Public Housing Magnets: Public Housing Supply and Immigrant Location Choices*

Gregory Verdugo†

December 7, 2011

Abstract

The supply of public housing and the participation rate of immigrants in public housing in Europe are significantly large. If immigrants are attracted by cities in which the public housing supply is substantial, their location choice could be distorted with respect to the theoretically most efficient location choice based on differences in labor market opportunities. This article investigates whether the initial location choices made by immigrants arriving in France between 1962 and 1999 were influenced by differences in public housing supply across urban areas. To deal with the potential endogeneity of changes in public housing supply over time, I use unexpected local population shocks unrelated to public housing and the fact that public housing construction was greater in cities heavily destroyed during the Second World War. The results indicate that increasing a city's public housing supply has a significant "magnetic effect" on new immigrant couples and that this effect is greater for non-European immigrants. I find no significant effect on single immigrants.

*I thank the Centre Maurice Halbwachs (CMH) for having made the data available. Most data used in this paper are available upon request for researchers from the CMH. The author also wishes to acknowledge the INSEE that provided the underlying data making this research possible. I thank Jean-François Royer from CREST-INSEE for his help to access the restricted version of the 1999 Census. I thank Karin Edmark, Thierry Magnac, Eric Maurin, Javier Ortega, Jean-Louis Pan Ké Shon, Gilles Saint-Paul, Karine Van Der Straeten, Etienne Wasmer, and seminar participants in many places for insightful comments. Data from the 1999 French Census used in this paper are confidential but the author access is not exclusive. For additional information, contact the author. This paper does not necessarily reflect the views of the Banque de France.

†Service des Analyses Microéconomiques, Direction des Etudes Microéconomiques et Structurelles, Banque de France, 31 rue Croix-des-petits-champs, 75049 Paris Cedex 01, France. Email: mailto:gregory.verdugo@banque-france.fr

Introduction

An important aspect of immigration is the settlement patterns of immigrants in the host country. Receiving countries have a greater capacity to absorb immigrants when the location choices respond more to changes in local economic conditions than to other factors. In regions with booming economies, immigrants are less likely to be unemployed and live on welfare. Borjas (2001) also emphasizes that immigration may help to equalize economic opportunities across cities if immigrants choose to live in places with a higher labor demand. The location choice of immigrants thus has important consequences for the impact of immigration and the economic outcomes of immigrants in the host country.

Empirically, Jaeger (2001) and Borjas (2001) find a significant response of immigrants to differences in wages and unemployment rates across locations.[1] However, the literature also emphasizes two other potential factors influencing the location choice of immigrants. The first factor is the existence of ethnic enclaves attracting immigrants from the same origins to places where their own ethnic group is concentrated (Bartel, 1989). Ethnic enclaves minimize the psychological cost of migration, and ethnic networks may facilitate the economic integration of immigrants. A second factor influencing the location choice of immigrants also found in the literature is the effect of differences in welfare availability across locations. It has been noted by Borjas (1999) among others that, unlike natives, immigrants may be disproportionately attracted to places with more generous welfare policies because, at least initially, they do not have local friendship ties and community attachments. As a result, several recent studies on the location choice of immigrants have investigated whether welfare has a "magnetic" effect on the location choice of immigrants.[2]

However, as far as I know, most studies investigating the impact of differences in welfare availability on the location choice of immigrants have focused on welfare programs providing

[1] The literature investigating the location choice of immigrants also includes, among others: Bauer et al. (2005) for the U.S., Pischke and Velling (1997) for Germany, Desplanques and Tabard (1991) and Jayet and Ukrayinchuk (2007) for France.

[2] For the U.S., Borjas (1999) found evidences of interactions between welfare and the location choice of immigrants. Applying a change in the welfare eligibility of immigrants in the US, Kaushal (2005) reported no impact. Giorgi and Pellizzari (2009) also studied the impact of differences in welfare benefits across European states on immigrant location choice and reported a positive effect. Boeri (2010) discusses the impact of welfare on immigrants in Europe.

monetary benefits. This paper examines the impact on immigrants of a welfare program that is still popular in Europe and is potentially highly persistent and for which the participation rate of some groups of immigrants is much higher than that of natives: public housing.

