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DECOMPOSING THE IMPACT OF IMMIGRATION ON HOUSE PRICES*

Rosa Sanchis-Guarner

ABSTRACT: Immigrant inflows affect local house prices by increasing housing demand when housing supply is fixed. In this paper, I show that we can formally decompose total demand changes into changes stemming from an immediate increase in population due to new arrivals ("partial effect") and additional changes in demand from relocated natives ("induced effect"). I propose a methodology to separately estimate these two effects using Spanish provinces' data from 2001- 2012. Applying an instrumental variables approach, I find that a 1 p.p. increase in the immigration rate increases average house prices by 3.3% and rents by 1%. Partial demand estimates are 24% smaller than the total estimates, due to immigrants and natives locating in the same provinces. The results show that accounting for the impact of immigration on native location choices is key to understanding net demand adjustments, as partial and total effects can significantly differ depending on native population mobility.

JEL Codes: J61, R12, R21
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1 Introduction

The study of the economic impact of immigration in the receiving regions has been a very active area of research in the last 30 years and continues to attract much attention from academics and policymakers. Recent large population displacements have renewed interest on the effects of large immigration inflows on the receiving regions. [1] Most of the theoretical and empirical contributions on this topic have originated from the analysis of their impact on labour markets (see Edo, 2019; Dustmann et al., 2016; Borjas, 2014, for reviews), but a large number of studies have investigated the effects on other outcomes such as crime (Bell et al., 2013; Ajzenman et al., 2022), education and occupational choices (Llull, 2018), political and social outcomes (Alesina & Tabellini, 2022) or welfare (Piyapromdee, 2021). When housing supply is fixed, immigrant inflows affect local prices by increasing housing demand. In this paper, I provide novel evidence on the impact of immigration on house prices, proposing a method to tease apart the effect due to demand stemming from new arrivals ("partial effect") and additional changes in demand from relocated natives ("induced effect"). The 2000s Spanish immigration wave is an ideal empirical set-up to study this as the country experienced sizeable increases in foreign-born population between 2000-2005 (González & Ortega, 2013), coupled with an unprecedented housing boom (Blanco et al., 2021).

A small number of papers have provided evidence on the impact of immigration on (consumption) goods prices (Lach, 2007; Cortés, 2008; Balkan et al., 2016), mostly finding negative effects. Zachariadis (2012) sums up the three theoretical effects of immigration inflows on prices: increase in aggregated demand (positive), increase in the search of low-priced goods (negative) and increase in supply of cheap labour for the production of labour-intense goods and services (negative). In the case of housing, which is by nature an inelastic good, the adjustment of its costs to an increase of local population might be different. [2] For a given housing and local population stocks, an inflow of foreign-born population intensifies spatial competition on the consumption of housing, which may initially push prices up. On the other hand, local population shocks might trigger internal migration across locations. Ultimately, the sign and magnitude of the total effect is ambiguous. The net impact is the result of three adjustments (Saiz, 2007): (1) increased demand from newly arrived immigrants ("direct" demand), (2) additional demand changes from relocated population ("induced" demand) and (3) changes in housing conditions (supply and occupancy density).

Most papers providing empirical evidence on the impact of immigration on house prices have estimated the net impact, paying less attention to the adjustments behind it. Many studies have found positive estimates of immigration on both local rents and house prices (Larkin et al., 2019).[3] In con-

[1] For example, very recent papers have looked at the receiving regions impact of large refugee waves. See Tumen (2016); Balkan et al. (2016) (from Syria to Turkey), Caruso et al. (2021) (from Venezuela to Colombia) and Trojanek & Gluszak (2022) (from Ukraine to Poland).

[2] See Cochrane & Poot (2021) for a review of this literature.

[3] Saiz (2003, 2007); Ottaviano & Peri (2012a); Sharpe (2019) provide estimates for the USA. Other studies include Hyslop

trast, a handful of papers (Hatton & Tani, 2005; Saiz & Wachter, 2011; Accetturo et al., 2014; Sá, 2015) have estimated negative impacts of immigration on average house prices, in particular when focusing on small neighbourhoods. More recently, Monras (2020) finds that in the long-run, an increase of Mexican-induced low-skilled workers immigration reduced housing and rental prices, and argues this is due to a large increase in the supply of construction workers, which pushed down construction costs. The displacement of natives from these areas is the main argument used to explain these negative findings but, except in Sá (2015), this channel is rarely explicitly estimated. In the case of Spain, a few papers have studied the role of immigration on housing markets using different levels of spatial aggregation with mixed findings (García-Montalvo, 2010; Nicodemo & Raya, 2012). Notably, González & Ortega (2013) attribute half of house price growth to immigration inflows in Spain during the period 1998–2008. Using Lewis & Peri (2015) terminology[4], my paper aims to provide a framework to interpret the coefficients according to their partial or total effect on prices, enabling a better understanding of the reduced form estimates. The two effects add up to the total demand effect.

I start by showing that, under certain assumptions, we can show that the total impact of immigration on house prices is the net sum of two effects: (1) the direct increase in local population from the new arrivals and (2) the additional changes from relocated population ("displacement"), both affecting housing demand. These adjustments relate to demand as long as supply is accounted for in the estimation equations. The first component is theoretically positive and it resembles a reduced-form demand coefficient. The second component can be positive or negative, depending on the impact that immigrants have on native mobility. The sum of both components determines the size and sign of the total "net demand" effect.

To provide the empirical counterparts of the decomposition, I proceed in three steps. In Step 1, I estimate a similar specification to Saiz (2007), regressing the annual local house price growth (sale and rental) on the immigration rates. I use data for Spanish provinces between 2001 and 2012. To be able to make causal claims about the estimates, I use a modified version of the standard immigration shift-share instrument and control for relevant local characteristics, area and province fixed effects. The baseline estimated semi-elasticities are approximately 1% for rents and 3.3% for house prices, for an increase in 1 percentage point (p.p.) of the immigration ratio.[5] These estimates correspond to the total

et al. (2019) who provide positive impact estimates for New Zealand; Degen & Fischer (2009) who find positive effects for Switzerland; Akbari & Aydede (2012) find positive but small impacts in Canada; Frostad (2014) estimates the impact in Norway; and Kürschner (2017) for Germany. Tumen (2016) study the impact of the Syrian refugees inflow on Turkish housing rents and finds a positive effect on high quality units and no effect on low-quality units.

[4] Lewis & Peri (2015) pg 4–5: "Traditionally the economic analysis has distinguished between short and long run effects of immigration. However, the so-called short-run effects are mostly a theoretical device to decompose a complex effect. When economists analyze the 'short-run effects' of immigrants they try to isolate the consequences of immigration when all other variables (including the stock of capital, the skill supply of natives and the technology and productive structure) are fixed. This should be called 'partial' effect. It is a way to understand and isolate a specific effect, not a way to forecast what happens, even in the short run.". In fact, Saiz (2007) refers to long and short-run impacts when allows for adjustments on native population and housing conditions (total) or does not (partial).

[5] My coefficients are directly comparable to Saiz (2007) estimates of around 0.9% for rents and 3.3% for prices, and

effect. In Step 2, I explicitly test the impact of immigrant inflows on native mobility and, in line with existing estimates for Spain (Fernández-Huertas et al., 2009), I find that immigrants attract natives to areas in which they locate (approximately 3 natives for each 10 immigrants). [6] To identify the impact of immigration on prices that is only due to increased immigrant housing demand ("partial" effect), in Step 3 I use population growth rate as the main regressor. I estimate the coefficient of this variable using solely the variation on population growth which is due to exogenous location of immigrants (predicted by the instrument). The estimated immigrant demand semi-elasticities using this methodology are 0.75% for rents and around 2.5% for house prices. The difference between the empirical estimates of steps 1 and 3 corresponds to the change in demand from natives locating in the region contemporaneously to the immigrant inflow, i.e. induced native demand.

The total and partial estimated coefficients correspond to demand effects when the role of housing supply is partialled out in the regressions. In the estimation of the baseline results, which already include province fixed effects, I also include local attributes in trends that relate to housing supply conditions. I further explore further the impact that directly controlling for housing stock changes has on the estimates, using an instrument for changes in housing stock. I find that they have very little additional effect on the coefficients. Multiple tests on the validity of the instrumental variable strategy are provided in section 5. The empirical results are robust across different specifications, to different constructions of the instrument and remain very similar when using long differences instead of year-to-year variations. In an online appendix (section B.3), I propose a simple model to explain the mechanisms behind the results: an inflow of immigrants increases housing costs in the receiving region but, due to the specialisation of natives and immigrants on different production sectors, it also attracts natives to the region, increasing house costs even further. The model predicts that the total demand effect would be larger than the immigrant demand effect, which is what I find in the empirical exercise.

The results of the paper highlight the importance of taking into account local population mobility when interpreting the effect of immigration on house prices, or any other local outcome affected by population changes. The impact of population mobility on the identification of aggregate local effects gained renewed interest after the publication of Borjas (2003). This paper criticised studies on regional labour market impacts of foreign-born inflows, claiming that the United States worked as a single labour market and that the existence of mobility across areas could hinder the estimation of regional effects. The lack of local effects could be the result of exit of native population after an inflow of immigrants resulting in a net zero or very small change in local labour demand. As total housing demand changes are the result of new and induced population changes, if these have oppos-

similar to other existing estimates (Larkin et al., 2019).

[6] This sizeable co-location estimate is rare, but not unique in the literature. In fact, even in the US the evidence on significant native displacement is not as robust as some authors have claimed as discussed in (Peri & Sparber, 2011). Mocetti & Porello (2010) and Wozniak & Murray (2012) find similar size estimates of co-location of natives and immigrants. In Spain, Fernández-Huertas et al. (2019) find co-location of natives and immigrants in newly developed neighbourhoods.

ing signs, the net estimates might be close to zero but masking sizeable partial adjustments. Previous papers have relied on the existing US evidence to argue that native area displacement due to immigration is small or not large enough to cancel out increased demand stemming from increased area population so its impact on the estimates is irrelevant, and thus discussed total and partial effects as equivalent.[7] However, my findings suggest that the impact of immigrants on native location can be non-negligible, so we need to be more careful about making these claims.

The current paper makes several contributions. First, from combining the estimating equations I show that the coefficient that captures total demand changes can be formally decomposed as the sum of direct (immigrant) demand changes plus additional demand shifts from relocated population (induced). This is the first paper to provide causal estimates of all the elements of this decomposition, and in particular, to identify the relationship between them. My paper provides a framework to better understand the demand adjustments on local house prices following a large immigration inflow. Second, to obtain causal counterparts of the decomposition elements, I exploit Spanish province level data between 2001 and 2012. The magnitude and timing of the immigration inflows during this period provides an appropriate setting with "quasi"-experimental variation (Fernández-Huertas et al., 2019). I construct a shift-share instrument that combines historical immigrant location patterns with predicted national inflows by country of origin obtained from a push-factors gravity model. This is an improved version of the standard ethnic networks instrument widely used in the immigration literature and it is the first time it is applied to Spanish data. The this modification of the traditional shift-share immigration predictor addresses many of the concerns raised in the recent shift-share instruments literature.[8] Third, in order to be able to interpret the results in terms of changes of prices and quantities in equilibrium in my estimations I condition out housing supply conditions. The main results are estimated including province fixed effects and additional province-level supply-related attributes interacted with year dummies. By doing this, my estimates correspond to demand impacts, which makes their interpretation more straightforward. Finally, I estimate the effects of immigrants not only on house prices, but also on rents. Even if renting is the primary housing tenure of immigrants, immigrants also purchase homes and indirectly affect prices when housing is bought as an investment good (González & Ortega, 2013). The evidence on the impact on rents complement the estimates on sale prices and provide a more complete picture of the effects on housing costs.

The rest of the paper is organised as follows. Section 2 provides the formal decomposition of the total impact. Section 3 describes the methodology:the empirical specifications (3.1), how the components of the decomposition can be estimated (3.2), the instrumental variables (IV) strategy (3.3) and the data (3.4). Section 4 presents the main regression results. Section 5 discusses the validity of the IV strategy and provides results robustness checks. Finally, section 6 presents the conclusions.

[7]This is the case for example in González & Ortega (2013).

[8]Notably Jaeger et al. (2018), Adao et al. (2019), Goldsmith-Pinkham et al. (2020) and Borusyak et al. (2022).

2 Decomposition of the effect of immigration on house prices

2.1 Total and partial effects

Let's assume a simple supply-demand framework where the observed local average house price is determined by the demand from local population and by the level of housing stock. When supply is fixed, changes in local average prices would occur when there are changes in the level of population in the province (by shifting housing demand). I focus on large local inflows of foreign-born population as the main driver of local population changes. Tables 1 and B.3 illustrate the magnitude of population changes in Spain during these years.

In my context, I do this for two reasons. First, during the period of analysis (2001–2012) most of the population changes in Spain stemmed from foreign-born population inflows (around 76%), especially before 2009. While the national annual growth rate of foreign-born population was on average over 13% (with peaks around 30% between 2001 and 2003), annual growth rates of native population remained stable and very low throughout (around 0.25%). Thus, aver the period, the average population growth rate amounted to around 1% per year. Second, for local changes in foreign-born population it is possible to construct a plausible source of exogenous variation, which is key in the empirical exercise and for the results in this section. One might want to assess the impact of changes in total local population (demand) on housing costs. However, as discussed by Sharpe (2019), finding a credible instrument to estimate the impact of *total population changes* on local prices is very difficult, and using two instruments for immigrants and natives separately is also problematic (Angrist & Pischke, 2009). What can be plausibly done is to capture changes in local population driven by immigration, isolating the impact of immigration on outcomes via their effect on the total number of local inhabitants. This can be achieved using the variation on total population driven by the exogenous source of immigration inflows.

The total effect of changes in foreign-born population on house prices has been previously estimated using variations of the Saiz (2007) empirical specification:

$$\Delta \log (hpr_{i,t}) = \beta_1 \left(\frac{\Delta FB_{i,t-1}}{POP_{i,t-2}} \right) + \theta^1_{i,t} + \epsilon^1_{i,t} \tag{1}$$

In this equation, β_1 captures the total impact of immigration on prices: the net total adjustments due to the immigration inflow plus additional changes in native location, where changes in housing supply conditions are be accounted for by θ^1), while ϵ^1 is a well-behaved error term. $\left(\frac{\Delta FB_{i,t-1}}{POP_{i,t-2}} \right)$ is the normalised immigration inflow during $t-1$ (or immigration rate) and it is measured in percentage points. β_1 is interpreted as the percent change in house prices for an increase in the rate of one percentage point.

Saiz (2007) discusses three channels affecting the "long-run" (total) adjustment of local house prices. Initially, an increase in foreign-born population in a given location raises total population and then pushes demand and prices in the "short run" (partial direct demand effect). The total net impact ("long-run") on prices also depends on changes on housing consumption (density), on housing supply (construction) and on the mobility of natives or previous residents (displacement) following the immigration inflow. [9] Unless we partial-out these adjustments, an estimate of β_1 would capture the combination of all these changes. Conditional on housing supply, an unbiased estimate of β_1 would be the net result of changes in housing demand which stem from the newly arrived population plus any changes in demand related to natives relocating *due to* the immigration wave. Saiz (2007) argued that the impact of immigrant demand (partial effect) cannot be separated from other demand changes induced by native population adjustments , affecting the interpretation of the estimated semi-elasticities.[10] The use of instruments and controls would produce unbiased estimates of the total effect coefficient, but it does not help with interpreting the channels driving it. Hence, the main contribution of my paper is to show that it is indeed possible to estimate the direct and induced changes in housing demand, while conditioning on housing supply conditions.

When estimating local average impacts of immigration inflows, one needs to take into account that changes in population in a given area affect the whole regions-cities system equilibrium. The relocation of population across regions within a country could hinder the identification of any area-level effects, as the effects of a local immigration inflow would dissipate throughout the country if natives relocate. If large population outflows are triggered by immigration, the net area impact would tend towards zero. In the analysis of the impact of immigration on local labour market outcomes, the existence of "native displacement" has been used as an explanation for the lack of robust estimates of the impact of immigration on wages across US labour markets (Borjas, 2006). When displacement exists, cross-region regressions would underestimate the effect of immigrants on local labour markets. However, native displacement might be quantitatively small, as discussed in Peri & Sparber (2011). Not only for regional wages, native mobility also affects the estimation of the average effect of immigration on local house prices. The impact of mobility on the interpretation of the net area estimates has been generally inferred from the sign of the total impact estimate. If the reduced-form estimate is positive it is assumed that natives are little or not "sufficiently" displaced by immigrants (González & Ortega, 2013). Thus the sign of the total effect is interpreted as a test of "native displacement" (Saiz,

[9] The theoretical discussion in Saiz (2007) uses short-run and long-run adjustments terminology. By short-run he refers to a situation where native mobility and housing supply cannot adjust and by long-run when they do. In the empirical results of his paper he uses both annual first-differences models and decennial long-differences models, but the interpretation of the coefficient is always the same – the total (reduced-form) effect. In fact, as discussed in Lewis & Peri (2015), when economists refer to the "short-run" effects of immigrants, they try to isolate the consequences of immigration on one variable keeping the rest fixed. These authors refer to this as a "partial" effect. Thus, in this paper I refer to "total" effect when I allow for adjustments in the three channels and to "partial" effect as the direct increase in demand from recently arrived population.

[10] *"There is no way to separate the effect of increased housing demand (immigration) from the potential decreased demand associated with potential native out-migration. Part of the local response to the treatment (immigration) can occur through native out-migration. In this case, we need to be careful about the interpretation of the coefficient of immigration on rents."* (Saiz, 2007, pg. 348).

2007). However, the impact of immigrants on native mobility is often not directly estimated in these studies.[11] The specific relationship between total and partial effects is derived and discussed in the next section.

