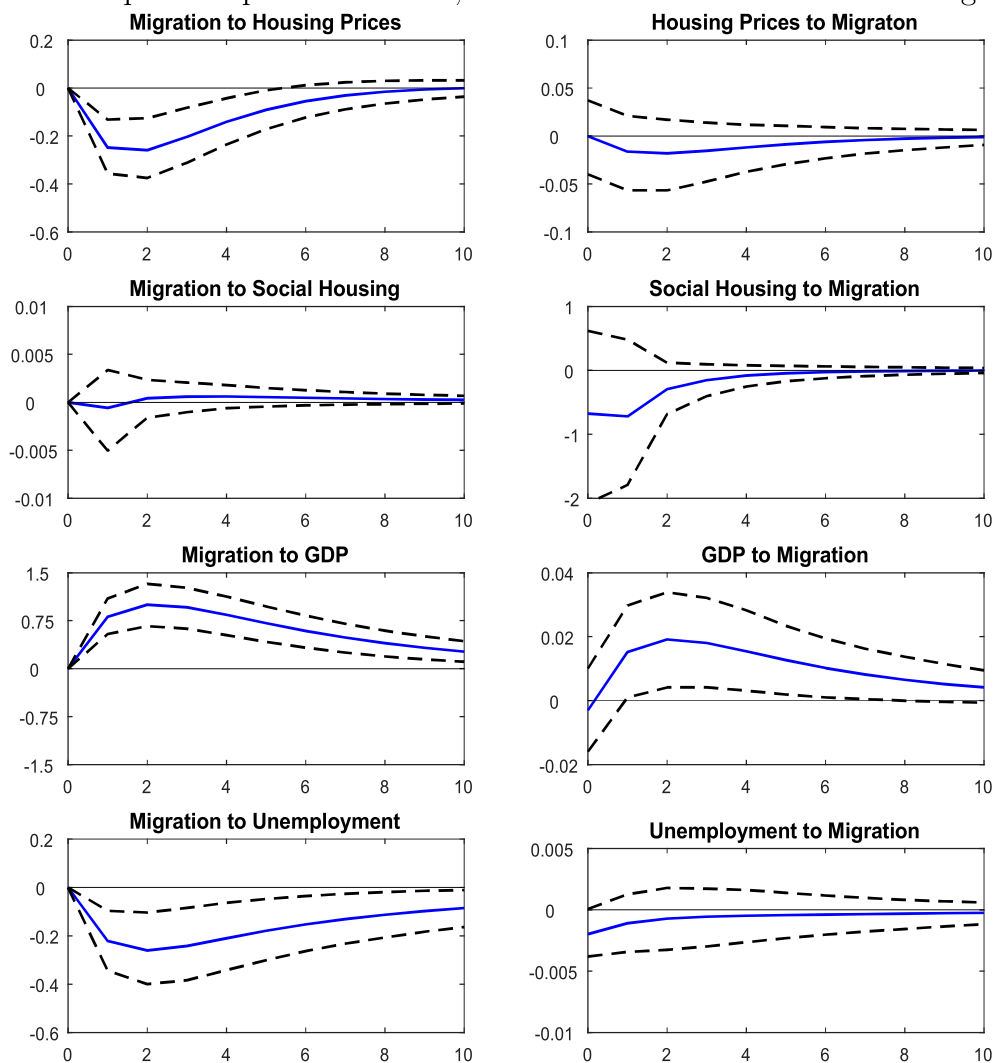
Notes: The solid line shows the estimated percentage change in the response variable due to a shock representing one percentage increase in the impulse variable. The dashed lines indicate the 90% confidence intervals that are generated by Monte Carlo with 5,000 repetitions.