There are at least two good reasons to investigate the influence of public housing on the location choice of immigrants. First, the fact that public housing is location-specific may introduce far more spatial variations in the benefits to settling across different areas than welfare programs providing monetary benefits, which are more easily changed in the short run.[3] As a result, immigrants, particularly those with a large family, may be significantly responsive to regional differences in public housing supply, particularly if the benefits of living in public housing for this group are substantial. Second, public housing remains a major policy issue in Europe. In 1996, public housing accounted for more than 40% of the housing stock in the Netherlands, 20% of the total housing stock in Austria, the United Kingdom, Sweden, and Denmark and more than 10% in Germany, Ireland, France, and Belgium (Priemus and Dieleman, 2002).[4] Investigating the impact of public housing on the location choice of immigrants is thus interesting for understanding several potentially unintended consequences of public housing policies that may adversely affect immigrants in their host country. Understanding whether public housing affects the geographical dispersion of immigrants in Europe is also informative with respect to determining some of the mechanisms that drive the economic integration of immigrants throughout most of Europe. In practice, several figures suggest that public housing should be a particular concern for immigration policymakers, considering that the participation rates of some groups of immigrants are much higher than the rate of natives. In Amsterdam, for example, it is estimated that more than 80% of the Turkish and Moroccan immigrants lived in public housing in 1990 (Musterd and Deurloo, 1997), while in London, 40% of the foreign-born residents are social tenants (Rutter and Latorre, 2008).

Using French data on the location of immigrants arriving between 1962 and 1999, this paper investigates whether differences across urban areas in the supply of public housing affected

[3]For example, in 1996, in the US, the Personal Responsibility and Work Opportunity Reconciliation Act denied legal noncitizens who arrived in the US access to means-tested federal benefits for the first 5 years of residence (Kaushal, 2005).

[4]On the other hand, in the US, the share of public housing declined consistently during the 1990s and represented only 3.3% of the occupied rental housing units in 2000 (Baum-Snow and Marion, 2009).

the initial location choice of new immigrants. France is a good case study for addressing these questions because, as previously stated, a remarkably large share of its population lives in public housing, and French data provide extremely precise information on public housing supply and participation. As in other European countries, participation rates in public housing are also much larger for immigrants than for natives. For example, in 1999, 16% of the natives but approximately 31% and 50% of immigrants from Algeria and Morocco, respectively, lived in public housing.[5] Finally, spatial variations in the public housing supply across localities are also significantly large. In 1999, the share of public housing units over total housing varied from 7% in the urban unit of Nice to 44% in the urban unit of Reims, with an average share of 20% across localities.

However, several factors complicate this evaluation. The main econometric issue is that unobserved city-specific factors may affect both the public housing supply and the utility of choosing the city for an immigrant. To deal with the potential correlation between cities unobservable factors, which are constant over time, and the public housing supply, I estimate discrete location choice models that control for time-invariant city characteristics as in Jaeger (2008). The model is thus identified with variations within cities over time in the supply of public housing. However, changes over time in several unobserved factors influencing the location choice may also be correlated with changes in the public housing supply. We account for this issue by comparing the location choice of male new immigrants with children to those of male immigrants who are single. Because households with a large family are more likely to be eligible for public housing, immigrants who are single could be used as a control group if the two groups react similarly to other observed and unobserved factors influencing the location choice. Differencing between groups will rid our estimates of some bias related with the correlation between changes in public housing supply and changes in a city unobserved factor influencing the location choice.

I confront the endogeneity of changes in the public housing supply within cities with an instrumental variable strategy inspired by the literature on the consequences of the durability of housing (Glaeser and Gyourko, 2005). Public housing is durable and cannot therefore

[5]Unless indicated otherwise, the figures on France reported in the introduction are from the author's tabulations of the French census.

be rapidly adjusted upward or downward when the population decreases. As a result, if public housing supply is not perfectly elastic with respect to population changes, population variations unrelated to public housing will affect the supply per household. As a first excluded instrument for the variations in public housing supply, I take advantage of plausibly exogenous population shocks across municipalities by constructing two "shift-share" variables (see Bartik (1991), Saiz (2010) and Saks (2008) for recent use of such instruments) to predict a counterfactual evolution of the public housing supply per household using differences in industrial composition across municipalities to predict population changes. I also construct a second instrument based on the fact that initial public housing constructions were larger in those cities that where most heavily destroyed during the Second World War. Differences in the intensity of destruction caused by the war are, in practice, a strong predictor for the cross-section dispersion of the share of public housing per household in 1990. Given that we use a differencing approach between singles and immigrants with a large family and given that our model is identified with changes in settlement patterns over time, our requirement for the exogeneity of these instruments relies on the hypothesis that they (the instruments) may be unrelated to changes over time in unobserved determinants of settlement patterns that affect new immigrants with a large family differently than singles.