2.2 Formal decomposition of the total effect

The first step to derive the channels behind the total impact of immigration on prices is to define the relationship between price growth and population growth:

$$\Delta \log (hpr_{i,t}) = \beta_2 \left(\frac{\Delta POP_{i,t-1}}{POP_{i,t-2}} \right) + \theta^2_{i,t} + \epsilon^2_{i,t} \tag{2}$$

where $\left(\frac{\Delta POP_{i,t-1}}{POP_{i,t-2}} \right)$ is the local population growth rate. $\theta^2_{i,t}$ includes variables that relate to housing supply conditions, as well as time and area fixed effects. In the empirical exercise, I use plausibly exogenous variation in local foreign-born stocks to estimate β_2. Hence, this coefficient captures the effect of immigration in prices through its impact on local population. Importantly, it is not the effect of changes of *total population* on prices because, as explained above, it is very difficult to find credible variation to estimate this. By including supply controls in the specification, it captures the impact via changes in housing demand.

Changes in local population are by definition the sum of changes in foreign-born and natives:

$$\Delta POP_{i,t-1} \equiv \Delta FB_{i,t-1} + \Delta NAT_{i,t-1} \tag{3}$$

Replacing (3) into (2):

$$\Delta \log (hpr_{i,t}) \quad = \quad \beta_2 \left(\frac{\Delta FB_{i,t-1}}{POP_{i,t-2}} \right) + \beta_2 \left(\frac{\Delta NAT_{i,t-1}}{POP_{i,t-2}} \right) + \theta^2_{i,t} + \epsilon^2_{i,t} \tag{4}$$

This expression decomposes local population growth into the two components that can be affected by immigration: immigrant $\left(\frac{\Delta FB_{i,t-1}}{POP_{i,t-2}} \right)$ and native population $\left(\frac{\Delta NAT_{i,t-1}}{POP_{i,t-2}} \right)$ rates.

The following equation defines how native rates are affected by immigrants:

$$\left(\frac{\Delta NAT_{i,t-1}}{POP_{i,t-2}} \right) = \beta_3 \left(\frac{\Delta FB_{i,t-1}}{POP_{i,t-2}} \right) + \theta^3_{i,t} + \epsilon^3_{i,t} \tag{5}$$

where β_3 captures the reaction of native mobility to an immigration inflow and its sign and size would inform about the existence of displacement or co-location. Plugging (5) into (4) and rearranging:

$$\Delta \log (hpr_{i,t}) \quad = \quad (\beta_2 + \beta_2\beta_3) \left(\frac{\Delta FB_{i,t-1}}{POP_{i,t-2}} \right) + \beta_2\theta^3_{i,t} + \theta^2_{i,t} + v_{i,t} \tag{6}$$

[11] With the exception of Sá (2015).

This expression shows that, when we use the components of local population, we can decompose the impact of immigration on prices in terms which depend on parameters β_2 and β_3, and other terms such as $\theta^2_{i,t}$, $\theta^3_{i,t}$ and $v_{i,t}$, where $v_{i,t} = \left(\beta_2 \epsilon^3_{i,t} + \epsilon^2_{i,t}\right)$. Coefficient β_2 is the effect of immigration in prices through its impact on local population (partial or *direct* demand impact) and β_3 is the impact of immigration rates on native population rates (*mobility*). The interaction of both parameters, $\beta_2\beta_3$, captures the changes in local demand due to natives relocating following the immigration wave (*induced* demand).

In order to obtain an expression that relates the total impact to its partial adjustment components, I turn to the empirical counterparts of equations (1) and (6). I start by differentiating equation (1) with respect to changes in the immigration rate:

$$\frac{\partial \Delta \log(hpr_{i,t})}{\partial\left(\frac{\Delta FB_{i,t-1}}{POP_{i,t-2}}\right)} = \beta_1 + \frac{\partial \theta^1_{i,t}}{\partial\left(\frac{\Delta FB_{i,t-1}}{POP_{i,t-2}}\right)} + \frac{\partial \epsilon^1_{i,t}}{\partial\left(\frac{\Delta FB_{i,t-1}}{POP_{i,t-2}}\right)} \tag{7}$$

Then, I also differentiate equation (6) with respect to changes in the immigration rates:

$$\frac{\partial \Delta \log(hpr_{i,t})}{\partial\left(\frac{\Delta FB_{i,t-1}}{POP_{i,t-2}}\right)} = \beta_2 + \beta_2\beta_3 + \beta_2 \frac{\partial \theta^3_{i,t}}{\partial\left(\frac{\Delta FB_{i,t-1}}{POP_{i,t-2}}\right)} + \frac{\partial \theta^2_{i,t}}{\partial\left(\frac{\Delta FB_{i,t-1}}{POP_{i,t-2}}\right)} + \frac{\partial v_{i,t}}{\partial\left(\frac{\Delta FB_{i,t-1}}{POP_{i,t-2}}\right)} \tag{8}$$

By definition, $\epsilon^1_{i,t}$, $\epsilon^2_{i,t}$ and $\epsilon^3_{i,t}$ (in $v_{i,t}$) are uncorrelated with $N\Delta FB_{i,t-1}$, so the last terms in equations (7) and (8) would be zero. The partial derivatives of the θ terms with respect to immigration rates would capture any biases in the estimation of consistent β_1, β_2 and β_3. $\frac{\partial \theta^1_{i,t}}{\partial\left(\frac{\Delta FB_{i,t-1}}{POP_{i,t-2}}\right)}$ captures the correlation between variables related to immigration rates and house prices in equation (1). The same applies to $\frac{\partial \theta^2_{i,t}}{\partial\left(\frac{\Delta FB_{i,t-1}}{POP_{i,t-2}}\right)}$: it captures the impact of variables correlated with changes in house prices and changes in population (from immigration inflows) in equation (2). Finally, $\frac{\partial \theta^3_{i,t}}{\partial\left(\frac{\Delta FB_{i,t-1}}{POP_{i,t-2}}\right)}$ captures the correlation between variables related to native and immigration rates at the same time in equation (5). These variables can include observables (which can be accounted for) and un-observables. I deal with omitted variable bias and fixed unobservables by using a large set of controls and time and area fixed effects. To eliminate endogeneity bias from un-observables I use an IV strategy, discussed in detail in section 3.3.[12] Therefore, in the empirical counterparts of these terms, the θ terms would not directly depend on the immigration rate and would be zero.

When the coefficients are consistently estimated we can match (7) and (8) to decompose the total impact $\hat{\beta}_1$ as follows:

$$\hat{\beta}_1 = \hat{\beta}_2 + \hat{\beta}_2\hat{\beta}_3 = \hat{\beta}_2\left(1 + \hat{\beta}_3\right) \tag{9}$$

where $\hat{\beta}_2$ is the impact of immigration on prices via its impact on the size local population changes

[12] When the coefficients are estimated by OLS, expression (7)=(8) would still empirically hold. What the IV strategy allows is to obtain a simplified expression where we do not need to evaluate all the cross-derivatives of the θ terms.

(direct "demand" impact) and $\hat{\beta}_3$ is the impact of immigration on local native population changes (native mobility or the so-called "native-displacement"). The term $\hat{\beta}_2\hat{\beta}_3$ captures the changes in prices that are due to additional changes in demand from relocated (native) population (induced demand impact). This term can be positive or negative depending on how native mobility is affected by immigrants. $\hat{\beta}_1$ is the total impact which captures the net sum of all these changes. These coefficients correspond to demand effects as long as supply conditions are partialled-out in the estimation.

Coefficient $\hat{\beta}_2$ can be interpreted as the semi-elasticity of prices with respect to changes in local population, which in our setting is estimated from an exogenous change in the foreign-born population (predicted by the instrument) and it is expected to be positive (for an inelastic normal good). It is the partial impact, when we do not consider additional adjustments in demand and we control for adjustments in supply. The sign and size of the net total impact $\hat{\beta}_1$ depends on the term $(1+\hat{\beta}_3)$, which captures the impact of native mobility on additional changes in local demand. This last term could be negative –if immigrants displace natives more than one-to-one–, positive but smaller than one –if immigrants displace natives but not one-to-one–, one –if immigrants have no impact on native mobility–, or greater than one –if immigrants and native co-locate.

If we ignored the impact that native mobility has on changing the total demand in the region via its impact on natives mobility, we could assume that $\hat{\beta}_1$ (total) and $\hat{\beta}_2$ (partial) are the same and that $\hat{\beta}_1$ corresponds to the changes in housing demand from immigrants. This would only be the case if natives are unaffected by immigrant inflows, e.g. $\hat{\beta}_3 = 0$. In reality, only the partial effect coefficient corresponds direct demand changes, while the total (demand) impact includes changes in demand from the exogenous change in population (from the immigrant inflow) and from endogenous change in (native) population. The decomposition provided in this paper clarifies this issue and the results in this paper provide the first joint estimates of all three impacts in a relevant context. The next sections explain the methodology used to obtained consistent coefficients of the three effects and shows that the decomposition in equation (9) holds exactly in the data.

3 Methodology

3.1 Empirical specifications

To obtain the estimates of equation (9) we estimate the following models:

$$\Delta \log(hpr_{i,t}) = \beta_1 \left(\frac{\Delta FB_{i,t-1}}{POP_{i,t-2}}\right) + \lambda_t + \gamma_i + \phi'(Z_i * \lambda_t) + \delta'\Delta X_{i,t-2} + \varepsilon^1_{i,t} \tag{10}$$

$$\Delta \log(hpr_{i,t}) = \beta_2 \left(\frac{\Delta POP_{i,t-1}}{POP_{i,t-2}}\right) + \lambda_t + \gamma_i + \phi'(Z_i * \lambda_t) + \delta'\Delta X_{i,t-2} + \varepsilon^2_{i,t} \tag{11}$$

$$\left(\frac{\Delta NAT_{i,t}}{POP_{i,t-1}}\right) = \beta_3 \left(\frac{\Delta FB_{i,t}}{POP_{i,t-1}}\right) + \lambda_t + \gamma_i + \phi'(Z_i * \lambda_t) + \delta'\Delta X_{i,t-1} + \varepsilon^3_{i,t} \tag{12}$$

Equation (10) estimates the total impact of immigrants on house prices, equation (11) the partial immigrant demand (direct) and (12) the impact of immigration on native mobility. The geographical units of observation are the 50 Spanish provinces i (NUTS3) and t denotes time periods (years 2002 to 2012). [13] $\Delta \log(hpr_{i,t/t-1})$ is the change of the natural logarithm of average house prices (sale or rental prices) in province i during year t (approximately the growth rate), while $\left(\frac{\Delta FB_{i,t-1}}{POP_{i,t-2}}\right)$ is the immigration rate during $t-1$ (and similarly for natives (NAT) and total population (POP)). The inflow of immigrants during $t-1$ is calculated as the change in the foreign-born population between the end of $t-1$ and the end of $t-2$. The same applies when I use native or total population changes.[14] The identification strategy aims to produce causal consistent estimates of β_1, β_2 and β_3. I use these coefficients to check if the decomposition holds in the data. Similarly to other authors, in the main results I use the immigration inflows lagged one period with respect to changes in prices. [15] Hence, inflows in the sample correspond to years 2001 to 2011, while price growth rates to years 2002 to 2012. λ_t are time fixed effects, γ_i are province fixed effects, $\Delta X_{i,t-2}$ is a matrix of province time-varying controls and Z_i is a matrix of province time-invariant attributes. In the most demanding estimations, I include province attributes interacted with time dummies ($Z_i * \lambda_t$). Finally, $\varepsilon_{i,t}$ is the random error term. In this set-up the coefficients of interest are in equations (10) and (11) are interpreted as a semi-elasticity: an increase in the rate of one percentage point has an effect on the change in prices of β percentage points. In equation (12), β_3 captures the number of natives that relocate for each additional immigrant (Peri & Sparber, 2011).

The first thing to notice is that we are estimating the relationship between immigration and house prices in growth rates and population changes. The first differences setting of equations (10)–(12) already eliminates any unobservable province characteristics which might be correlated with the outcomes and the population rates *in levels*. However, there could still exist some confounders at the area level correlated province outcomes and treatments *in changes*. These could be of two types: unobservable and observable province characteristics. Vector Z_i contains time-invariant province attributes that control for the fact that provinces with different levels of these characteristics might have different growth trends. This includes variables related to the nature of housing supply (share of residential secondary homes, share of residential empty homes, share of households which own a secondary home) and housing consumption (share of renters and log average home square me-

[13] I exclude the African territories due to their historical particularities and the lack of reliable data.

[14] Using normalised inflows instead of (log) net inflows as the measure of "immigration" eliminates any unobservables that might equally affect both the numerator (immigration inflow) and the denominator (original province population). Standardising immigration inflows by initial population stock also deals with the fact that regions of different sizes have different population and house price dynamics (Card, 2001; Peri & Sparber, 2011; Wozniak & Murray, 2012). Scale effects can induce spurious correlation between higher inflows and higher price changes. This correlation could arise due to the fact that the average and standard deviation of both variables are likely to be proportional to the total population in the province. In addition to the standarisation, I control for the effect of initial population trends by either including it directly in the specifications or by using province fixed effects. Additionally, this format allows a more straightforward comparison to existing estimates in papers that have used similar specifications.

[15] I also investigate other lags as a robustness test in table 6.

ters per person; for foreign-born and for natives separately) in 2001. I also include information about the economic structure in 2001 (share of employed in construction sector and share of employed in services sector) and some other attributes potentially correlated with location decisions and price growth (log road distance to Madrid, length of coastline in 100s of kilometres, log of rain precipitation (January). Finally the share of developable land and average ruggedness index in the province aim to capture the factors related to the potential growth of housing supply. When I use province fixed effects the province attributes included in Z_i drop. In the most demanding specification, I include province fixed effects and province attributes interacted with time dummies, which control for differential growth trends by level of attribute.

Vector $\Delta X_{i,t-2/t-3}$ contains time-varying province characteristics (in changes). I control for changes in output per capita, conditions in the financial sector, unemployment rate, education levels and infrastructure endowments. As contemporaneous changes of these factors could well be the result of population changes ("bad controls"), I use an additional lag with respect to population rates variable, e.g. changes in the variables during $t-2/t-3$ (one period before the inflows in $(t-1)$ and two periods before the change in prices $(t/t-1)$). The results change very little if the time-varying controls are contemporaneous to the population rates or excluded.

Other factors potentially inducing bias in the estimates could be unobservables. Time fixed effects λ_t control for common unobserved shocks affecting all Spanish provinces in a given year. Province fixed effects γ_i control for time-invariant province heterogeneity. When including both sets of fixed effects, the specification corresponds to a first differences fixed effects estimation (e.g. growth regression with fixed effects). In this model, the coefficients are estimated off the within-province time variation in outcome and treatment changes, conditional on the controls and time fixed effects. To deal with time-varying unobservables I use an IV strategy which is described in section 3.3.

3.2 Estimation of the components of the decomposition

In this section I describe the empirical issues related to the estimation of the components of the decomposition and their interpretation as total or partial demand estimates. This relates to the consistent estimation of the coefficients, by using plausible exogenous variation, and controlling for supply-related factors in the estimation. Firstly, for equation (9) to hold, the three β coefficients must be consistently estimated. The IV estimation of the β_1 coefficient in (10) yields a consistent estimate of the total impact of immigration on prices. To obtain the right-hand-side components on (9) I also need to estimate coefficients β_2 (partial demand impact) and β_3 (native mobility).

As suggested by Peri & Sparber (2011), to test the impact of immigration on native mobility I use a normalised change in native population in the left-hand-side and estimate equation (12). The sign and size of $\hat{\beta}_3$ captures the relationship between immigration inflows and native location and measures how many natives relocate in response to one immigrant arrival. If a sizeable causal relationship exists, we need to be more cautious about the interpretation of $\hat{\beta}_1$. The results of the estimation of the impact of immigration on native mobility are discussed in section 4.2.

In the theoretical framework of section 2 I refer to the interpretation of $\hat{\beta}_1$ and $\hat{\beta}_2$ as total and partial *demand* coefficients. This interpretation relies on conditioning out housing supply conditions (housing density and homes supply) when estimating these parameters. Changes in housing density might have evolved very different for provinces that received more or less immigrants during this period. To investigate this we define provinces with *high* and *low* immigration status during the 2001-2011 period, where we classify the province over and below the median immigration rate between 2001 and 2011. Panel (a) of Figure A.1 shows that, while it grew for both, these two groups differed greatly on the evolution and the level of the share of foreign-born population over time.

Table 1 on the other hand displays the total population, total and per person housing stock in years 2001 to 2012, for all provinces and by high and low immigration status. Both variables increased dramatically during the period, but housing density remained relatively stable, moderately very little over. This suggests that intensive construction together with large immigrant inflows kept the rate of houses/population relatively constant over the period. The table also shows the the average housing stock per person was relatively different between provinces with high and low immigration in 2001. However, panel (b) in figure A.1 shows that these differences are not statistically significant in any year during the period, and converge by 2012. Hence, I consider the province fixed effects to sufficiently control for the differences in housing consumption across provinces over time.