Figure A6: Impulse response functions - Model 2, without Paris region



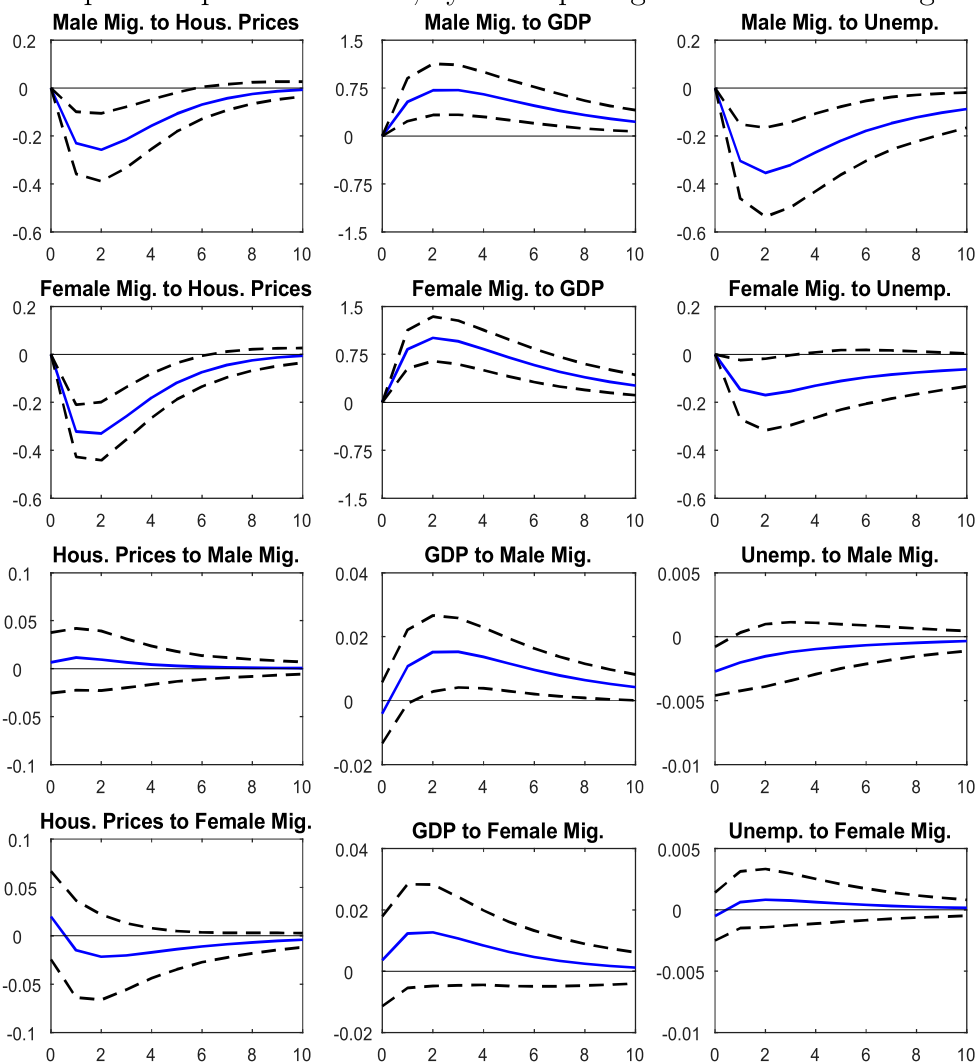
Notes: The solid line shows the estimated percentage change in the response variable due to a shock representing one percentage increase in the impulse variable. The dashed lines indicate the 90% confidence intervals that are generated by Monte Carlo with 5,000 repetitions.