To further probe the possibility of reverse causality, I estimate models that include lagged and future public housing supply. Lagged stocks cannot respond to future immigrant flows while the inclusion of the lead public housing supply tests for endogenous shifts in supply responding to immigrant flows. Reassuringly, I found small and statistically insignificant effects of future variables.

Our findings show that European and non-European immigrants react quite differently to the availability of public housing. Quantitatively, the estimates controlling for location fixed effects indicate that for non-European new immigrant couples with children, a one-standard-deviation increase (approximately 5%) in the number of public housing units per household in the city increases the probability of an immigrant choosing a city with "average" characteristics by approximately 25%. This effect is fairly robust across alternative specifications, such as changes in the choice set. On the other hand, I find statistically insignificant effects of public housing supply on European immigrants.

The paper proceeds as follows. Section I discusses the theoretical framework. Section II describes the data and some relevant institutional details about public housing and immigration in France. Section III presents the empirical model. Section IV discusses the main estimation results. Section V concludes the paper.

1 Theoretical Framework

We present a brief discussion of the theoretical framework to set the stage for the empirical work. We aim to highlight the conditions under which we may interpret the location decisions as influenced by differences in public housing supply. To do so, we consider a simple version of the Rosen-Roback framework (Moretti, 2011). For simplicity, we neglect the effect of differences in amenities across locations. We begin by assuming that workers are risk neutral, and the indirect utility of a worker in city i is given by the differences between wages and housing costs: $U_i = w_i - c_i$. However, unlike the original model, where all workers are identical, we assume individuals differ with respect to their family size, that is, there are families (denoted f) and singles (denoted s). Wages are similar for both types of individuals within cities. There are two types of housing available in each city, that is, private housing and public housing. Private housing is accessible to families and singles at the same price, but public housing is only available to families.[6] Housing costs differ between private and public housing in city i and are denoted, respectively, by r_i and r_{si}. To simplify the discussion, we assume that local public housing authorities adjust rents in public housing such that the differences in rents in public and private housing are similar across cities and equal to k: $r_i - r_{si} = k$.[7] Apartments in public housing in city i are obtained with probability p_i, which depends on the supply of public housing in the city. The fact that housing costs are lower for public housing implies that individuals with a large family always prefer to live in public housing. Expected housing costs in city i for a family in private housing at the beginning of the period are thus given by

[6]The assumption that there are two types of housing is also made in a different context by Beaudry et al. (2010) to explain the differential location choice of skilled and unskilled.

[7]This assumption is quite reasonable given the existing evidences that public housing authorities partially adjust rents in public housing with respect to differences in living cost (Le Blanc et al., 1999).

$(1-p_i)r_i + p_i r_{si} = r_i - kp_i$ and are decreasing in p_i.

What does the assumption that there are two types of housing with different prices imply about the equilibrium sorting between families and singles? A first implication of the model is that because individuals differ in their family size and current housing status, the relative utility across locations differs across individuals. Suppose there are only two cities indexed by 1 and 2 and consider the decision of inhabitants in city 1 to move to city 2. We assume workers take wages and house prices for each locality as given when deciding where to locate, and we denote ΔU_2^l the difference in utility in city 2 with respect to city 1 for an individual of type l and $\Delta U_2^{f,pv}$ and $\Delta U_2^{f,pu}$ the relative utility of moving for families initially living in private and public housing. Individuals decide to move $\Delta U_2^l = U_2^l - U_1^l > 0$. Suppose $p_2 > p_1$, that is, city 2 provides a higher supply of public housing and $\Delta w_2 - \Delta r_2 < 0$, thus economic outcomes in city 2 net of private housing costs provide a lower utility. In this case, we have $\Delta U_2^{f,pv} > \Delta U_2^s > \Delta U_2^{f,pu}$.[8] In this framework, individuals in city 1 with a large family who are not already living in public housing are more likely to respond to differences in the probability of obtaining an apartment in public housing. If there are no moving costs, the model predicts that public housing will change the composition of the city by attracting individuals who are eligible for public housing. The model also predicts that families living in public housing are less likely to move because they may lose the benefits of public housing when moving, which is consistent with the empirical evidences indicating that public housing inhabitants are less mobile (Gobillon, 2001).