Table 1: Residential density in Spain 2001-2012

	ALL PROVINCES			LOW IMMIGRATION			HIGH IMMIGRATION		
Year	*Population*	*Housing stock*	*Average stock/pop*	*Population*	*Housing stock*	*Average stock/pop*	*Population*	*Housing stock*	*Average stock/pop*
2001	41,692,558	20,988,378	0.545	15,940,732	7,728,691	0.520	25,751,826	13,259,687	0.571
2002	42,573,670	21,440,413	0.549	16,044,778	7,877,993	0.527	26,528,892	13,562,420	0.570
2003	43,055,014	21,878,187	0.554	16,105,694	8,018,749	0.535	26,949,320	13,859,438	0.573
2004	43,967,766	22,368,785	0.555	16,241,244	8,177,697	0.541	27,726,522	14,191,088	0.570
2005	44,566,232	22,877,640	0.560	16,346,618	8,336,402	0.548	28,219,614	14,541,238	0.573
2006	45,054,694	23,443,569	0.568	16,419,649	8,519,092	0.558	28,635,045	14,924,477	0.579
2007	46,008,985	23,983,886	0.571	16,581,611	8,709,850	0.565	29,427,374	15,274,036	0.578
2008	46,593,673	24,518,341	0.580	16,688,538	8,898,996	0.574	29,905,135	15,619,345	0.585
2009	46,864,418	24,856,498	0.587	16,739,433	9,024,952	0.582	30,124,985	15,831,546	0.591
2010	47,029,641	25,054,029	0.591	16,771,796	9,113,294	0.587	30,257,845	15,940,735	0.594
2011	47,100,501	25,196,069	0.595	16,756,995	9,182,064	0.593	30,343,506	16,014,005	0.598
2012	46,961,924	25,328,848	0.603	16,678,085	9,234,733	0.601	30,283,839	16,094,115	0.605

Notes: Spanish Department of Housing and Annual Population Registers. High/low immigration status is defined by being above or below the median in the foreign-born inflow (2001-2011) over population in 2001 (0.0783). The low immigration provinces are (sorted from lower to higher): Caceres, Jaen, Cordoba, Cadiz, Orense, Badajoz, Zamora, Lugo, Palencia, La Coruña, Salamanca, Pontevedra, Leon, Asturias, Sevilla, Valladolid, Vizcaya, Albacete, Guipuzcoa, Granada, Cantabria, Avila,Ciudad Real, Burgos and Huelva. The high immigration provinces are (sorted from lower to higher): Alava, Soria, Navarra, Valencia, Teruel, Segovia, Huesca, Cuenca, La Rioja, Las Palmas, Zaragoza, Santa Cruz de Tenerife, Murcia, Barcelona, Toledo, Madrid, Castellon, Malaga, Baleares, Lleida, Tarragona, Girona, Alicante, Almeria, Guadalajara.

Table 1 also shows that a large number of housing units were built between 2001 and 2012 (over 5.2 million, more than 400,000 per year). House construction could be directly correlated with immigration inflows if immigrants locate in areas where house construction is higher. I account for the effect of housing supply on house prices in two ways. Firstly, in the estimation of (10)-(12) I include a set of time-invariant province attributes and trends related to housing supply (explained above). However, if we want to truly condition-out the impact of changes on housing supply on price growth, we need to include time-varying changes in housing supply as an additional control variable. This variable would remove potential bias arising from the fact that immigrants might be locating in areas where construction is growing faster (to work in this sector or due to higher availability of homes) and that house construction also affects housing costs via increasing supply of housing units. Including time-varying supply changes in the estimation as an additional control is very problematic, because even if lagged, housing construction is a "bad control" given that construction is directly affected by immigration.[16] In section A.1, I discuss the impact that directly including a measure of changes in housing stock as an additional control (also instrumented) has on the estimates. Given the demanding empirical specification and the set of fixed effects and controls, the impact of adding this additional variable on the main estimates is negligible. As including this additional control barely affect the main estimates, I refrain from doing this and rely on the specifications (10)-(12) as specified above.

Once we have considered the impact of native mobility and partialled-out supply conditions on the total estimates, we need to apply a method to pin down the partial demand coefficient β_2. To identify the variation in total population that arises only from immigrant inflows, I replace immigration rate in (10) by a normalised population inflow variable (or population growth rate) and estimate the effect of changes in population using solely the variation which is due to exogenous location of foreign-born (predicted by the instrument). In practice, this corresponds to the estimation of model (11) using an instrument for $\left(\frac{\Delta POP_{i,t-1}}{POP_{i,t-2}}\right)$. When I do this, $\hat{\beta}_2$ is estimated from exogenous variation in local population which is only related to the arrival of immigrants, and thus I am able to isolate the partial demand changes from other adjustments in demand due to other population relocation. To be able to recover unbiased estimates, besides using controls and fixed effects, I implement a shift-share-type IV strategy explained in the section 3.3. In the results section I show that the decomposition outlined in section 2 holds exactly in the data when the β coefficients of equations (10) to (12) are estimated consistently using the instrument described in the next section.

3.3 Construction of the immigration rate instrument

Even after including fixed effects and control variables, consistent estimation of the β coefficients requires the regressors of interest to be uncorrelated with the time-varying part of the error term. Unobserved factors could still induce omitted variables or endogeneity biases. There is no prior on

[16] González & Ortega (2013) do find a sizeable impact of immigrants on home construction in Spain for a similar time period than mine.

the direction of the bias. For example, in the case of the total impact (10), the estimated β_1 could be upward biased if immigrants are going to provinces with positive shocks or unobserved better economic prospects, while it would be downward biased if, for some reason, immigrants locate in province in which prices are growing slower (conditional on all the controls). To estimate consistent causal parameters I use an instrumental variables (IV) strategy.[17]

The instrument I construct is an imputation or prediction of the immigration rate: $\left(\frac{\Delta FB_{i,t-1}}{POP_{i,t-2}}\right)$, which is defined as the immigration inflow during $t-1$ divided by total population at the end of $t-2$. I construct the instrument adapting the "shift-share" methodology, which has extensively been used before, for instance by Card (2001) or Ottaviano & Peri (2006). Intuitively, a province-year immigrant stock imputation is constructed by distributing year-to-year variation of the stocks of immigrants by nationality (or by country of origin/birth) – the "shift" or "shock"– across different areas, using some location pattern –the "share"–, to allocate this magnitude. The immigration rate in province i that we want to instrument is:

$$IMM\_rate_{i,t-1} = \frac{\Delta FB_{i,t-1}}{POP_{i,t-2}} = \frac{FB\_inflow_{i,t-1}}{FB_{i,t-2} + NAT_{i,t-2}} = \frac{FB_{i,t-1} - FB_{i,t-2}}{FB_{i,t-2} + NAT_{i,t-2}} \tag{13}$$

The most commonly used shift-share instrument builds up on the fact that, to take advantage of social and economic established networks, immigrants tend to disproportionately locate in areas where immigrants from the same nationality or ethnicity have located before (ethnic networks instrument). I use past (1991) location patterns ("share" distribution) to predict current location patterns. For the national yearly immigration inflow ("shift") by nationality, I use country-of-origin-specific predicted inflows based on a gravity push factor model. The product of the shift-share produces annual imputations (*predictions*) of the stock of foreign-born for each nationality in each province. To compute imputed province $FB_{i,t}$ stocks in (13) I sum these imputations over nationalities.

To impute the annual immigrant population by province and by nationality of origin, I first calculate the share of immigrants that were located in that province in the base year.[18] This share corresponds the proportion of immigrants located in a particular province i over the total immigrants from the same nationality n in Spain in 1991 (sum r provinces over a total of $R = 50$):

$$share^n_{i,1991} = \frac{FB^n_{i,1991}}{\sum_r^R FB^n_{r,1991}} = \frac{FB^n_{i,1991}}{FB^n_{Spain,1991}} \tag{14}$$

The imputed foreign-born stock of a specific nationality n in province i at time t, $imp\_FB^n_{i,t}$ is calculated allocating yearly total national stocks ($FB^n_{Spain\_i,t}$) by nationality weighted by their historical location share calculated as (14). I then sum this across nationalities to calculate the total imputed

[17] Additionally, using instrumental variables would remove attenuation bias in the OLS and fixed effects estimates if there is substantial measurement error in the immigration numbers.

[18] The list of the nationalities used appears in Table B.5.

foreign-born stock in province i at time t, $imp\_FB_{i,t}$:

$$imp\_FB_{i,t} = \sum_{n}^{N} \left(share^{n}_{i,1991} * FB^{n}_{Spain\_i,t} \right) = \sum_{n}^{N} \left(imp\_FB^{n}_{i,t} \right) \tag{15}$$

Shift-share instruments need to be relevant (sufficiently correlated with the variable they predict) and conditionally exogenous uncorrelated with unobserved shocks) in order to be valid and yield consistent estimates. The relevance of the instrument can be assessed by the value of the F-statistic of the instrument in the first stage of the 2-stage-least-squares (2SLS) regressions, and additionally by using weak identification tests. Even if one could worry that the level of immigrant stocks in particular nationality-province pairs was too low in 1991, the F-statistics in the results tables show that the instrument is strong enough, even conditional on a large set of controls and province fixed effects (Jaeger et al., 2018).[19]

For the exogeneity condition to be met, both elements involved in the construction of province yearly immigrant predictions must be orthogonal to local shocks related to the outcome variable, conditional on controls. Regarding the local share of immigrants by nationality in the base year, the exclusion restriction requires that the only channel through which foreign-born geographical distribution in 1991 affects current changes in house prices/native locations is through its influence on shaping the current immigrants location patterns, conditional on controls and province and time fixed effects. In other words, the unobserved factors determining the location of immigrants in one province with respect to another in the base year (1991) have to be uncorrelated with the relative economic prospects of the two provinces during the period of analysis (2001-2012). I consider 1991 to be separated enough from 2001 for this condition to be reduced, and the province fixed effects to capture across-provinces heterogeneity correlated to past location patterns to a great extent. Nevertheless, in the robustness checks, I perform a battery if test to assess the validity of the choice of 1991 as base year.

We also require that the national total stock of immigrants for a given nationality in a given year, $FB^{n}_{Spain\_i,t}$, to be exogenous to specific province unobservable shocks. It is unlikely that shocks that have driven immigrants of a different nationalities to specific locations in a given year (what we are instrumenting for) are uncorrelated to shocks in neighbouring provinces. To solve this issue, a similar strategy to Saiz (2007), Ortega & Peri (2016) and Monras (2020) is adopted. I predict annual total stock and inflow of immigrants by country of origin from the results of estimating gravity migration model which depends only on push factors in origin. Details of this procedure are given in the appendix (section B.2.1). I use the models to obtain predictions of foreign-born stocks and inflows in years 2001 to 2012 for each nationality. I multiply this by the shares and sum across nationalities to obtain

[19] All the result tables in section 4 provide the Kleinbergen-Paap statistic (test of weak identification), which is robust to non-i.i.d error terms, and corresponds roughly to the t-stat of the included instruments in the first-stage to the square. In some tables I also include the first-stage coefficients.

province-year imputations of foreign-born stocks and inflows.

To compute the instrument for the immigration rate (13), in the numerator I use the imputed prediction of foreign-born inflows ($imp\_pred\_FB\_inflow_{i,t-1}$). In the denominator I need to compute a (lagged) prediction of population stocks. This is composed of foreign-born stocks and native stocks ($FB_{i,t-2} + NAT_{i,t-2}$ in equation (13)). For the first component, I can use the lagged imputed prediction of foreign-born stocks ($imp\_pred\_FB_{i,t-2}$). However, as discussed in section 3.2, the number of total natives residing in a given province might depend on the number of foreign-born in the same location or on unobservables correlated with house price growth. For this reason, I use a similar shift-share strategy to compute a prediction for the location of natives $imp\_NAT_{i,t-2}$, based on past location patterns. Details on this procedure are also given in the appendix (section B.2.2).

With all this, I compute the instrument as:

$$IV\_IMM\_rate_{i,t-1} == \frac{imp\_pred\_FB\_inflow_{i,t-1}}{imp\_pred\_FB_{i,t-2} + imp\_NAT_{i,t-2}} \tag{16}$$

Prediction 16 is used to instrument $\left(\frac{\Delta FB_{i,t-1}}{POP_{i,t-2}}\right)$ in equations (10) and (12) and to instrument $\left(\frac{\Delta POP_{i,t-1}}{POP_{i,t-2}}\right)$ in equation (11). I use $IV\_IMM\_rate_{i,t-1}$ in the main IV estimation results and different versions of it in robustness checks. In section 5.1 I discuss and test the validity of this IV approach and in section 5.2 I check if the results are robust to using different definitions of the shift and share in the construction of the instrument.

3.4 Data and descriptive statistics

To investigate the impact of immigration on housing costs, I exploit a panel of Spanish provinces for the period 2001-2012. This setup very suitable to study this question because during this period large immigration inflows were coupled with a period of housing sector boom (2001–2007) followed by a bust (2008–2012), which provides large variation in the data. Average annual rent and house price growth during the period was around 3 and 5%, while the average population growth rate was around 1%, mostly due to foreign-born population growth. Figure A.2 shows the evolution of immigrant stocks and inflows and of housing costs during these years. Foreign-born stock increased from around 2.5 million people in 2001 to 6.7 million in 2012, an increase of almost 160%. In these years, the share of foreign-born over total population rose almost 10 percentage points (from 4.8% to 14%). This increase was particularly remarkable for high-immigration provinces (as shown in panel (a) of Figure A.1. Between 2001 and 2008, the annual inflows of foreign-born were over 400,000 persons per year, and even after the start of the recession they remained between 70 and 100,000. With respect to the nationality of the immigrants, panel (b) in Figure A.2 shows that biggest immigrants inflows by origin stemmed from Latin American and Eastern Europeans, which had moderate presence in 2001, followed by EU-15 and North Africans. The change in the most important sending nationality groups could reduce the strength of the instrument, but it helps with concerns raised by Jaeger et al. (2018) and Goldsmith-Pinkham et al. (2020).

Until 2008, average local housing costs also increased considerably, in particular house sale prices. As the middle panels of the figure show, during the housing boom years average house prices increased around 108%, and even with the fall that followed, on average they increased almost 65% during the 2001–2012 period. For rents, the increase was of around 37% between 2001 and 2012, around 30% until 2008. The annual increase slowed down after 2008 but nevertheless remained slightly above 1%. [20]

Regarding the spatial distribution of the variables of interest, Figure A.3 shows the distribution by quintile of the share of foreign-born population in 2001 and 2012. We can see that some of the core economic provinces are in the top 2 quintile in 2001 (Madrid, Catalonia, Valencia, Basque Country); in tourism-oriented locations like the Islands, Malaga, Murcia and Alicante, where wealthy European foreigners locate; but also in poorer areas (Galicia, Extremadura), with large proportions of 50-60s out-migrant returnees. By 2012 we have new provinces with high proportions of foreign-born in Castilla and Aragon. The change in the spatial pattern in the location of foreign-born shows that we have much variation to exploit in the empirical exercise, and this can adopt a very demanding empirical strategy. The top maps of Figure B.1 depicts the location of the changes in foreign-born population between 1991 and 2001 (IV base year and start of the period of analysis) and 2001 and 2012 (start and end of period of analysis): we can observe variation in the location of the changes in both maps. Finally, the bottom maps of the figure show the spatial distribution of the growth of housing prices between 2001 and 2012. We observe remarkable differences for rental and sale prices, with a lot of the higher growth concentrated in areas where prices were traditionally lower (south and west of Spain). When put all together, there is no clear spatial pattern connecting the location of immigrants and housing cost growth so we need to turn to the regression analysis in order to be able to extract meaningful conclusion.

Table 2: Housing cost and population rates summary statistics

Variables	Time period	Mean	Std. Dev.	Min	Max
Change of log rental prices	2002/01–2012/11	0.028	0.017	-0.014	0.083
Change of log sale prices	2002/01–2012/11	0.047	0.094	-0.157	0.276
Inflow of population during t over population in (t-1)	2001–2011	0.010	0.012	-0.012	0.061
Inflow of foreign-born during t over population in (t-1)	2001–2011	0.009	0.008	-0.005	0.046
Inflow of natives during t over population in (t-1)	2001–2011	0.002	0.006	-0.017	0.031

To carry out the empirical analysis I used data from multiple sources. Immigrant and popula-

[20]The range of variation of rent price growth is much smaller than that of sale prices, as can be seen in the standard deviations in table 2. This is because average rental prices calculations are based on prices in properties currently rented (whose rents grows slowly or tied to national CPI indices) and on properties newly rented (where we see higher price increases when tenancy agreements change). Also, rent prices capitalise consumption of housing as a service, while sale prices growth also has an speculative component when housing is used as an investment asset.

tion data comes from the Municipal Population Registers (*Padrón Municipal*), which keeps an annual record of all registered individuals in a municipality over time regardless of their legal immigration status. This is the most reliable source to study the impact of the size of immigration on area economic outcomes. House house price data was obtained from Uriel-Jiménez et al. (2009), who provide an improved version of the Housing Department Average Province House Price Index. Data on rents was obtained combining data from the Housing Department and the National Institute of Statistics. Finally, data on the controls comes from several sources including the National Institute of Statistics, the Public Works (Housing) Department, the European Environmental Agency and the 2001 Census. Full details on the data sources are provided in section B.1 in an online appendix. Summary statistics for the main variables in the analysis are provided below in Table 2. The full list of controls is provided in the descriptive statistics table (Table B.4) and in the results tables notes.

4 Results

4.1 The total effect of immigration on house prices

In this section I present the results of the estimation of the total impact, e.g. β_1. Table 3 presents results of the estimation of equation (10) by ordinary-least-squares (OLS), for rents (panel A) and for house (sale) prices (panel B) [21]. Each column presents a specification that includes different sets of controls and fixed-effects. In all specifications the standard errors are clustered at the province level and robust to heteroskedasticity, and I include year fixed effects to control for national shocks. Specifications range from more to less demanding in terms of data variation: OLS results (column 1) to first differences province fixed effects with attribute trends model (column 5). The list of control variables is specified the notes of Table 3, and it is the same in all result tables unless specified. Coefficient β is a semi-elasticity and it can be interpreted as the growth of housing costs in percentage points for a 1 percentage point (0.01 units) increase in the rate. As explained above, in this case β corresponds the total demand estimate and captures the combined impact of changes in demand from immigrants and natives.