Figure A7: Impulse response functions, alternative measure of social housing supply



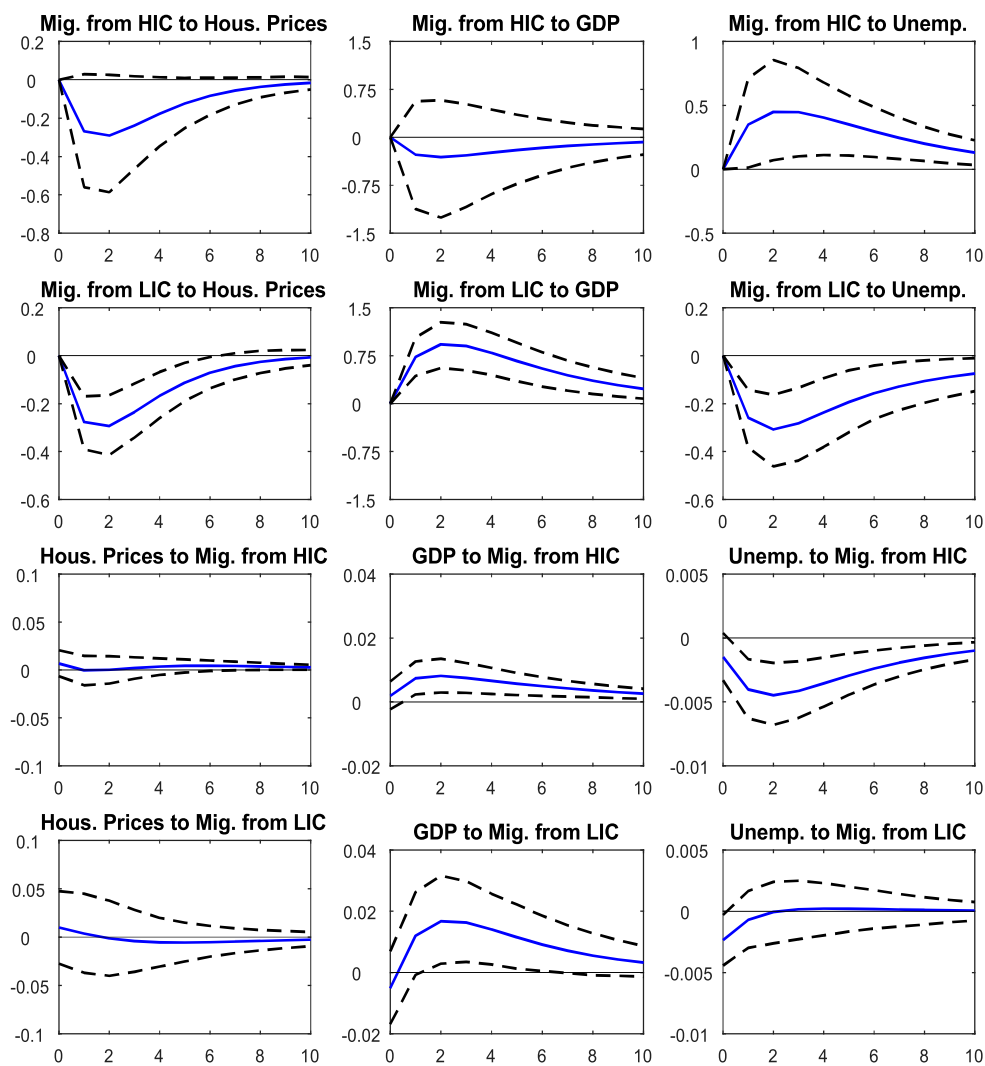
Notes: The solid line shows the estimated percentage change in the response variable due to a shock representing one percentage increase in the impulse variable. The dashed lines indicate the 90% confidence intervals that are generated by Monte Carlo with 5,000 repetitions.

Figure A8: Impulse response functions, by decomposing male and female migration rates



Notes: The solid line shows the estimated percentage change in the response variable due to a shock representing one percentage increase in the impulse variable. The dashed lines indicate the 90% confidence intervals that are generated by Monte Carlo with 5,000 repetitions.

Figure A9: Impulse response functions, by decomposing with the immigrants' country of origin



Notes: *LIC* and *HIC* stand for low-income countries and high-income countries, respectively. The solid line shows the estimated percentage change in the response variable due to a shock representing one percentage increase in the impulse variable. The dashed lines indicate the 90% confidence intervals that are generated by Monte Carlo with 5,000 repetitions.

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