Following Borjas (2001), assume that there are fixed costs to moving. Fixed costs may be explained by the existence of family or friendship ties, for example, which may deter a native's internal migration as emphasized by Mincer (1978). This implies that utility will not be equalized across locations in equilibrium and that there will exist wage differentials exclusive of housing costs across cities. For sufficiently high migration costs, all internal migration may be deterred. Consider the case of newly arrived immigrants. Unlike natives, they have already paid the fixed costs of migration and thus should be more responsive to differences in welfare benefits across localities (Borjas, 1999). Immigrants with a family need only to have

[8]We have $\Delta U_2^s = \Delta w_2 - \Delta r_2$ for singles, $\Delta U_2^{f,pv} = \Delta w_2 - \Delta r_2 + (p_2 - p_1)k$ for individuals with a large family in private housing in city 1, and $\Delta U_2^{f,pu} = \Delta w_2 - \Delta r_2 - (1-p_2)k$ for families in public housing in city 2.

$\Delta U_2^{f,pv} > 0$ to select city 2 instead of city 1. Income maximization hypothesis combined with the assumption that there are high fixed costs of migration and two types of housing generate strong predictions. The main insight of the model is that new immigrants with a large family should be more responsive to differences in public housing availability across locations. As a result, if public housing influences the location choice of families, we should observe that the location choice of families and singles differ. The fact that immigrants with a large family have access to public housing while immigrants who are single are not eligible provides a natural control group to investigate the effect of public housing.

From the recent literature on immigrant segregation and on the existing empirical evidences, there are additional reasons to expect that public housing may have a different effect on the location choice of immigrants with respect to natives. Recent work by Bouvard et al. (2009) provides several evidences that immigrants, particularly non-European immigrants, are discriminated against in the private housing market in France. On the other hand, other studies have suggested that immigrants cannot be discriminated against in public housing as public housing allocation follows a strictly delimited administrative process (Algan et al., 2011). A differential effect of public housing between immigrants and natives might thus also rely on the existence of discrimination in the private sector housing market. If the discrimination of immigrants in the private housing market implies housing costs in the private market will be higher for this group, differences in public housing supply will be even more attractive for immigrants than it is for natives.

These strong theoretical implications follow from a framework using very restrictive assumptions. A first issue is that the fertility decisions may be influenced by the public housing supply. Individuals who are single with higher preferences for children may also react in anticipation to differences in public housing if they plan to have children in the future. Similarly, fertility decisions may also depend on the public housing supply in the city. That is, immigrants may first choose their location decision according to factors unrelated to public housing and then decide to have children based on the supply of public housing.[9] A second issue is that

[9]I am not aware of a paper that explores whether public housing availability influences fertility. However, recent research by Dettling and Kearney (2011) has found that an increase in housing price decreases fertility for households who are not homeowners.

public housing supply is considered as exogenous in the model. Public housing projects may be constructed as a response to immigrant inflows in a given city that were related with other factors. Finally, the assumption that individuals who are single are not eligible to public housing is a simplification. While it is true that most of the existing public housing units are large apartments designed for families, there are also some public housing units available for singles. However, as emphasized in Table 3 below, the participation rate of new immigrants with a large family is three times higher than that of singles; therefore, the assumption that public housing is more attractive to families seems reasonable. In the empirical analysis, I attempt to take these issues into consideration.

2 Institutional Background

We first present the data used in the paper in a first subsection. We present the characteristics of public housing and discuss the historical context of public housing constructions in France in a second and third subsection. A fourth subsection documents the evolution of the participation rates of immigrants in public housing.