The first column of Table 3 shows the results obtained when we only include year dummies. It reports a simple correlation is 0.34 for rents and 0.6 for house prices. In columns 2 and 3 I add province attributes (time-invariant characteristics) and the province fixed effects (which are collinear with the attributes). The estimates increase substantially with respect to column 1. In column 4 I add the province attributes interacted with year dummies, which control flexibly by trends on the levels of the attributes correlated with immigration rates and housing costs. This is particularly relevant for characteristics which are quite different across provinces, like house tenancy and consumption patterns by natives and foreigners. The coefficients increase again and remain very similar when we

[21] As they are not informative, I do not report the coefficients for the control variables

Table 3: Total demand effect estimates – OLS\FE results

	(1)	(2)	(3)	(4)	(5)
PANEL A: Change in log rent prices t\t-1					
Immigration rate (t-1)	0.340**	0.482***	0.447***	0.592***	0.675***
	(0.130)	(0.134)	(0.126)	(0.202)	(0.203)
Adjusted R^2	0.57	0.60	0.68	0.68	0.68
PANEL B: Change in log sale prices in t\t-1					
Immigration rate (t-1)	0.604**	1.069***	1.468***	1.947***	2.024***
	(0.291)	(0.331)	(0.374)	(0.596)	(0.584)
Adjusted R^2	0.85	0.85	0.86	0.86	0.86
Province attributes		Yes			
Province FE			Yes	Yes	Yes
Province attributes * Year FE				Yes	Yes
Time-var controls (t-3\t-2)					Yes

Notes: Significance levels: * p<0.05, ** p<0.01, *** p<0.001. Each column presents results from a different specification. All regressions include year dummies and use 550 observations (50 provinces over 11 periods). t=2002\2012. Clustered (province) standard errors in parenthesis. *Time-varying controls* (lagged two periods e.g. t-2\t-3) include change of log current GDP per capita, change of log of number of credit establishments, change of unemployment rate, change average years of education employed, change share working-age-population without any degree, change of log transport infrastructure and change of log urban infrastructure. *Province attributes* (time invariant) include share of residential secondary dwellings, share of residential empty dwellings, share of households which own a secondary home, share of employed in construction sector, share of employed in services sector, share of foreign-born renters (residents in family homes), share of natives renters (residents in family homes), log average sqm dwelling per person foreign-born, log average sqm dwelling per person natives; all these in 2001. They also include log road distance to Madrid, length of coastline in 100s of kms, log of rain precipitation (January), share of developable land (Corine 2000) and average ruggedness index. *Province attributes * Year FE* interact the time-invariant characteristics with year dummies.

further add time-varying controls in column 5. The model in this last column, where we include time and province fixed effects, province attribute flexible trends and time-varying controls is the most demanding one, and the baseline specification in the rest of the paper. In the last column the estimated semi-elasticities are around 0.67 for rents and for sale prices is 2. Even if informative, these coefficients roughly correspond to partial correlations between prices and immigration rates.

In order to be able to make causal claims about the estimates, I implement the IV strategy explained in section 3.3. Table 4 presents the results using the instrument as defined in equation (16), using 1991 as base year and the gravity-predicted immigration inflows to construct the shift-share instrument. I depict the coefficients for the baseline specification of Table 3 column 5. Column 1 shows the semi-elasticity of rents and column 2 for house (sale) prices. I estimate the models using 2 stages least squares (Correia, 2018). The table shows the coefficients for the second-stage (panel A), which correspond to the total demand semi-elasticities and the first-stage estimates (panel B), which show the relationship between the predicted and instrumented immigration rates. I also report the weak identification test (F-stat Kleibergen-Paap), which informs about the relevance of the instrument and the mean values of the outcomes and the rates. All specifications show that the instrument is very

Table 4: Total and partial demand effect estimates – IV 2SLS results

	(1)	(2)	(3)	(4)	(5)
	Rent Prices (t)	Sale Prices (t)	Native Rate (t-1)	Rent Prices (t)	Sale Prices (t)
PANEL A: Second-Stage estimates					
Immigration rate (t-1)	0.986**	3.278**	0.308***		
	(0.465)	(1.236)	(0.088)		
Population rate (t-1)				0.754**	2.506***
				(0.353)	(0.881)
PANEL B: First-Stage estimates					
Immigration rate SSIV		0.684***		0.894***	
		(0.143)		(0.180)	
Weak identification test (KP)		22.94		24.78	
Mean Outcome (Y)	0.0285	0.0474	0.0016	0.0285	0.0474
Mean Rate (X)		0.0089		0.0105	
All province FE and controls	Yes	Yes	Yes	Yes	Yes

Notes: Significance levels: * p<0.05, ** p<0.01, *** p<0.001. Rents and sale prices are in log changes. All specifications include province FE, province attributes*YearFE and time-varying controls, as described in Table 3. Clustered (province) standard errors in parenthesis. Obs=550.

strong. As expected, in all specifications the standard errors increase when using IV.

Compared to the OLS fixed effect estimates, the IV coefficients are much larger. This suggests that immigrants are moving, conditional on the controls and the area fixed effects, to provinces which are experiencing negative shocks in the growth of rents, and therefore the estimates of Table 3 are downward biased. Given that we are controlling for a wide set of time-varying economic factors and time-invariant attributes, it is quite reasonable that, conditional on all these controls, immigrants locate in places where housing is more affordable. [22] In addition, as discussed above, the downward bias of the OLS estimates could be due to measurement error, either due to poorly measured raw population register number or due to the fact that the total foreign-born number masks substantial nationality-mix heterogeneity across provinces and the IV better captures the average treatment effect.

I find a semi-elasticity of around 1% (0.986) for rental prices and of 3.3% (3.278) for sale house prices, for an increase in the immigration rate of 1 percentage point. These numbers are very similar to existing IV estimated elasticities, such as Saiz (2007) (who finds 0.9 for rents and 3.2 for prices), Degen & Fischer (2010) who find 2.7 for Swiss prices, González & Ortega (2009) who find 3.2 for house prices and Ottaviano & Peri (2007) who find 0.7 for rents and up to 2 for prices.

The coefficient of the effect on house prices is larger than of rents, consistent with existing estimates in the literature. This is not surprising as in this period, due to the intense housing construction, the stock of for-sale housing is much more elastic than the one of rental properties. Hence, the scope

[22] González & Ortega (2013) and Farré et al. (2011) find the same direction for the OLS-IV bias.

for price growth and the mean value standard variation of the two prices are different (as shown in Tables 2 and 4), and higher for sale prices. If we standarise the semi-elasticities, the difference between the two effects become much smaller.

4.2 The effect of immigration on native location

In this section I discuss the estimate of the impact of immigration on native mobility, in order to assess the difference between the total and partial demand effects formalised in 2. It involves the IV estimation of coefficient β_3 in equation (12). The sign and size of this coefficient inform about the existence of native displacement or co-location (attraction). The estimated coefficient is shown in column 3 of Table 4, which indicates that for each 10 immigrants that settle in a province, around 3 natives relocate there due) to the immigration inflow. The time period of analysis is 2001–2011. The effect is estimated contemporaneously in order to match the equation (9), but in the following sections I also explore the timing of these adjustments. This non-negligible attraction estimate suggests that the difference between partial and total impacts would be sizeable.

My findings suggest that natives and immigrants are contemporaneously co-locating in the same provinces. The attraction or co-location estimate, although rare, has also been found in other papers (for example Mocetti & Porello, 2010; Wozniak & Murray, 2012). While most authors have argued the existence of "native´-fly", the empirical findings on native displacement are inconclusive (Amior, 2021). Peri & Sparber (2011) argued that displacement could not be as quantitatively relevant as previously thought, at least in the case of the US. The expected direction of the relocation effect might also depend on the geographical size of the unit of analysis and the underlaying characteristics of the locations (Larkin et al., 2019). [23]

One potential explanation for the attraction result is that immigrants might be complementary to natives, due to different tastes or skill levels, and thus positively affect their location decisions. Recently, the immigration impacts literature has focused in the research of these complementarities (Ottaviano, 2014). Besides enhancing productivity through improved task specialisation (Peri, 2012), immigrants might have desirable attributes for natives. For example immigrants could be specializing in producing goods and services which are desirable for natives (Ottaviano et al., 2013), increasing their consumption opportunities (Mazzolari & Neumark, 2012) or allowing female workers to increase their labour supply (Barone & Mocetti, 2011). In order to provide some intuitions on the co-location finding, section B.3 in the (online) appendix develops a simple theoretical framework where

[23] There exist some previous evidence that also points towards native-immigrant co-location in Spain. Fernández-Huertas et al. (2009) find a comparable result to mine for a long-differences estimation from population growth regressed on the immigration rate for the period 2001-2008. Their prediction is of 1 native for each 10 immigrants. They argue that this number is sufficiently small to have an impact on compensation or reinforcement of the impact of immigration inflows on the housing or the labour markets[24]. Fernández-Huertas et al. (2019), using finer spatial data, also find co-location of natives and immigrants in newly created suburban communities, and only find mild native displacement in neighbourhoods where housing supply is constrained.

natives and immigrants specialise in different sectors (high-skill natives in the tradable sector and low-skill immigrants in the non-tradable local services sector). In the model, an inflow of immigrants reduces the price of local services making locations more attractive to natives, who co-locate with the immigrants. I provide some correlations that show this mechanisms could be credible, especially in provinces which are receiving most natives and immigrants.

4.3 The partial effect of immigration on house prices

In this section, in order to estimate the partial demand effect, I apply the methodology described in section 3.2. I use population growth rate as the main regressor in equation (11) and instrument it with expression (16). This instrument predicts exogenous foreign-born location. Conditional on controls and fixed effects, the predicted-by-the-instrument population growth second-stage estimate only captures changes in population due to immigrant inflows. This coefficient captures the impact on house prices stemming from changes in foreign-born demand, abstracting from the induced demand due to other population changes. By doing this, the estimated coefficient corresponds to a direct immigrant demand elasticity (partial impact), independent from demand changes from relocated natives.

The results of this exercise are shown in columns 4 and 5 of Table 4. The instrument is very strong in all specifications, as shown in the weak identification test, and predicts almost 90% of population growth. The estimated semi-elasticities are 0.75% and 2.5%, for an increase in population growth (due to immigration) of one percentage point, for rents and sale prices respectively. If we combine the coefficients of columns 4 and 5 with the co-location estimate of column 3, we observe that the decomposition equation (9) hold exactly in the data. The partial demand semi-elasticities are almost 24% smaller than the total demand effect, due to the increase in demand induced by the native relocation process. For example, for sale prices the total demand impact of an increase of immigration in one percentage point is 3.3%, of which 2.5% is due to direct immigrant demand and 0.78% to additional demand for relocated natives. This insight is new in the literature, and highlights the importance of the framework laid out in section 2.

4.4 The timing of the adjustments

While the empirical strategy has so far focused on providing total and partial demand estimates, in this section, I explore the timing of the adjustments for all three sets of results. First I explore if using longer lags of data, instead of one-year differences as in the main results, changes the results. In Table 5, I use three (panel A) and five-year (panel B) long-differences (LD) to construct the immigration and population rates and the instrument. The outcome variables also correspond to 3 and 5-year growth rates. The top panel uses 150 observations and the bottom one 100. [25]

Table 5: The timing of the adjustments: long-time differences

	(1)	(2)	(3)	(4)	(5)
	Rent Prices (LD)	Sale Prices (LD)	Native Rate (LD)	Rent Prices (LD)	Sale Prices (LD)
PANEL A: 3-year differences					
Immigration rate (LD3)	1.523***	2.143*	0.208		
	(0.413)	(1.157)	(0.208)		
Population rate (LD3)				1.261***	1.774*
				(0.382)	(0.884)
Weak identification test (KP)		16.38		11.17	
PANEL B: 5-year differences					
Immigration rate (LD5)	1.027**	2.132**	0.430**		
	(0.405)	(0.986)	(0.172)		
Population rate (LD5)				0.718***	1.491**
				(0.261)	(0.656)
Weak identification test (KP)		20.54		14.07	

Notes: Significance levels: * $p<0.05$, ** $p<0.01$, *** $p<0.001$. The sample period is 2011/2001. Both the outcomes and the population/immigration rates are calculated using 3-year (2002/2005, 2005/2008 and 2008/2011, Obs=150) or 5-year (2001/2006 and 2006/2011, Obs=100) time differences. Clustered (province) standard errors in parenthesis. The estimations include region and time dummies, time-varying LD controls and province attributes, as described in Table 3.

The table reports the coefficients for the total (columns 1-2) and partial (column 5) demand effects and native mobility (column 3). The results show that the main results hold when allowing for longer periods for the adjustments to take place. The instrument is still strong and the Kleinbergen-Paap statistics well above the 10 rule-of-thumb threshold. The coefficients are slightly larger for rent prices and smaller for sale prices. The coefficient for native mobility is not significant in the top panel and slightly larger in the bottom one. If we use the coefficients of Table 5, the decomposition expression (9) is still valid. Overall, the results are similar to those in Table (16) and confirm my findings even when I allow for adjustments in prices over longer periods of time.

A second exercise to study the timing of the adjustments is to change the lag structure of the rates. The results are shown in Table 6. Here, I test the robustness of the results to using a contemporaneous, one lagged or two-lagged inflow of immigrants, using them one at a time or combined. I show the results for native mobility (panel A), and total demand effects (panels B and C). The baseline estimates are shown in bold for comparison purposes.

The table shows that for native mobility the lag that seems to matter most is the contemporaneous: immigration rate lagged one or two periods with respect to the native rate is insignificant, and when

[25]There are two main changes in the estimated specification. First, because the LD setting allows for the adjustments in prices from the inflows to take place at any time during the 3 or 5-year time period, I do not impose any lag structure in the specification. Therefore, in these results the long-difference changes in the outcomes and main regressors are contemporaneous. Secondly, given that I have fewer observations, I use a slightly less demanding specification than in Table (16), and use 17 regions (NUTS2) fixed effects (instead of province ones). However, I still include a large set of controls which are listed in the notes of the table.

Table 6: The timing of the adjustments: timing of effects - lags

	(1)	(2)	(3)	(4)	(5)
PANEL A: Native population rate in t					
Immigration rate (t-2)			-0.120		-0.180
			(0.114)		(0.116)
Immigration rate (t-1)		0.074		-0.048	
		(0.102)		(0.116)	
Immigration rate (t)	**0.308*** **			0.340***	0.307*
	(0.088)			(0.127)	(0.156)
Observations	550	500	450	500	450
Weak identification test (KP)	22.94	20.25	16.61	6.79	4.85
PANEL B: Change in log rent prices t\t-1					
Population rate (t-2)			0.567*		0.342
			(0.306)		(0.329)
Population rate (t-1)		**0.754****		0.482	0.690*
		(0.353)		(0.405)	(0.409)
Population rate (t)	0.989**			0.823*	
	(0.443)			(0.443)	
Observations	550	550	500	550	500
Weak identification test (KP)	13.26	24.78	24.09	8.19	9.54
PANEL C: Change in log sale prices t\t-1					
Population rate (t-2)			3.116**		2.890*
			(1.498)		(1.602)
Population rate (t-1)		**2.506*** **		2.336***	0.694
		(0.881)		(0.827)	(0.998)
Population rate (t)	2.055			0.512	
	(1.902)			(2.232)	
Observations	550	550	500	550	500
Weak identification test (KP)	13.26	24.78	24.09	8.19	9.54

Notes: Significance levels: * $p<0.05$, ** $p<0.01$, *** $p<0.001$. All specifications include province FE, province attributes*YearFE and time-varying controls, as described in Table 3, adjusted accordingly depending on the lag used. The sample period is 2002/2012 for house prices and 2001/2011 for native mobility. Clustered (province) standard errors in parenthesis. Obs=550.

we add both at the same time (contemporaneous and lagged), it is the contemporaneous one that matters. For prices, we see a different picture. For rental prices the contemporaneous and lagged rates all affect rents growth, as we would expect as immigrants consume housing from the moment they settle. For sale house prices only the lagged rates affect price growth, which would be explained by the fact that it takes more time for immigrants to be able to purchase homes. [26]

[26]It is worth mentioning that in panel estimations such that of table 6 it is difficult to exactly pin-down the timing of the effects, as we are exploiting within province variation (conditional on many controls). This exercise then is similar to a horse-race regression where we investigate which lag seems to have a stronger explanatory power, but should not be read as an infallible test of the time-structure to be used in the specification.

5 Validity tests

5.1 IV strategy discussion

In this section I discuss the validity of the IV strategy implemented. Two conditions must apply for the shift-share prediction to be an appropriate instrument. For the exclusion restriction to be valid, conditional on all controls, the only channel through which the predicted immigrant stocks affect the housing costs growth must be via its effect on current immigrant stocks. This implies that historical settlement pattern of immigrants by nationality/country of origin in the base year (share component in equation (15) has to be sufficiently lagged that, conditional on controls, it is orthogonal to unobservables correlated with current housing costs growth (exogeneity condition). At the same time, the instrument has to be sufficiently strong in its prediction of the current immigrant location patterns (relevance condition). I provide two pieces of evidence to test these conditions. The first one relates to the exogeneity of the instrument and the second one, to its relevance. Finally, I also correlate the instrument with changes in house sale prices before the period of analysis, particularly during 1994-1998 where there was mild housing boom. This exercise aims to test for the existence of pre-trends.

The base year for the construction of the instrument, 1991, is 10 years before my observation period starts, which is substantially longer than in other applications[27]. However, one could still think of unobservable shocks correlated with housing costs and location decision of immigrants that existed in 1991 that still affect both aggregates today (even conditional on all the region trends and province attributes/fixed effects). To test this, similarly to Farré et al. (2011), I regress the share of foreign-born population in 1991 on 1990-91 economic factors and then the change in this share during my observation period (2001-2012) on the same variables. The aim of this exercise is to show that the determinants of the geographical distribution of the mass of immigrants in 1991 does not perfectly predict their location during my period of study. The results are shown in Table 7.