2.1 The Data

The empirical analysis draws data from the 1962, 1968, 1975, 1982, 1990 and, the 1999 censuses. The sampling rate for the individual file is 5% for the 1962 Census, 20% for the 1975 Census and 25% for the 1968, 1982, 1990, and 1999. Such high sampling rates guarantees I can study small subpopulations of immigrants separately without too much sampling errors (Aydemir and Borjas, 2011). To approximate the relevant local labor market from which the characteristics determine the location choice, I use 57 urban areas with more than 100 000 inhabitants in 1990. Urban areas aggregates municipalities between which there are no discontinuities across construction. These urban areas were chosen by more than 85% of immigrants in 1968 and 1990.[10] The boundaries of urban areas is maintained constant across censuses by

[10] If one excludes European immigrants, the percentage increases to 92% in 1968 and 90% in 1990.

matching municipalities with the national municipality code.[11]

I restrict the sample to men and women aged 16 to 60 and exclude students and individuals in the military.[12] As usual, an immigrant is defined as a foreign-born individual who is a non-citizen or naturalized French citizen. To identify newly-arrived immigrants, I use the reported location of an individual at the time of the previous census.[13] A "new immigrant" is therefore defined as an immigrant who declared to be living abroad at the time of previous census.

The main characteristics of immigrant inflows to France between 1962 and 1999 are reported in panel A of Table 1.[14] With respect to the period before 1975, annual immigrant inflows during the 1990s have been divided by two; even so, however, they remained large. The most recent period is also characterized by an increase in the share of family-based migration. Row 3 in panel A of Table 1 indicates that the share of male new immigrants declined from 60% before 1975 to 46% during the 1990s. Columns 3 and 4 of Table 2 document the distribution of, respectively, immigrants and new immigrants in France across the largest urban areas. As in many countries, immigrants tend to be spatially concentrated. For example, traditional immigrant cities such as Paris, Lyon or Marseille have a much larger share of immigrants than other cities. The figures also indicate that new immigrants tend to locate disproportionately in Paris. While 25% of the population in Paris are natives, 50% of the new immigrants are located there.

In this paper, the term public housing refers to what the French call social housing (*logement social*) or 'HLM' (*Habitation à Loyer Moderé*, which means, literally, housing with moderated rents). Public housing thus designates constructions financed by the government and managed by local public housing authorities that are rented. Information on whether a dwelling is in the public-rented sector is available from the Census of Housing and thus does not come from current resident reports. To estimate the public housing supply per urban area across years, I use the exhaustive dwelling files (100% extract) from the 1982, 1990 and 1999 Censuses, which

[11] Each municipality has had a unique administrative identifier constant over time. Boundaries of municipalities are also stable over time with very few exceptions which have been accounted for.

[12] However, the population count used to select urban areas included in the analysis includes all individuals.

[13] Unlike US Census data, there is no variable indicating the arrival year for each foreign-born individual until the 1999 Census.

[14] Reassuringly, these estimates are very similar to those based on administrative data from the National Immigration Office (Tavan et al., 2005, p.70).

Table 1: New Immigrants and Public Housing Supply in France 1968-1999

A. New Immigrants in France 1968-1999					
Arrival Period	1962-68	1968-75	1975-82	1982-90	1990-99
Total Number (in thousands)	915	1 053	707	663	689
Number per year	152	150	101	95	77
Share of new immigrants over total immigrant stock	28.3%	27.1	17.5	15.9	16.0
Proportion of Male	60.2%	59.4	50.6	49.9	46.8

B. Estimated Changes in Public Housing 1968-1999					
Year	PH Units Stock (in thousands)	Nb per year	Pct Change	PH per Capita	Std.
1945	275				
1968	1 395		400%	7.4%	2.9
1975	2 239	121	60	11.1	4.1
1982	2 725	69	22	13.6	4.8
1990	3 093	46	14	15.7	7.3
1999	3 454	40	12	17.1	5

Notes for panel A: New immigrants are immigrants who declared to have lived abroad during the previous census year. Author's tabulations from 1968, 1975, 1982 and 1999 Censuses. *Notes for panel B*: Only primary residences in urban areas with more than 10 000 inhabitants in 1990 are included in the calculations. *Pub. Housing per Capita* and *Std.* columns reports respectively the average and standard deviation of the public housing supply per capita across the 57 cities with more than 100 000 inhabitants in 1990. The public housing unit stock is estimated retrospectively using building construction dates from the 1999 Census of Housing. *Sources*: Author's tabulations from 1999 Census of Housing and the 1968, 1975, 1982, 1990 and 1999 Censuses of Population.

provide information about all dwellings and buildings existing in France during those years. However, the 1968 and 1975 censuses did not collect information on public housing. For those years, I retrospectively approximate the number of public housing units per urban area using the indicated construction year of each building with the 1982 census of dwelling. Because there were no destructions of public housing units between 1968 and1982, this method is a relatively accurate approximation for those years.