The explanatory variables include the log of disposable income, the log of average wage (region wage bill over workers), the share of different sectors (construction, services and industry) in the regional value-added (the exclude category is the agriculture sector), the unemployment rate for natives and foreign-born workers, the log housing density (number of residential housing units per square km) and the share of built-up land over total land (to control for urbanisation) and additional controls related to geography (area, coast dummy and length of coastline and distance to Madrid). In column 1, the model has high predictive power (R^2 is around 0.82) and most of the regressors are significant.[28] When I regress this same set of variables on the change of share of foreign-born popu-

[27] For example Saiz (2007) uses data from 1985–1998 and the base year is 1983 and Sá (2015) uses data from 2003 and the base year is 2001.

[28] Note, this are partial correlations and the coefficients of individual variables have to be interpreted conditional on everything else.

Table 7: IV validity checks: Base-year validity regressions

	(1) Share of FB in 1991	(2) Change share FB 2001\2012
Log disposable income	-0.026**	-0.032
	(0.010)	(0.052)
Log of average wage	-0.013	-0.006
	(0.020)	(0.095)
Share of GVA construction	0.200**	-0.580
	(0.091)	(0.533)
Share of GVA services	0.113***	0.011
	(0.036)	(0.333)
Share of GVA manufacturing	0.085**	-0.027
	(0.036)	(0.288)
Natives unemployment rate	-0.030	-0.165
	(0.028)	(0.179)
Immigrants unemployment rate	-0.076**	-0.175
	(0.037)	(0.185)
Log homes per sqkm	0.038***	0.042
	(0.012)	(0.066)
Share of built-up land over total	-0.283*	0.025
	(0.146)	(0.460)
Log area	0.022**	0.043
	(0.010)	(0.054)
Coast dummy	-0.009	0.015
	(0.006)	(0.032)
Length of coastline (100s of kms)	0.001***	0.001
	(0.000)	(0.004)
Log road distance to Madrid (kms)	0.006	-0.006
	(0.004)	(0.011)
Constant	0.014	0.183
	(0.143)	(0.746)
Observations	50	50
Adjusted R2	0.82	0.43

Notes: Significance levels: * p<0.05, ** p<0.01, *** p<0.001. Source data: Census 1991, Spanish Regional Accounts, Corine Landcover 1990. GVA stands for Gross Value Added. Economic values in 1991, share of build-up land in 1990. The omitted category is share agricultural GVA. 50 observations, one per province.

lation over the 2001-2012 period none of the coefficients is significantly different from zero and the explanatory power of the model is much lower. This test is supportive of the appropriateness of using 1991 as base year. In case we consider 1991 still too close to the start of the period, in the robustness checks provided in the next section, I also use 1981 as base years to construct the instrument (which remains strong) and the results remain fairly similar.

The second instrument validity exercise relates to the relevance of the instrument. I construct alternative instruments with placebo shifts or placebo shares combining them with the actual share or shifts (those used in the construction of the instrument 16). By placebo I refer to using shares and shifts which I expect to have little strength predicting current immigrant location patterns. I use these placebo instruments to check if the results hold. The purpose of this exercise is to prove that it is the exact combination of the (nationality-specific) 1991 location patterns and the gravity predicted

Table 8: IV shift-share placebos: native mobility and partial demand effects

	(1)	(2)	(3)	(4)	(5)	(6)
	Baseline	**Change Share**		**Change Shift**		
	c91&grav	random	c1940	USA	AUS&NZ	NAT_GR
PANEL A: Native population rate (t-1)						
Immigration rate (t-1)	0.308***	0.391**	0.151	-0.349	0.284	0.123
	(0.088)	(0.188)	(0.274)	(0.952)	(0.259)	(0.252)
Weak identification test (KP)	22.94	3.24	3.80	0.76	3.14	4.87
PANEL B: Change log rent prices in t\t-1						
Population rate (t-1)	0.754**	1.737*	-0.256	-0.972	0.312	1.337
	(0.353)	(1.013)	(1.250)	(7.072)	(0.615)	(1.423)
Weak identification test (KP)	24.78	3.23	2.30	0.18	2.34	2.80
PANEL C: Change log sale prices in t\t-1						
Population rate (t-1)	2.506***	3.128	-2.358	-17.823	6.615	5.368
	(0.881)	(1.974)	(4.692)	(49.532)	(5.419)	(4.917)
Weak identification test (KP)	24.78	3.23	2.30	0.18	2.34	2.80
All province FE and controls	Yes	Yes	Yes	Yes	Yes	Yes

Notes: Significance levels: * p<0.05, ** p<0.01, *** p<0.001. All specifications include province FE, province attributes*YearFE and time-varying controls, as described in Table 3. Clustered (province) standard errors in parenthesis. Obs=550.

national inflows that produces a reliable strong instrument. I show the results of this exercise in Table 8, only for native mobility (panel A) and the partial demand estimates results (panels B and C). The first column shows the baseline estimates, which uses the instrument where the share is based on the 1991 provincial foreign-born stocks by nationality and the shift is based on the gravity estimates of Table B.1.

In columns 2 to 4 I change the province share and interact it with the gravity national prediction (baseline shift). In column 2 I randomly distribute each nationality immigrants across provinces and I multiply this random allocation by the gravity-model predicted annual inflow by nationality. The results are significant for native mobility but weakly and not-significant for house prices. In column 3, I use nationality-province immigrant stock information from the 1940 census. This shows that such old past location patterns do not have any predictive power for the current ones. The instruments used in columns 2 and 3 are very weak, with KP tests below 4.

In the last three columns of the table I use the 1991 share (baseline share) and interact it with a placebo annual shift by nationality. In columns 4 to 5, I use the immigrants by nationality (inflow and stock) going to other rich countries which are very far and have different production structures than Spain. I use inflows to the USA (column 4) and to Australia and New Zealand (column 5), and I find the instrument becomes very weak and coefficients very small and insignificant. Finally, in column 6 I use the annual growth in province population predicted by the natural movement of population (births-deaths), and distribute it based on the total share of foreign-born population in 1991 in each province. The coefficients of interest are also highly insignificant and the instrument is weak. These results show that is the precise the combination of a relevant share and a relevant shift that gives raise to a strong instrument that predicts the annual location of immigrations by nationality in each of the provinces.

In the last validity exercise I test if the instrument correlates with pre-trends in the outcome variable, only with house sale prices as there is no pre-2001 data for province rental prices. I use information from the Valuation Society (*Sociedad de Tasación*) for several years between 1990 and 2000. In this decade there was an economic crisis (1990-1994) followed by high (1994-1998) and moderate (1998-2000) house price recovery. In Table A.2, I correlate the instrument with growth rates in these three periods, either using the whole panel and year-by-year data (columns 1-3) or a 10-year long-difference version of the instrument and one cross section. The results of the table show there is no correlation between price dynamics in the 1990-2000 decade and our instrument, which strengthens the validity of my instrumental variables strategy.

5.2 Robustness checks

In this section, I present additional results in order to check the robustness of the findings. I focus on the native mobility and partial estimates coefficients, but the results are also robust for the total demand estimates. The robustness results for rents are displayed in Table 9. In the first column I show the baseline estimate in both tables shows the baseline elasticity estimate for comparison purposes.

Column 2 uses foreigners instead of foreign-born population. A fraction of the Latin American immigrants that settled in Spain held Spanish passports and were able to settle as nationals, so the number of foreigners is smaller than the number of foreign-born immigrants. The coefficient using this measure is very similar for all three outcomes, with the instrument (for which we use 1991 foreigner settlement patterns) stronger than in the baseline results. In columns 3 and 4 we change the base year for the computation of the instrument, with location patterns based in 1981 (older) and 2001 (more recent) ethnic networks. The instrument is weaker using 1981 information and stronger using 2001, as we would expect, and in both cases with a KP-test over 10. With 1981 shares the result for rents does not hold, but for native mobility and sale prices is very similar to the baseline. With 2001 shares all three results hold, and the coefficients are larger than in the baseline. However it is unlikely that the exclusion restriction holds with 2001 settlement patterns. Given this, the preferred base year remains 1991. In column 5 I use a different grouping of countries to construct my instrument, 49 countries instead of the 104 nationalities depicted in table B.5. The coefficients are very similar to column 1 and the instrument remains very strong.

Finally in column 6 I use an alternative instrument based on the gateways/ports of entry.[29] The

[29] In a similar manner to González & Ortega (2013).

Table 9: Robustness checks: native mobility and partial demand effects

	(1)	(2)	(3)	(4)	(5)	(6)
	Baseline	Foreigners	c1981	c2001	49COB	Gates
PANEL A: Native population rate (t-1)						
Immigration rate (t-1)	0.308***	0.299***	0.429**	0.396***	0.291***	0.368***
	(0.088)	(0.084)	(0.178)	(0.110)	(0.089)	(0.100)
Weak identification test (KP)	22.94	28.25	8.95	42.01	19.49	29.03
PANEL B: Change log rent prices in t\t-1						
Population rate (t-1)	0.754**	0.881**	-0.011	0.839***	0.694*	0.902**
	(0.353)	(0.367)	(0.492)	(0.290)	(0.369)	(0.367)
Weak identification test (KP)	24.78	30.92	11.38	51.57	19.98	42.01
PANEL C: Change log sale prices in t\t-1						
Population rate (t-1)	2.506***	2.519***	2.278**	3.191***	2.540**	3.409***
	(0.881)	(0.925)	(1.054)	(0.805)	(0.957)	(1.041)
Weak identification test (KP)	24.78	30.92	11.38	51.57	19.98	42.01
All province FE and controls	Yes	Yes	Yes	Yes	Yes	Yes

Notes: Significance levels: * $p<0.05$, ** $p<0.01$, *** $p<0.001$. All specifications include province FE, province attributes*YearFE and time-varying controls, as described in Table 3. Clustered (province) standard errors in parenthesis. Obs=550.

intuition is that different nationality immigrants will locate disproportionably in regions which are more accessible to them. I first locate 50 ports of entry using 6 different travel modes (listed in Table B.6). For 113 nationalities I calculate the share of immigrants in 2000 that used those different modes of transportation using data from the National Immigrant Survey 2007. This gives me nationality-specific variation which is necessary in order to avoid perfect collinearity with the province attributes and fixed effects of specifications (10)–(12). Then, for each province in Spain I calculate a weighted-by-road-distance and port size (using data on air and boat passengers in 2001) nationality-specific accessibility index. I calculate a weighted measure of how accessible a province is for each nationality from all the ports of entry, where the numerator is the port-size and the denominator is share of migrants that use that particular mode. I normalise this province accessibility measure and use it to distribute the nationality-specific gravity-model inflow/stock in every year from 2001 to 2011. Using the model with province fixed effects, province attribute flexible trends and all the controls the instrument is very strong and the results are very similar to the baseline. It is very reassuring to find very similar results using a complete different share based on accessibility.

A second robustness check is to test if leads of the immigration and population rates affect past outcomes. The results for this are provided in Table A.3. Column 1 shows the baseline estimates for native displacement (contemporaneous) and population rate (lagged one period). In columns 1 and 2 I use one lead with respect to the outcome variable. i.e. (t+1). We find no impact on mobility and weak impacts for sale and rental prices. When we use two leads (t+2) we find no effect in either outcome. Columns 4 and 5 use the leads in conjunction with the baseline lag for the outcomes and we find that the chosen time structure seem to be the one which remains significant even after controlling for leads, a bit less so for rental prices. All in all, these falsifications checks provide additional robustness to the findings.

In the online appendix I present some additional results using different population groups and different measures of the outcomes. The results also remain very similar.

6 Conclusions

This paper draws the attention to a highly overlooked issue in the estimation of average area effects of immigration: the role of local population displacement on the adjustment of local demand and prices. The total impact of increases of foreign-born population on housing markets results from a combination of their direct impact on housing demand, their impact on native mobility and their impact on housing supply. The estimation of these well-identified reduced-form total effects is with no doubt of interest for policy makers. However, as well as the net impact, we might be interested on understanding the mechanisms driving its adjustment. Previous research has shown the importance of taking into account the role of (population) displacement when estimating the impact of local policy interventions (Blundell et al., 2004; Mayer et al., 2017; Einiö & Overman, 2020). Null impacts of a local policy might be the results of positive and negative impacts cancelling each other out over space. Hence, it is important to account for the impact that mobility has on the estimation of aggregate local estimates. As I show in this paper, this issue is also important when assessing the impact of immigration on house prices, as the immigration wage also alters the spatial distribution of native population. Mine is the first paper that provides joint estimates of the total price impact of immigration and of all its components and, in doing so, it sheds light in the adjustment channels of local house prices to population shocks.

A better understanding of the impact of immigrant demand on housing costs is of great interest to urban economics. Howard (2020) shows that immigration amplifies labour market shocks via its impact on housing demand. Incorporating better knowledge into the adjustment of housing markets is crucial when analysing the impact of immigration on urban labour markets and how demographic changes affect house prices (Gong & Yao, 2022). Second, large immigration inflows impact local population generating shocks that dissipate across locations. (Monras, 2020) provides seminal work embedding the impact of immigration on housing markets into a quantitative spatial equilibrium framework. He provides new insights on the role of low-skilled immigrants, who work largely in the construction sector and in the long run reduce construction costs and house prices. In contrast, my paper highlights the importance of accounting for native population relocation and leaves the door open for further work incorporating this channel of adjustments into formal models. Finally, housing capital was a very important driver of wealth accumulation in Spain during the 2000s (Blanco et al., 2021). Correctly understanding the role of the large immigration inflows on housing stock capitalisation is also crucial when aiming to study wealth accumulation over time and locations, in order to better disentangle the price changes related to housing consumption demand from those related to investment or changes in supply.

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Appendix

A.1 The role of housing supply

Depending on adjustment of housing supply, increases in housing demand following the immigrant inflows could have different effects on average local house prices. If housing density is increasing or many new dwellings are being built, the impact of increased demand on prices could be mitigated. How housing construction costs are affected can have very important effects on the evolution of prices, particularly in the medium to long-run, as highlighted by Monras (2020), and compensate increases in demand resulting in slower growth of prices. On the one hand, as discussed in section 3.2, housing density remained very stable during the time period of analysis. On the other hand, the increased in population was coupled with a large number of new housing units built, as shown in table 1. As controlling directly for changes in housing supply is problematic, I first start by adding controls and trends that capture socio-economic, geographic and other aspects related to house construction (see notes in Table 3 for the exact list of variables). Thus, the results of Table 4 include a large set of housing supply-related province attributes or province fixed effects which capture time-invariant supply characteristics and thus allow the interpretation of $\hat{\beta}_1$ and $\hat{\beta}_2$ as demand coefficients.

In this section, I explore further the role played by housing supply on potentially mitigating the increase in prices. I directly investigate the mitigating impact that time-varying supply changes might have on house prices by including the growth in the stock of dwellings as an additional control variable. I control for changes in housing supply in order to remove potential omitted variable bias from the demand estimates. I am not interested on the specific estimated coefficient of this variable but on the effect that introducing it has on the immigration and population rate coefficients.

Table A.1: Effect of controlling for housing supply on the estimates (partial)

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Change log rental prices in t\t-1				**Change log sale prices in t\t-1**			
Population rate (t-1)	0.754**	0.638*	0.581*	0.601*	2.506***	2.537***	2.451***	2.457***
	(0.353)	(0.321)	(0.317)	(0.322)	(0.881)	(0.859)	(0.854)	(0.842)
Log change housing stock (t-2)		0.229*		0.302*		-0.061		0.097
		(0.121)		(0.159)		(0.535)		(0.834)
Log change housing stock IV (t-2)			-8.596**				-2.754	
			(4.076)				(23.649)	
Test weak-identification POP RATE	24.78	25.54	25.27	12.77	24.78	22.35	22.14	12.77
Test weak-identification HSUPPLY				49.33				49.33
All province FE and controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Notes: Significance levels: * p<0.05, ** p<0.01, *** p<0.001. All specifications include province FE, province attributes*YearFE and time-varying controls, as described in Table 3. Clustered (province) standard errors in parenthesis. Clustered (province) standard errors in parenthesis. Obs=550

The results are presented in Table A.1. Columns 1–4 show the results for rents and columns 5–

8 for (sale) prices. Columns 1 and 5 repeat the results of the baseline specifications of Table 4. In columns 2 and 6 I introduce changes in log housing stock (in $t-2$) as an additional control variable. This variable is lagged two periods with respect to the outcome in the same manner as the time-vaying controls, but using one lag (t-1) generates very similar results. This coefficient is significant at 10% level for rents on the left panel (total impact) but insignificant and small for sale prices. The coefficients remain very similar even if I remove all the supply-related province attributes from the specifications or if I use the change in (t-1).