2.2 Institutional Framework

Rents in public housing are much lower than in the private sector with existing estimates indicating that rents were an average of 40% lower than in the private sector during the 1990s

Table 2: Major Urban Area Characteristics in 1990

Urban area	Total Population	Public Housing per Capita	Immigrants to Population	Share of New Immigrants	Share of Natives
Paris	9 316	22.1%	19.3%	51.8%	25.5%
Lyon	1 262	20.1	14.7	3.5	3.5
Aix-Marseille	1 230	15.8	11.6	2.5	3.4
Lille	959	24.6	9.8	1.7	2.6
Bordeaux	696	16.4	7.6	1.4	2
Toulouse	649	14.4	10.1	1.6	1.9
Nice	517	7.8	13.8	1.6	1.4
Nantes	495	19.8	3.8	0.6	1.5
Toulon	437	10.7	8.9	0.5	1.2
Grenoble	404	16.2	15.8	1.1	1.1
Strasbourg	387	19.8	14.4	1.6	1.1
Rouen	380	30.9	6.7	0.6	1.1
Valenciennes	338	18.5	6.5	0.3	0.9
Antibes	335	7.1	15.5	1.3	0.9
Nancy	329	21.4	7.7	0.6	0.9
Lens	323	19	4.2	0.1	0.9

Notes: Column (1) reports the total population of the urban area including all individuals in thousands. Column (2) reports the proportion of public housing units per capita. Only primary residence and inhabited housing are included in the calculations. Population taken into account in the calculations of the other columns is restricted to men and women between 16 and 60 not in school and not in the military. *Sources*: 1990 Census. Author's calculation.

(Le Blanc et al., 1999) and approximately 30% lower during the 1970s (Durif and Marchand, 1975). Public housing management is decentralized at the municipal level and eligible families can apply in any municipality, regardless of their current location or nationality, and eligibility rules are uniform throughout France. The only requirements to be eligible for public housing are to be legally living in France (as a French citizen or a migrant with a valid residence permit) and to be living under a certain threshold of income per unit of consumption, which is usually rather high. Eligibility depends on the income per unit of consumption and family size, which must be below a threshold that varies across regions (Algan et al., 2011, see for details). As a result, the number of eligible families is approximately three times greater than the available space in public housing. Consequently, average waiting periods can be quite long, For example, in the Paris region, recent figures indicate that average waiting times may be in excess of five years (Guillouet and Pauquet, 2011).

In practice, the availability of public housing in a given urban area appears to be proportional to the supply of public housing per households at the local level. Figure 1 presents the relationship between the log of average waiting times across French regions for new public housing inhabitants living in urban areas with more than 100 000 inhabitants and the log of public housing per households estimated using the 1996 and the 2002 housing condition surveys.[15] The figure indicates a clear negative relationship between average waiting times and the public housing supply across regions, while the region Ile-de-France, which designates the Paris regions, stands out as an obvious outlier.[16] Overall, this suggests that public housing supply per capita might offer a reasonable proxy for the availability of public housing at the urban area level. As detailed figures on waiting times for public housing across localities are not available, even for the recent period, I use the number of public housing units per households that can be calculated across urban areas from 1968 to 1999 to capture differences in public housing availability across locations.

[15] For reasons of confidentiality, these surveys do not contain information on the municipality or agglomeration of residences. Waiting times are not available for earlier housing condition surveys. Four sparsely populated regions with fewer than 15 individual observations have been aggregated with the neighboring region. The small sample size of 2,490 observations does not enable me to compute specific waiting times for immigrants.

[16] A regression of the log waiting time on the log of public housing per capita, excluding the Paris region, provides a parameter (standard error) of -0.32 (0.13).

Historically, the public housing program in France started much later than it did in the US[17], after the Second World War. War destruction created severe housing shortages and, as a result, the government implemented a system of rent control in 1948. As an unintended effect, this law reduced new constructions drastically in spite of a high demand. In 1958, during a period of rapid economic growth, the freshly elected center-right Gaullist government launched a massive construction plan for public housing intended to both increase the supply of housing and improve its poor quality. Panel B of Table 1 documents the estimated evolution of the public housing unit stocks from 1945 to 1999 using the stock of public housing units in 1999. The figures indicate that the share of public housing units constructed before 1945 is negligible. However, more than half of the public housing stock in 1999 was constructed between 1945 and 1975. Construction rates were particularly large until 1975. In fact, between 1968 and 1975, the number of public housing units increased annually by 121 000, which increased the public housing per capita from 7.4% to 11.1%. Construction rates decelerated during subsequent decades, but they remained, nonetheless, relatively large, increasing the public housing per capita from 11% to in 1975 to 17% in 1999.

Public housing constructions were undertaken in cooperation between the central government and local authorities (mainly municipalities) through local public housing agencies.[18] Politically independent municipalities could, in practice, veto any construction projects, such that it was impossible for other layers of the local or national government to construct without their agreement.[19] Over the period, public housing constructions were durable. That is, up until 1999 there has been no destruction of public housing units. Moreover, there were no policies for converting public housing unit apartments into condominiums as in the US (Stébé, 2007).

As a result of the decentralized mechanisms of construction, there are large spatial disparities in the public housing supply across urban areas in France. Table 2 documents the

[17] In the US, public housing emerged with the Housing Act of 1937 (Hunt, 2009)

[18] There exist approximately 820 different public housing agencies (‘*organismes HLM*’) in France. The boards of these organizations are typically composed of local politicians from different levels of the French local and national administration.

[19] The vote in 1999 by the socialist government for a law called *Solidarity and Urban Renewal* (SRU) (see, e.g., Selod (2004) for details) illustrates the central role of municipalities. This law was endorsed by the socialist government to homogenize the supply of public housing across municipalities. To do so, each municipality, under the law, is mandated to have at least 20% of their inhabited construction in the public housing domain. Municipalities not reaching this threshold have to pay severe fines, which are proportional to the gap between their actual supply and the 20% threshold.

heterogeneity in the public housing supply across the 16 largest urban areas in 1990.[20] The figures indicate that the level of public housing per household varies dramatically across urban areas. For example, the urban area of Antibes had a very low supply in 1990 with 12% public housing per capita, while the largest supply of public housing is found in Le Havre at 48%. The last column in panel B of Table 1 indicates the standard deviations of public housing per capita across 57 urban units of more than 100 000 inhabitants in 1990, indicating the level of dispersion of the public housing supply across urban areas. The dispersion in supply across urban areas is significant, increasing from 3 percentage points in 1968 to 7 percentage point in 1990, while it subsequently decreased to 5 percentage points in 1999.

The spatial disparities in public housing supply can be explained by several factors. First, part of the initial stock of public housing is related to destruction caused by war. I indicate below that public housing constructions before 1968 were much greater in cities that had been bombed and partially destroyed during the Second World War such as Troyes, Reims and le Havre (Florentin, 1997). War destructions provided space within central cities to construct large public housing projects that were more easily funded by the government given the large housing shortages in cities where war destruction had been significant (Voldman, 1997). Second, a large share of public housing has been constructed in places with higher demands for housing during the 1960s. Given that public housing is durable, population changes over the period and differences in initial public housing stock explain a large share of the variations in public housing supply (Verdugo, Forthcoming). Obviously, variations in local attitudes to public housing may also have been relevant (Subra, 2006). While left-wing politicians are usually more in favor of public housing, the effect of differences in political partisanship on the evolution of the supply appears negligible during this period. Verdugo (Forthcoming) finds no effect of differences in political preferences across municipalities to explain changes in supply between 1968 and 1999. Finally, several pieces of evidence suggest that the initial public housing share in 1968 is not related with immigration. During the 1960s, immigrants were discriminated against with respect to public housing, and their access was generalized only during the 1970s. The fact that initial public housing stocks were not directly related to initial immigration

[20] Arbitrarily, the main municipality of the urban area is defined as the most populated municipality of the area.