However, using the observed growth of housing stock as an additional control variable is highly problematic. Even if lagged two periods with respect to the outcome variable, and one period with respect to the immigrant and population rates, this variable is likely to be endogenous. Unobservable province trends could be affecting both the growth in prices and the construction of new housing units, particularly in a context of housing market boom where there were expectations of high capital gains. In fact, as the results in González & Ortega (2013) suggest, immigrants also have a direct impact on dwelling construction, so the growth of housing stock is a "bad" control by definition. To deal with this, I use an instrument for the stock of housing in a given province. I construct a similar predictor as in Saiz (2010), a shift-share type instrument for changes in housing stock, by combining province national share of developable land in year 2000 (initial spatial distribution – the share) and changes in total annual national stock (excluding the own province changes – the shift).[30]

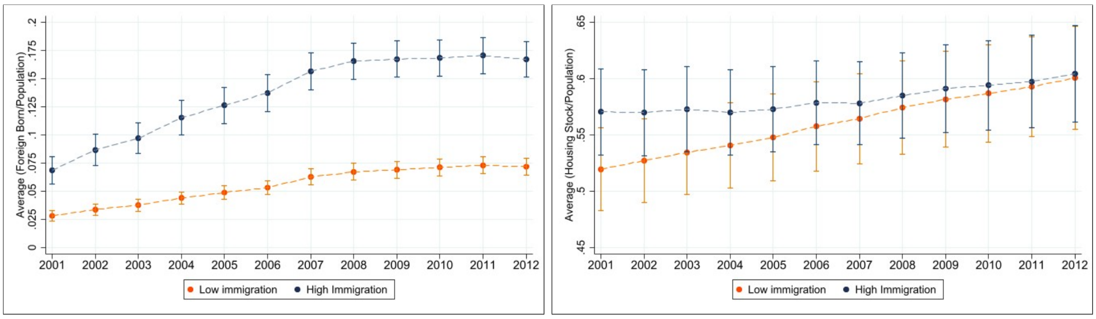
In columns 3/7 I use the predicted change in log housing stock IV as an additional control and in 4/8 I use it to instrument for the time-varying change in housing supply. The housing-supply IV is very strong (Kleinbergen-Paap values over 49). All along the coefficients of the impact of the rates on housing costs remain very similar to the baseline estimates, particularly for house sale prices, and lose a bit of significance for rental prices. These results suggest that changes in housing supply did not have any major additional impacts on the demand estimates (neither the total nor the immigrant demand ones). We need to keep in mind that there are already a large set of controls included in the specifications that are likely already capturing a substantial share of the variation due to changes in housing supply. It is then not so surprising to find insignificant coefficients for this variable. In addition, Table 1 shows that the very high intensity of dwelling construction was coupled to a large population growth such that the house per person rate remained almost constant during the period. Even if many houses were constructed during this period, the population inflows were so substantial that increases in supply did not decrease the pressure of demand on housing costs. Given that directly controlling for housing supply does not seem to have any large effect on the baseline estimates I believe that the set of controls and fixed effects I already include in the regressions do a good job

[30] I exclude the own province stock from the national stock for two reasons: to avoid using the exact figure I am trying to instrument for in the calculation of the predicted stock and to have province-specific time variation that is not fully collinear with the included time and province fixed effects. The results are very similar if we use 1990 availability of developable land. Full details on what is included in developable land are provided in an online appendix in section B.1.1.

partialling out the impact of supply changes and this allows me to interpret my estimates as demand coefficients.

A.2 Appendix tables and figures

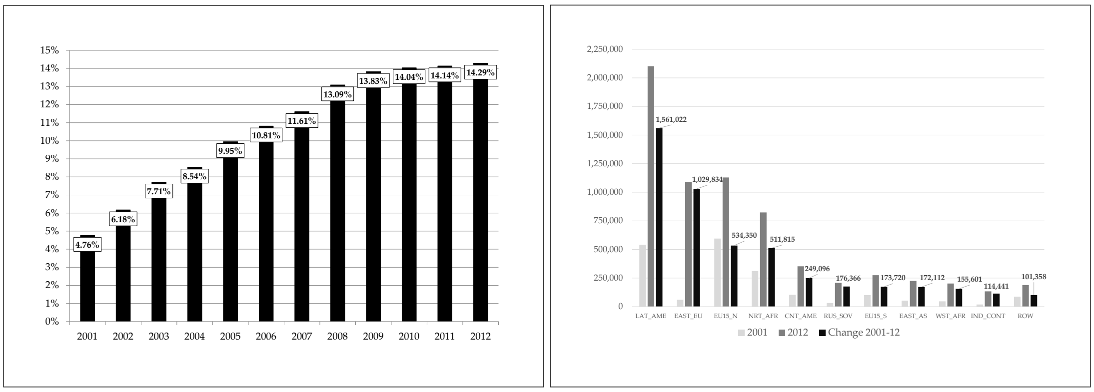
Figure A.1: Share of foreign-born population and housing stock per person – 2001-2012



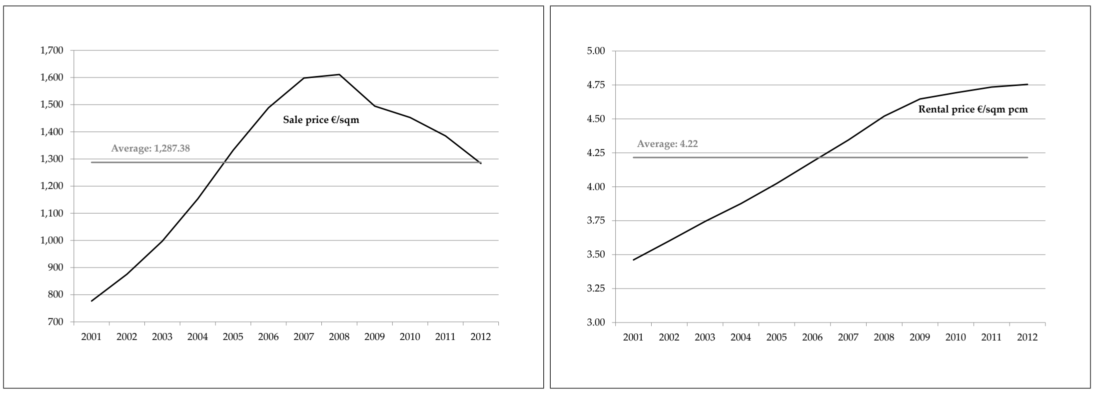
(a) Average FB/population (b) Average dwellings/person

Sources: Spanish Department of Housing and Annual Population Registers. High/low immigration status is defined by being above or below the median in the foreign-born inflow over population during the 2001-2012 period (0.0783). Points depict province averages in the variable, while bars show the 95% confidence intervals.

Figure A.2: Foreign-born shares and house price growth over time



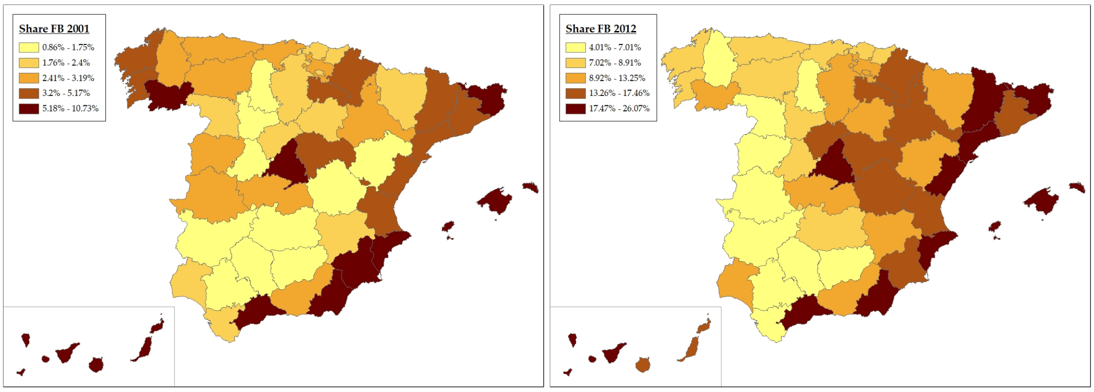
(a) Share of foreign-born population (b) Foreign-born stocks and inflows by nationality



(c) House sale price growth (d) House rental price growth

Sources: Spanish Department of Housing, IVIE and Annual Population Registers

Figure A.3: Spatial distribution of shares of foreign-born 2001 and 2011



(a) Share of FB in 2001

(b) Share of FB in 2012

Sources: Annual Population Registers

Table A.2: Correlations between Immigration rate SSIV and pre-period House Prices

	(1)	(2)	(3)	(4)	(5)	(6)
Change Log HP between	1990\94	1994\98	1998\00	1990\94	1994\98	1998\00
Immigration rate IV (t-1)	-5.848	-0.354	2.153			
	(3.582)	(1.712)	(1.358)			
Immigration rate IV (LD10)				0.070	-0.103	0.070
				(0.355)	(0.648)	(0.355)
Observations	550	550	550	50	50	50
Model	Panel	Panel	Panel	LD10	LD10	LD10
Sample	2001-11	2001-11	2001-11	2011	2011	2011
R^2	0.10	0.04	0.07	0.62	0.45	0.62

Notes: Significance levels: * p<0.05, ** p<0.01, *** p<0.001. Each column shows the coefficients of running a different model. The main regressor is the immigration instrument, either year predictions (1-33) or long-difference (4-6). The outcomes of the models are the change in log (growth rates) of province average house prices for periods 1990-94, 1994-98 and 1998-2000, using data from the Valuation Society (*Sociedad de Tasación*). Columns 1 to 3 use the 2001-2011 panel and control for province attributes, (province attributes * year) and time-varying controls as in table 3, but not for province FE (as the outcomes are time invariant). Clustered (province) standard errors in parenthesis. Columns 4-6 regress the 10-year long-difference of the instrument on the growth of house prices. These models include province attributes and 10-year LD time-varying controls. Robust standard errors in parenthesis.

Table A.3: Additional results: timing of effects - leads

	(1)	(2)	(3)	(4)	(5)
PANEL A: Native population rate in t					
Immigration rate (t)	**0.308*****			0.386***	0.360***
	(0.088)			(0.125)	(0.091)
Immigration rate (t+1)		-0.032		-0.301	
		(0.171)		(0.202)	
Immigration rate (t+2)			0.196		-0.087
			(0.194)		(0.180)
Observations	550	550	500	550	500
Weak identification test (KP)	24.94	18.84	11.90	6.94	4.20
PANEL B: Change in log rental prices t\t-1					
Population rate (t-1)	**0.754****			0.443	-0.342
	(0.353)			(0.401)	(1.122)
Population rate (t+1)		1.990*		1.258	
		(1.012)		(0.808)	
Population rate (t+2)			1.259		2.050
			(0.809)		(2.322)
Observations	550	500	450	500	450
Weak identification test (KP)	24.78	5.21	5.33	4.36	0.95
PANEL C: Change in log sale prices t\t-1					
Population rate (t-1)	**2.506*****			2.189***	1.741
	(0.881)			(0.766)	(2.139)
Population rate (t+1)		4.297*		2.217	
		(2.392)		(2.000)	
Population rate (t+2)			2.913		0.677
			(1.898)		(4.650)
Observations	550	500	450	500	450
Weak identification test (KP)	24.78	5.21	5.33	4.36	0.95

Notes: Significance levels: * p<0.05, ** p<0.01, *** p<0.001. All specifications include province FE, province attributes*YearFE and time-varying controls, as described in Table 3. The sample period is 2002/2012 for house prices and 2001/2011 for native mobility.

Online Appendix

B.1 Data

B.1.1 Data sources

The spatial unit of analysis in the paper is the province (NUTS3). I exclude Ceuta and Melilla because of their particular history and lack of data. I use data from 2002 to 2012 – the population data is dated in January so it refers to the beginning of the period.

I use data on total, foreign-born and native population from the Spanish population municipality registers (yearly). The number of residents in a municipality is registered by the city councils in an administrative register called the Municipal Register (*Padrón Municipal*). An annual record of the municipal register, dated on the 1^{st} January of each year, is obtained from its updates. This dataset provides precise information on the population figures, on a yearly basis. It is also more accurate than other population sources because it collects the total number of foreign-born residents even if they are illegal immigrants[31]. Immigrants are identified using foreign-born population (by country of birth), not nationality. The figures are dates at the beginning of the natural year (1^{st} of January).

Even if this data is available since 1996, I focus on the period after 2001 for several reasons. First, Fernández-Huertas et al. (2009) and Bertoli et al. (2011) recommend the use of population data coming from the population registers (*Padrón*) from 2001 because its reliability improves after that year. Secondly, it is after 2001 that the stock of foreign-born starts increasing significantly. It could be the case that most entries started in 2001 or that the stocks started to be correctly measured after that year. To mitigate measurement error I then focus on 2001-2012 for the main analysis. Thirdly, the rents data is only available from 2001 so focusing on this time period allows us to compare the rental and sale prices results over the same time period. Finally, using the housing boom and bust allows adoption of a demanding estimation strategy as there is more variance in the house price growth data.

House price data comes from Uriel-Jiménez et al. (2009), published by the Valencian Institute of Economic Research (henceforth IVIE) jointly with the BBVA Foundation (FBBVA). The database covers the period 1990-2007 and the IVIE prices are calculated using the original data from the (previously) Spanish Housing Department (*Ministerio de Vivienda*). The Housing Department (now inside Public Works Department) official data provides the average price per square meter on dwellings sales in the private sector. It is provided every quarter for all the provinces. The IVIE dataset of house prices is constructed by re-weighting the official prices provided by the Housing Department to take into account the location of the dwelling and when it was built. As the IVIE data is only available

[31] However, it has two disadvantages. For confidentiality issues, data availability on the characteristics of the population is limited (only age, gender and nationality). In addition, the immigration figures may be over-estimated because immigrants have to actively cancel their register when they move out of the country (if they move within the country their new register cancels out the old one). For this reason, it is a good source to study the effect of immigration inflows but not so good for outflows.

until 2007, the dataset was expanded to 2012 by applying the provincial price growth rates from the Housing Department official data series and adjusting the series for changes in base years. In any case, the empirical results are robust to using the Housing Department official data series. Data on rental prices comes from the Housing Department and the National Institute of Statistics (*INE*). I combine data from the National Observatory of Rented Properties (*Observatorio Estatal de la Vivienda en Alquiler*) and the consumer price indices (CPI provinces - rents component) to calculate the average rental price per square meter of the each province, from 2001 to 2012.

As a control and to construct the housing supply changes instrument I use the share of developable land in the province. The share of developable land in 2000 is obtained combining "developable" categories from the EU EEA Corine Land Cover 2000 dataset. Total area and total developable area[32] were calculated using GIS and raster maps of land use year for 2000, provided by the *Corine* Land Cover data project (European Environment Agency). I calculated the stock of dwellings in the different years combining data from the Spanish Housing Department. Data on the housing stock is available from 2001. Using the entry and exit flows, I calculated a rate of depreciation and I updated the stock of the dwellings combining the depreciation rate and construction of dwellings data.

I also use time-invariant province characteristics (attributes) in the specifications without province fixed effects. Summary statistics for these variables (and the year in which they are measured) provided in Table 2.

The data sources for the attributes and time-varying controls variables are diverse and most of them were obtained from the Spanish National Statistics Institute (INE) and the 2001 Census. Other sources include the National Geographical Institute, *La Caixa* Spanish Economic Yearbook (*La Caixa Anuario Económico de España*), Banking Annual Yearbook (*Anuario de la Banca*), Bank of Spain Statistics (*Banco de España*), the 1991, 2001 and 2011 Population and Dwelling Censuses, the Housing Department (now Public Works Department), the IVIE-BBVA Human Capital Statistics (*Estimación de las Series de Capital Humano 1964-2013*), the IVIE-BBVA Regional Capital Stock Statistics *(Series históricas de capital público en España y su distribución territorial (1900-2012))*, the Spanish Regional Accounts (*Contabilidad Regional de España*).

Data for the IV validity regressions comes from the OECD International Migration Statistics and the Spanish National Statistics Institute. Data for the construction of the gateways instrument comes from the Spanish Port Authority (*AENA*), the INE National Immigrant Survey 2007 (*Encuesta Nacional de Inmigrantes*), the National Ports Statistics Yearbooks (*Anuarios Estadísticos Puertos del Estado*) and an online road atlas for ports-entry distances. The list of ports was selected looking at several sources

[32]The categories included in developable land are: Green urban areas, Non-irrigated arable land, Permanently irrigated land, Rice fields, Vineyards Fruit trees and berry plantations, Olive groves, Pastures, Annual crops associated with permanent crops, Complex cultivation patterns, Land principally occupied by agriculture, Agro-forestry areas, Broad-leaved forest, Coniferous forest, Mixed forest, Natural grasslands, Moors and heartland, Sclerophyllous vegetation and Burnt areas.

with the main ports, airports, stations and roads and looking at those that were larger, busier and closer to the border countries.

B.2 Further details on the construction of the instrument

B.2.1 Gravity estimations

In order for the instrument to be valid, both terms in expression (15) have to be orthogonal to local shocks related to immigration inflows and house price growth. Local shocks have a direct impact on total immigration inflows to Spain as these depend on national shocks which are just a combination of local shocks. For this reason, instead of directly using national inflows by nationality in (15), I construct a prediction based on factors that are plausibly exogenous to local shocks. Following Saiz (2007) and Ortega & Peri (2012), I use a gravity-type model that only contains push-factors from origin to predict the total inflow from nationality n to Spain in a given year t.[33] The estimated equation is:

$$\ln\left(FB\_inflow_{from\_n\_to\_Spain,t}\right) = \rho' \ln\left(ECON_{n,t-1}\right) + \gamma_n + \lambda_t + \xi_{n,t} \tag{B.1}$$

where $ECON_{n,t-1}$ is a matrix of (lagged) time-varying economic conditions of the sending country (log of gross domestic output in real terms, log of total population, percentage of urban population, percentage of internet users, indexes of globalisation, conflict and governance controls and dummies of belonging to the EU27/EUROZONE). I include year dummies λ_t and country-specific dummies γ_n. The countries of origin are listed in Table B.5. I estimate a similar model using foreign-born stocks on the left hand side (in this case the economic variables are lagged two years because population is measure on the 1^{st} of January). The time period used to obtain the coefficients is 2001 to 2012.

The economic and institutional variables come from the World Bank World Development Indicators and Governance, the Globalisation Indices come from the Swiss Economic Institute (KOF Globalization Index), information on EU and EUROZONE membership comes from Wikipedia, the Major Episodes of Political Violence variables come from the Centre for Systemic Peace. Data is available for the 113 countries (the source countries in the Population Registers data are 119, including 6 categories for "other" and "stateless/unknown") of Table B.5, which represent more than 99% of the inflows into Spain for the period.

Results for different specifications are showed in Table B.1, for the total national inflows (column 1) and for the national foreign-born stocks (column 2). The specifications include country dummies and year dummies. All the models have high predictive power. From the results in Table B.1 I recover the predicted inflows to and predicted stocks of foreign-born in Spain from nationality n for every year 2001-2012. I use the prediction from estimates from column 1 for the construction of the

[33] And equivalently for imputed predicted stocks.

Table B.1: Gravity equations: immigrant inflow\stock by country

	(1)	(2)
	(Log) Inflow during t	(Log) Stock in Jan t
L1 or L2 Log current	-0.314	0.282
GDP in Bill USD	(0.528)	(0.227)
L1 or L2 Log of GDP	0.498	-0.03
Deflator	(0.558)	(0.275)
L1 or L2 Log population	-0.033	-2.497***
in 1000s	(1.266)	(0.577)
L1 or L2 Life expectancy	0.165**	0.080**
in years	(0.071)	(0.033)
L1 or L2 Percentage of	-3.602*	-1.543
agricultural land	(1.944)	(0.945)
L1 or L2 Percentage of	4.34	-0.194
urban population	(2.825)	(1.375)
L1 or L2 Internet take-up	-1.018	-1.229***
per 100 people	(0.632)	(0.260)
L1 or L2 Unemployment	4.107**	-0.354
rate total	(1.817)	(0.722)
L1 or L2 Share Services	-4.715**	-1.404
in Value-Added	(1.957)	(1.309)
L1 or L2 Share Industry	-5.577**	-1.979
in Value-Added	(2.129)	(1.465)
L1 or L2 Index political	0.01	0.004
globalisation	(0.008)	(0.004)
L1 or L2 Index social	0.006	-0.002
globalisation	(0.014)	(0.007)
L1 or L2 Index economic	0.014	0.010**
globalisation	(0.011)	(0.005)
Source country belongs to	0.442*	0.346***
the European Union	(0.239)	(0.128)
Source country belongs to	-0.205	-0.085
the Euro Zone	(0.326)	(0.137)
Observations	1,134	1,356
Adjusted R2	0.87	0.98

Notes: Significance levels: * p<0.05, ** p<0.01, *** p<0.001. Clustered (country) standard errors in brackets. t=2001/2012. The number of countries in the sample is 113. Note that sometimes country inflows are zero or negative so the number of observations in column 1 is smaller than in column 2. All models include country-specific fixed effects and year dummies. All models include Political/Social/Economic Globalisation Indices, WB Governance indicators and Major Episodes of Political Violence variables (coefficients not reported). EU/EUROZONE membership dummy changes over time as new countries join the Union. The economic explanatory variables are lagged one (L1) or two (L2) periods depending on the outcome variables used (inflow or stocks).

instrument, and I use the rest of the specifications estimates for the robustness checks of section 5.2. I use the coefficients of the gravity model to predict immigrant inflows and stocks by nationality in each year of the period of analysis. For example, the imputed predicted foreign-born inflow for each nationality n to each province i at time t becomes:

$$imp\_pred\_FB\_inflow^{n}_{i,t} = \left(pred\_FB\_inflow^{n}_{Spain,t}\right) * share^{n}_{i,1991} \tag{B.2}$$

To obtain the total imputed predicted inflow to each province i at time t we sum (B.2) across nationalities:

$$imp\_pred\_FB\_inflow_{i,t} = \sum_{n}^{N} \left(imp\_pred\_FB\_inflow^{n}_{i,t}\right) = imp\_pred\_\Delta FB_{i,t-1} \tag{B.3}$$

I use the lagged (B.3) in the construction of instrument (16).

B.2.2 Prediction for native location

I use past census data to predict the numbers of natives residing in province i in year t. Total natives in a province are the sum of those born and residing there and those who were born somewhere else in Spain and have moved there. I use a strategy that follows the same intuition as the shift-share immigration instrument. In contrast to the immigrants prediction, in this case we need to predict both magnitudes, i.e. stayers and movers. Therefore, we need to define a historical share and a time-varying shift for both types of natives. Instead of countries, the origin-destination geographical units are now the Spanish provinces. I use the province of birth of the native in the same way as the nationality in the case of foreign-born. The strength of the instrument is now based on the historical (im)mobility persistence of different Spanish locations (for stayers) and the "ethnic" networks (for movers). Some regions have historically had larger mobility propensities (Galicia), and some bilateral internal migration flows are based on historical location patterns (for example Galicians in Madrid or Andalusians in Cataluña).

A person born in a given province b can either stay where he/she was born (stayers) or can move and reside in a different province $i \neq b$ (movers). R is the total number of provinces in Spain in which natives can locate. For consistency, I use native location patterns from census 1991 as base year. I define the share of stayers in province i as the proportion of natives born and living in a province over all the natives born in the province (regardless of where they reside) in 1991 (I also use 1981 in the robustness checks). In this case, the province of birth and residence is the same, i.e $i = b$. The stayers share is defined as follows:

$$share^{b}_{i(i=b),1991} = \frac{natives^{b}_{i=b,1991}}{\sum_{i}^{R} natives^{b}_{i,1991}} \tag{B.4}$$

Share (B.4) is multiplied by the total natives that are living in the same province where they were born in year t. This gives the predicted number of stayers in a given province i in year t.

The share of movers is calculated differently. For a given province of birth b there are 49 potential province destinations where the mover can reside. I therefore need to calculate further 49 shares which represent the proportion of movers residing in a specific province i over the total number of movers originating from province b. The movers share is defined as proportion of natives born in b but residing in i over all the natives born in b but residing somewhere else:

$$share^{b}_{i(i\neq b),1991} = \frac{natives^{b}_{i\neq b,1991}}{\sum^{R}_{i\neq b} natives^{b}_{r,1991}} \tag{B.5}$$

Share (B.5) is multiplied by the total number of natives living outside the province they were born in year t (subtracting the natives living in the province for which we want to calculate the prediction, similarly to the case of the foreign-born prediction). This predicts the number of natives born in b living in province i (where $i \neq b$) in year t. For a given province of birth, there are 49 movers predictions.

To obtain the number of natives living in each province i at time t, I sum the prediction for stayers and the 49 predictions for each potential province of residence (movers) in each year. This gives $imp\_NAT_{i,t}$ which is used in the construction of (16).

B.3 Theoretical Model

In this section I propose a simple spatial equilibrium model to explain the results presented in section 4. I find that the total effect is larger than the immigrant demand because of increased induced demand by relocated natives and that both effects are positive. While the literature largely assumes that natives would be displaced from a region where a large inflow of immigrants arrive, I find a strong robust co-location effect. Recent theoretical and empirical developments have focused on analysing the impact of immigrants on the labour markets from a general equilibrium approach where we take into account the specialisation, skill-mix and technological changes that come about with immigrant inflows (Lewis & Peri, 2015; Peri, 2016). Some authors have proposed models predicting positive impacts of immigration on firm performance via production complementarities (Ottaviano & Peri, 2012b) and thus a native wage-enhancing effect of immigration. Existing theoretical models of the impact of immigration on house prices allow for potential co-location of immigrants and natives if some sort of complementarity (via production or consumption) exists (Saiz, 2007; Sá, 2015).

The model proposed below is a special case of a general model with two type of workers: native and foreign-born. When each type of worker specialises in one sector (tradable or non-tradable) and a change in the conditions in the source country pushes immigrants to the receiving country, the model predicts an increase on local rents and native and immigrant contemporaneous co-location. This result is consistent with the empirical findings of the previous sections and with existing evidence for Spain that native and workers specialise in different sectors (Amuedo-Dorantes & de la Rica, 2011; De la Rica et al., 2014) and that foreign-born workers provide cheaper local services (González

& Ortega, 2010; Farré et al., 2011). The model can be also be generalised to include a second tradable good which is produced only with foreign-born labour, in which case the model predicts that an inflow of immigrants increases rents, decrease native wages and crowds-out native labour in the receiving region[34].

B.3.1 Model Set-up

I present here a one-region (in our case province) model where a worsening in economic conditions in the country of origin generates an inflow of immigrants into the region. This fits well with the empirical instrumental variables strategy used in this paper where the inflow of immigrants is predicted using a push-factors gravity model (which captures the changes in conditions in origin) and the specific number of immigrants that locate in a particular region depends on local factors. The model is suitable for an immigration wave where immigrants move from their country of origin to a specific region and not when they move within regions of the receiving country[35].

In this set-up the spatial equilibrium for immigrants is determined by comparing the receiving city conditions with those of the sending country. After the inflow of foreign-born population settles in the region, how the wages, native population and rents (housing costs) react depends on whether the foreign-born population is employed in the tradable sector or not. In the case presented in this section foreign-born workers only work in the non-tradable services sector, while all the tradable good is produced with native labour. This assumption generates model predictions which fit the empirical results obtained above.

Let us assume a region r where two goods are produced, a tradable one C and a non-tradable one S. The third good is non-tradable housing H, which is just an endowment. There are two types of labour in the economy: native labour L^n and foreign-born labour L^f. All the tradable good is produced with native labour and all the non-tradable good is produced with foreign labour[36]. The production functions of these goods are:

$$Y_c = (L^n_c)^{\alpha}(N_c)^{1-\alpha} \tag{B.6}$$

$$Y_s = A_s L^f_s \tag{B.7}$$

where Y_c is the total production of good C, Y_s is the total production of good S, N_c is the second factor of production in the tradable good sector and α and $(1-\alpha)$ are the input shares in the Cobb-Douglas tradable good production function. A^s is the non-tradable good specific productivity shifter

[34] A full depiction of this model and its predictions is available upon request. The model could also be extended to add the Ottaviano & Peri (2012a) set-up where natives with different levels of skills would specialise in different sectors and the inflow of (low-skilled) immigrants would have differential effects on the wages and housing consumption decisions for both types of natives (high and low skilled).

[35] Which was the case during the immigration wave in Spain during the 2000s.

[36] In here I discuss the tradable and non-tradable sector produced with native and foreign-born workers but we could also think about a tradable high-skill and non-tradable low-skill sector where immigrants would provide low-skill labour.

(exogenous).

Individuals, both foreign and natives, yield utility from the consumption of three goods: the tradable good C, the non-tradable good S and housing H. The utility function is Cobb-Douglas and the shares of consumption of the goods are β, γ and $(1-\beta-\gamma)$:

$$u = c^{\beta} s^{\gamma} h^{1-\beta-\gamma} \tag{B.8}$$

Natives earn wages w_n and foreign-born workers earn (nominal) wages w_f, and they have no other sources of income. Rents (housing costs) are denoted by π. Prices of goods are given and denoted by p_c and p_s. The endowments of the second production factor and of housing stock are denoted by $\overline{N}_c$ and $\overline{H}$. The outside-option level of utility for native and foreign-born workers are given by $\overline{u}$ and $\overline{u}_f$. The relationship between the native and foreign-born baseline utility level is given by expression (B.23) and discussed in section B.3.3.

The equilibrium of the economy is characterised by the following equations:

$$\begin{aligned}
&\textit{ZPC1:} \quad p_c = \frac{w_n}{\alpha}\left(\frac{N_c}{L^n}\right)^{\alpha-1} && \text{(B.9)}\\
&\textit{ZPC2:} \quad p_s = \frac{w_f}{A_s} && \text{(B.10)}\\
&\textit{MCC1:} \quad w_f L^f = \gamma\left[w_n L^n + w_f L^f + \pi\overline{H}\right] && \text{(B.11)}\\
&\textit{MCC2:} \quad \pi\overline{H} = (1-\beta-\gamma)\left[w_n L^n + w_f L^f + \pi\overline{H}\right] && \text{(B.12)}\\
&\textit{SEC1:} \quad \overline{u} = \frac{w_n}{p_c^{\alpha} p_s^{\gamma} \pi^{1-\beta-\gamma}} && \text{(B.13)}\\
&\textit{SEC2:} \quad \overline{u}_f = \frac{w_f}{p_c^{\alpha} p_s^{\gamma} \pi^{1-\beta-\gamma}} && \text{(B.14)}
\end{aligned}$$

Equations (B.9) and (B.10) are the zero profit conditions (ZPC) for the tradable and non-tradable goods C and S, and equations (B.11) and (B.12) are the market clearing conditions (MMC) for the non-tradable goods S and H. Equations (B.13) and (B.14) define the spatial equilibrium conditions (SEC) where the optimality conditions for the maximisation problem of consumption for natives n and foreign-born workers f are equal to their outside-option level of utility $\overline{u}$ and $\overline{u}_f$.

By combining the MCC (B.11) and (B.12) we obtain:

$$w_f L^f = \frac{\gamma}{1-\beta-\gamma}\pi\overline{H} \tag{B.15}$$

$$w_n L^n = \frac{\beta}{1-\beta-\gamma}\pi\overline{H} \tag{B.16}$$

From the ZPC of the tradable sector (B.9) we derive the demand for native labour, decreasing in w_n:

$$L^n = \left[\frac{p_c\alpha}{w_n}\right]^{\frac{1}{1-\alpha}} N_c \tag{B.17}$$

By replacing (B.17) into (B.16) we obtain:

$$\frac{(p_c\alpha)^{\frac{1}{1-\alpha}} N_c}{w_n^{\frac{\alpha}{1-\alpha}}} = \frac{\beta}{1-\beta-\gamma}\pi\overline{H} \tag{B.18}$$

Using the SEC (B.13) and (B.14) into (B.18) we obtain:

$$\frac{(p_c\alpha)^{\frac{1}{1-\alpha}} N_c}{\left(\frac{\overline{u}}{\overline{u}_f}w_f\right)^{\frac{\alpha}{1-\alpha}}} = \frac{\beta}{1-\beta-\gamma}\pi\overline{H} \tag{B.19}$$

From the ZPC (B.10) and combining it with the SEC (B.14) we obtain:

$$w_f = \left[\frac{p_c^{\beta}}{A_s^{\gamma}}\right] \overline{u}_f^{\frac{1}{1-\gamma}} \pi^{\frac{1-\beta-\gamma}{1-\gamma}} \tag{B.20}$$

Combining equations (B.19) and (B.20) and solving for rents we have the following:

$$\pi = \Phi_1 \overline{u}_f^{-\frac{\gamma\alpha}{(1-\gamma)(1-\alpha)+\alpha(1-\beta-\gamma)}} \tag{B.21}$$

Finally, using equations (B.20) and (B.21) we can solve for the foreign-born workers wages w_f:

$$w_f = \left[\Phi_2 \overline{u}_f^{\frac{(1-\gamma)(1-\alpha)+(1-\gamma)\alpha(1-\alpha-\beta)}{(1-\gamma)(1-\alpha)+\alpha(1-\beta-\gamma)}}\right]^{\frac{1}{1-\gamma}} \tag{B.22}$$

where Φ_1 and Φ_2 are functions of prices, parameters and endowments.

B.3.2 Equilibrium adjustments to immigrants inflows

A decrease of the outside-option $\overline{u}_f$ for the foreign-born workers has the following effects:

$$\begin{aligned}
\Downarrow \overline{u}_f &\longrightarrow \Downarrow w_f && (\textit{equation B.22}) \\
\Downarrow \overline{u}_f &\longrightarrow \Uparrow \pi && (\textit{equation B.21}) \\
\Uparrow \pi &\longrightarrow \Downarrow w_n && (\textit{equation B.18}) \\
\Downarrow w_n &\longrightarrow \Uparrow L^n && (\textit{equation B.17}) \\
\Downarrow w_f \;\; \Uparrow \pi &\longrightarrow \Uparrow L^f && (\textit{equation B.15})
\end{aligned}$$

Intuitively, a worsening in the (economic) conditions in the countries of origin of the foreign-born workers decreases the value of their outside option ($\overline{u}_f$) and workers would migrate to the region. For a given housing stock and housing quality, housing rents π increase. Equation (B.21) predicts that the increase in foreign-born population decreases nominal wages w_f and thus make the non-tradable services cheaper (ZPC B.10). Foreign-born workers would be willing to work in the region for lower *real* wages. They enter the region until (equation B.14) is in equilibrium.

The decrease in the price of local services makes the region more desirable for natives, who enter the region until (equation B.13) is in equilibrium again. This in turn increases rents even further and decreases native nominal wages[37]. However, the *real* wages of natives are unchanged as local non-tradable prices have become cheaper after the foreign-born inflow. In the new equilibrium the region has increased foreign-born and native population, higher rents and lower nominal wages for natives and foreign-born[38]. The real wages for natives are the same or higher than before (depending on how much immigrants decrease local service prices and what share are these on the consumption of natives). Appendix section B.3.3 presents a brief model extension to explain why immigrants of different origins might locate in provinces in different proportions, providing a theoretical micro-foundation for the use of the ethnic networks shift-share instrument.

B.3.3 Model extension: explaining heterogenous location patterns

The discussion in section B.3 gathers all foreign-born workers under the common subscript f so it applies to each source-country separately. But in reality the province foreign-born population is the sum of immigrants from a variety of nationalities, and in each region the composition of the inflow and stock of foreign-born workers can be different. The average impact of the immigration wave on the average housing costs depend on the local mix of nationalities. In order to allow for heterogeneity in the location patterns of immigrants by nationality we define the foreign-born workers outside-option $\overline{u}_f$ as follows:

$$\overline{u}_f = \frac{\overline{u}}{\psi(\frac{1}{\overline{u}_f^*}, A_r)} \tag{B.23}$$

For foreign-born workers, the baseline (native) outside-option utility level $\overline{u}$ is normalised by the function ψ which is a measure of the relative "attractiveness "of a particular region r relative to the foreign-born worker's source country. This measure of relative attractiveness is itself related to two arguments:

[37]In the long run housing supply can change, either directly with the construction of new dwellings, or indirectly by increasing the density inside the dwellings (quality of housing can also be affected as noted by Saiz & Wachter (2011)). To make the model more intuitive we assume that housing stock is fixed. In the empirical exercise the results suggest that conditioning out for changes in housing supply does not change the demand impact of immigration on prices so this assumption is compatible with the empirical framework.

[38]Which are not necessarily the same as they work in different sectors and their level depends on parameter values.

- the "un-attractiveness "of the source country measured by $\frac{1}{\overline{u}_f^*}$ where $\overline{u}_f^*$ can be interpreted as some measure of utility available in the source country.

- the "attractiveness "of each region r to foreign-born workers, codified in an (exogenous) amenity endowment A_r. This can be interpreted as an ethnic specific amenity, say the presence of a pre-existing community *"à la Card"*.

Assuming super-modularity about ψ assures that attractive locations become relatively more attractive when push factors in the source country become stronger. This is given by:

Assumption 1: Monotonicity of ψ

$$\frac{\partial \psi(.,.)}{\partial(\frac{1}{\overline{u}_f^*})} > 0$$
$$\frac{\partial \psi(.,.)}{\partial(A_r)} > 0$$

Assumption 2: Supermodularity of ψ

$$\frac{\partial^2 \psi(.,.)}{\partial(\frac{1}{\overline{u}_f^*})\partial(A_r)} > 0$$

This set-up provides micro-foundation for why $\overline{u}_f$, even in spatial equilibrium, might differ across regions (i.e. the complementarity of push factors in the source-country and region-specific pull factors such as the presence of a pre-existing community). In consequence, for the same push factors in two different source countries, immigrants of different nationalities would locate in different proportions across different regions. When the attractiveness of source countries falls ($\overline{u}_f^*$ falls), high A_r locations attract large numbers of immigrants and all the mechanisms highlighted in section B.3.2 operate. For low A_r locations, the effects are weaker for the same fall in $\overline{u}_f^*$. In addition, this expression also formalises the channels through which the *shift-share* instrument discussed in section 3.3 predicts the changes in foreign-born population of a particular nationality in each province and year: a combination of an exogenous time-varying nationality specific push-component (predicting the number of immigrants coming to Spain in a given period – the shift) and time-invariant nationality-specific province preference (predicting the amount of immigrants that will locate in a particular region – the share).

B.3.4 Suggestive evidence in model results

The model is consistent with empirical findings for Spain suggesting that immigrants and natives specialise in different occupations (Amuedo-Dorantes & de la Rica, 2011), that during the immigration boom industries changed their skill-mix as a consequence of the low-skill inflow (particularly the sectors retail, construction, hotels and restaurants and domestic services) (González & Ortega, 2010) and that immigrants decreased the price of local services (Farré et al., 2011). I below show some simple results to add some evidence of the channel of adjustment at play, namely that regions where immigrants locate experience lower growth in non-tradables, and this is particularly important in provinces where more natives relocate.

First I construct an province-year non-tradable goods price index using the CPI components by consumption subgroup in each province and its weights. I keep the relevant services that I believe to be non-tradable and more related to sectors where immigrants work, and re-weight them to be representative by province and year. I use data for the Spanish National Institute for years 2001 to 2011. The index includes hospitality services (restaurants, hotels) and social services, where immigrants are disproportionably employed. These services, even if non-tradable, are more elastic than housing, so we can expect to find smaller impacts.

Table B.2: Relationship between non-tradable price growth and immigration – 10-year LD

	(1)	(2)	(3)	(4)	(5)
Outcome: Growth Non-tradable Price index LD10					
Immigration rate (LD10)	0.053	-0.608***	-0.742***	-0.714***	-1.475*
	(0.349)	(0.053)	(0.228)	(0.000)	(0.763)
Sample	All Provinces	Top 50 Natives	Top 70 Natives	Top 50 Nat+Imm	Top 70 Nat+Imm
Observations	50	25	35	18	27
Weak identification test (KP)	8.62	8.24	4.93	.	0.21

Notes: Significance levels: * p<0.05, ** p<0.01, *** p<0.001. This table depicts coefficients regressing province 10-year price growth for a non-tradable goods and services on the immigration ratio (instrumented). I include time-varying controls and province attributes.

I compute the 10-year long difference of this index and regress it with a 10-year difference of the immigration rate, instrumented with the 10-year rate IV prediction. I show the results in Table B.2. Column 1 shows the results for all 50 province. I also estimate the model in the subsample of provinces where I expect the effect to be concentrated, e.g. the top 70 and 50% provinces where more natives locate (cols 2-3) and the top 70 and 50% by location of immigrants and natives during the 2001-2011 period (cols 4-5). Even if the instrument is not very strong, the correlations when I focus on the most affected provinces (cols 2-4) are negative and significant, providing suggestive evidence that immigration inflows are correlated with lower growth of non-tradable services.

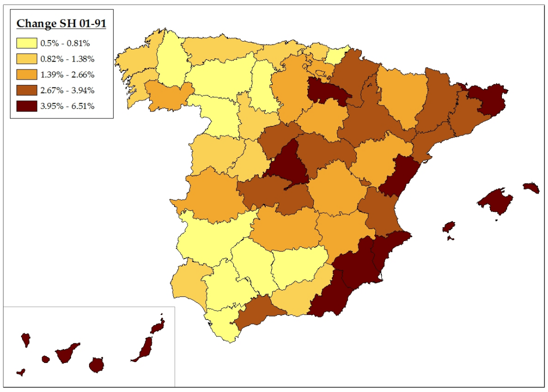
B.4 Additional figures and tables

Table B.3: Population stocks and share foreign-born – 2001-2012

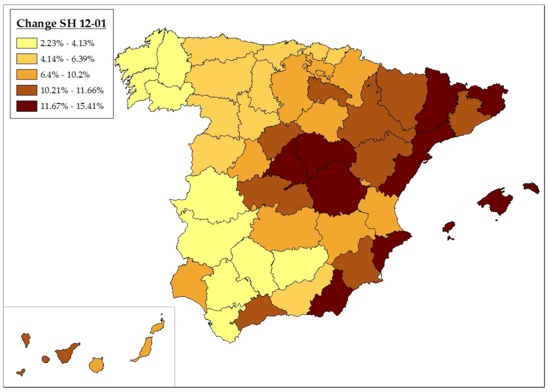
Year	ALL PROVINCES			LOW IMMIGRATION			HIGH IMMIGRATION		
	Natives	*Foreign Born*	*Average share FBorn*	*Natives*	*Foreign Born*	*Average share FBorn*	*Natives*	*Foreign Born*	*Average share FBorn*
2001	39,117,562	2,574,996	0.048	15,498,819	441,913	0.028	23,618,743	2,133,083	0.069
2002	39,290,201	3,283,469	0.060	15,512,597	532,181	0.034	23,777,604	2,751,288	0.087
2003	39,379,643	3,675,371	0.067	15,515,382	590,312	0.038	23,864,261	3,085,059	0.097
2004	39,592,858	4,374,908	0.080	15,548,571	692,673	0.044	24,044,287	3,682,235	0.116
2005	39,746,666	4,819,566	0.088	15,571,949	774,669	0.049	24,174,717	4,044,897	0.126
2006	39,824,392	5,230,302	0.095	15,574,697	844,952	0.053	24,249,695	4,385,350	0.137
2007	39,985,892	6,023,093	0.110	15,587,903	993,708	0.063	24,397,989	5,029,385	0.157
2008	40,150,524	6,443,149	0.116	15,612,779	1,075,759	0.067	24,537,745	5,367,390	0.166
2009	40,285,297	6,579,121	0.118	15,626,799	1,112,634	0.069	24,658,498	5,466,487	0.168
2010	40,378,949	6,650,692	0.120	15,622,363	1,149,433	0.071	24,756,586	5,501,259	0.169
2011	40,370,390	6,730,111	0.122	15,576,385	1,180,610	0.073	24,794,005	5,549,501	0.171
2012	40,352,241	6,609,683	0.120	15,521,352	1,156,733	0.072	24,830,889	5,452,950	0.167

Notes: Spanish Department of Housing and Annual Population Registers. High/low immigration status is defined by being above or below the median in the foreign-born inflow over population during the 2001-2012 period (0.0783).

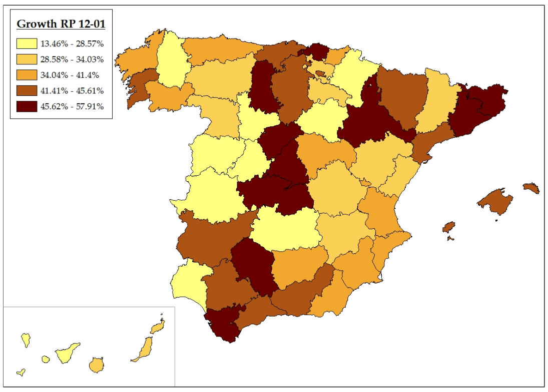
Figure B.1: Spatial distribution of changes in share of foreign-born, rents and prices growth



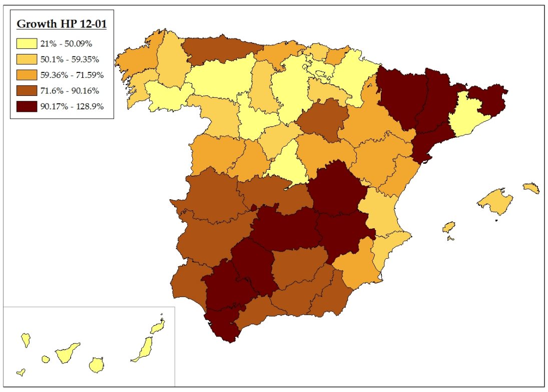
(a) Change in share of FB in 1991/2001



(b) Change in share of FB in 2001/2012



(c) Growth of rental prices 2001-2012



(d) Growth of sale prices 2001-2012

Sources: Spanish Department of Housing, Observatorio del Alquiler, INE, IVIE and Annual Population Registers.

Table B.4: List of variables with summary statistics

Variables	Time period	Mean	Std. Dev.	Min	Max
Time-varying controls					
Change of log current GDP per capita	2000/99-2010/09	0.042	0.038	-0.086	0.132
Change of log of number of credit establishments	2000/99-2010/09	0.005	0.030	-0.126	0.098
Change of unemployment rate	2000/99-2010/09	0.003	0.030	-0.134	0.097
Change average years of education employed	2000/99-2010/09	0.124	0.183	-0.678	0.857
Change share WAP without any degree	2000/99-2010/09	-0.013	0.034	-0.267	0.230
Change of log transport infrastructure	2000/99-2010/09	0.074	0.043	-0.035	0.370
Change of log urban infrastructure	2000/99-2010/09	0.069	0.054	-0.038	0.357
2001 Census characteristics					
Share of residential secondary homes	2001	0.352	0.090	0.154	0.568
Share of residential empty homes	2001	0.148	0.024	0.085	0.191
Share of households which own a secondary home	2001	0.138	0.042	0.073	0.225
Share of employed in construction sector	2001	0.126	0.021	0.083	0.170
Share of employed in services sector	2001	0.599	0.061	0.505	0.761
Share Foreign Born renters (residents in family homes)	2001	0.383	0.144	0.122	0.608
Share Natives renters (residents in family homes)	2001	0.079	0.026	0.044	0.159
Log average sqm dwelling per person Foreign Born	2001	3.487	0.084	3.245	3.622
Log average sqm dwelling per person Natives	2001	3.395	0.121	3.145	3.665
Province attributes					
Log road distance to Madrid	Time-invariant	5.792	0.648	3.517	6.581
Length of coastline in 100s of kms	Time-invariant	1.568	2.820	0.000	14.280
Log of rain precipitation (January)	Time-invariant	3.712	0.583	2.777	5.364
Share of developable land (Corine)	Time-invariant	0.854	0.074	0.466	0.961
Average ruggedness index	Time-invariant	110.314	48.249	31.030	247.614

Table B.5: List of countries of origin

Austria	Slovakia	Congo	Philippines
Belgium	Albania	Ethiopia	Indonesia
Denmark	Bosnia Herzegovina	Guinea Bissau	Japan
Finland	Croatia	Kenya	South Korea
France	Slovenia	South Africa	Thailand
Iceland	Serbia and Montenegro	Zaire	Vietnam
Liechtenstein	Macedonia	Costa Rica	Bangladesh
Luxembourg	Turkey	Cuba	India
Norway	Algeria	Dominica	Nepal
Netherlands	Egypt	El Salvador	Pakistan
United Kingdom	Morocco	Guatemala	Saudi Arabia
Germany	Tunisia	Honduras	Iraq
Sweden	Burkina Faso	Nicaragua	Iran
Switzerland	Benin	Panama	Israel
Cyprus	Cape Verde	Dominican Republic	Jordan
Greece	Cote d'Ivoire	Mexico	Lebanon
Ireland	Gambia	Argentina	Syria
Italy	Ghana	Bolivia	Ukraine
Portugal	Guinea	Brazil	Moldova
Andorra	Guinea Equatorial	Colombia	Belarus
Bulgaria	Liberia	Chile	Georgia
Hungary	Mali	Ecuador	Armenia
Poland	Mauritania	Paraguay	Russia
Romania	Nigeria	Peru	Kazakhstan
Malta	Senegal	Uruguay	Australia
Latvia	Sierra Leone	Venezuela	New Zealand
Estonia	Togo	Canada	
Lithuania	Angola	United States of America	
Czech Republic	Cameroon	China	

Table B.6: List of ports of entry

Mode	*Gate Name*	**Mode**	*Gate Name*
Raft	Las playitas en Fuerteventura	Boat	Puerto de Alicante
Raft	Punta de Tarifa	Boat	Puerto de Santa Cruz de Tenerife
Raft	Playa del Ingles Las Palmas	Boat	Puerto de Bilbao
Walk	Puesto fronterizo Ceuta	Boat	Puerto de A Coruña
Walk	Puesto fronterizo Melilla	Boat	Puerto de Malaga
Plane	Aeropuerto de Madrid-Barajas	Boat	Puerto de Valencia
Plane	Aeropuerto de Barcelona-El Prat	Boat	Puerto de Ceuta
Plane	Aeropuerto de Palma de Mallorca	Road	La Jonquera
Plane	Aeropuerto de Malaga	Road	Irun/Hendaya
Plane	Aeropuerto de Gran Canaria	Road	Roncesvalles
Plane	Aeropuerto de Tenerife Sur	Road	Candanchu
Plane	Aeropuerto de Alicante	Road	Puigcerda
Plane	Aeropuerto de Lanazarote	Road	Valenca, Portugal
Plane	Aeropuerto de Ibiza	Road	Badajoz
Plane	Aeropuerto de Fuerteventura	Road	Ayamonte
Plane	Aeropuerto de Menorca	Rail	Latour de Carol
Plane	Aeropuerto de Tenerife Norte	Rail	Portbou/Cerbere
Plane	Aeropuerto de Bilbao	Rail	Canfranc
Plane	Aeropuerto de Valencia-Manises	Rail	Hendaya
Plane	Aeropuerto de Sevilla	Rail	Vigo/Tui
Plane	Aeropuerto de Santiago de Compostela	Rail	Fuentes de Ouoro
Boat	Puerto de Barcelona	Rail	Valencia de Alcantara
Boat	Puerto Bahia de Algeciras	Rail	Badajoz
Boat	Puerto de Santander	Rail	Madrid Estacion de Atocha
Boat	Puerto de Almeria-Motril	Rail	Barcelona Estacion de Francia

Table B.7: Robustness checks: other population measures

	(1)	(2)	(3)	(4)	(5)	(6)
Population variable:	FBORN	WAFB	FOREIGN	WAFOR	POP	WAPOP
PANEL A: Native population rate (t-1)						
Rate (t-1)	0.308***	0.298***	0.302***	0.270***		
	(0.088)	(0.085)	(0.088)	(0.079)		
Weak identification test (KP)	22.94	23.75	22.28	22.85		
PANEL B: Change in log rental prices t\t-1						
Rate (t-1)	0.986**	0.955**	0.967**	0.866**	0.754**	0.669**
	(0.465)	(0.451)	(0.464)	(0.405)	(0.353)	(0.306)
Weak identification test (KP)	22.94	23.75	22.28	22.85	24.78	22.58
PANEL C: Change in log sale prices t\t-1						
Rate (t-1)	3.278**	3.174**	3.215**	2.876***	2.506***	2.224***
	(1.236)	(1.193)	(1.231)	(1.066)	(0.881)	(0.745)
Weak identification test (KP)	22.94	23.75	22.28	22.85	24.78	22.58
All province FE and controls	Yes	Yes	Yes	Yes	Yes	

Notes: Significance levels: * p<0.05, ** p<0.01, *** p<0.001. All specifications include province FE, province attributes*YearFE and time-varying controls, as described in Table 3. Clustered (province) standard errors in parenthesis. Clustered (province) standard errors in parenthesis. Obs=550

Table B.8: Robustness checks: other outcomes

	(1)	(2)	(3)	(4)	(5)
LHS variable:	RENTS	HP IVIE	HP AVER	HP 2Q	HSTOCK
PANEL A: Total effect					
Immigration rate (t-1)	0.986**	3.278**	3.784***	3.621***	0.996***
	(0.465)	(1.236)	(1.139)	(1.186)	(0.354)
Weak identification test (KP)			22.94		
PANEL B: Partial effect					
Population rate (t-1)	0.754**	2.506***	2.893***	2.769***	0.761***
	(0.353)	(0.881)	(0.802)	(0.841)	(0.258)
Weak identification test (KP)			24.78		
All province FE and controls	Yes	Yes	Yes	Yes	Yes

Notes: Significance levels: * p<0.05, ** p<0.01, *** p<0.001. All specifications include province FE, province attributes*YearFE and time-varying controls, as described in Table 3. Clustered (province) standard errors in parenthesis. Clustered (province) standard errors in parenthesis. Obs=550

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