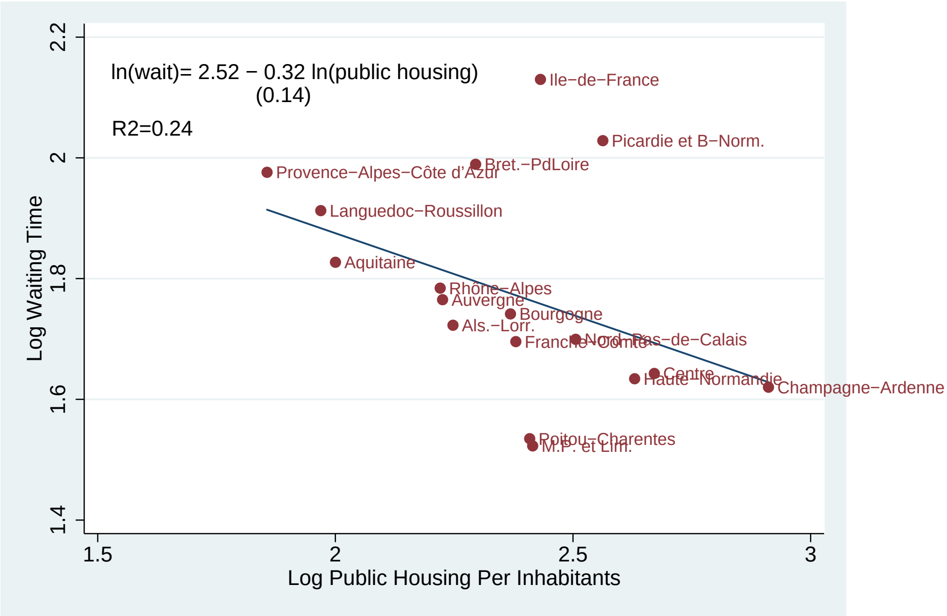

Figure 1: Waiting Times and Public Housing Supply across Regions

Notes: the Figure plots the relationship between log public housing per inhabitants and log average waiting time for public housing reported by new inhabitants across French regions. *Sources:* Public Housing per Inhabitants from the 1999 Census. Waiting times for public housing the 1996 and the 2002 Housing Condition surveys.

patterns rules out a direct relationship between the initial settlement patterns of immigrants and a part of the future public housing stock. I provide empirical evidences on these issues in the rest of the paper.

2.3 Public Housing and Immigration

The relationship between public housing and immigration changed dramatically between the 1960s and the 1970s because public housing was not initially easily accessible to immigrants during the 1960s as the government wanted to discourage family-based migrations and therefore provided incentives of return migration to immigrants during periods of economic downturn (Weil, 2005). In practice, strict rules were limiting the eligibility of immigrants in public housing as public housing agencies required immigrants to first maintain residency for 10 years and to have children (Schor, 1996, p.214). The number of immigrants in public housing was also limited by quotas. In some regions, no more than 6.5% of the units in housing projects could be occupied by immigrants. As a result, the participation rate of immigrants was much lower than that of natives: Pinçon (1976) reports that in 1968, only 5.5% of the foreign workers in the urban area of Paris lived in public housing versus 15.3% of the natives in the same area lived in public housing. However, these discriminatory policies combined with large immigrant inflows during the 1960s resulted in many immigrants living in slums around the French cities. In 1970, Lequin (2006, p.410) finds 113 slums in the Paris region alone. In response to increasing pressure from public opinion to eliminate immigrant slums, the policy changed drastically at the beginning of the 1970s when a new government was elected and decided to end immigrant discrimination regarding access to public housing. Finally, after 1981, the new socialist government passed a law that explicitly forbade discrimination of immigrants in public housing.

As a result of the increased availability of public housing to immigrants, available evidences indicate that in recent years, non-European immigrants settled disproportionately in public housing.[21] Table 3 documents the participation rates of immigrants between 1982 to 1999.[22] From 1982 to 1999, the share of immigrants participating in public housing increased by 10 to15 percentage points for immigrants from Africa and Maghreb, while this figure did not change much for natives or European immigrants during the same period. As a result, in 1999, approximately half of immigrants from Maghreb lived in public housing, a difference of 34 percentage points compared to the natives of France.

Finally, it is important to note that the participation rate in public housing varies widely with respect to family status. Most public housing units are large dwellings specifically designed for families, and as a result, the participation rates of households with only one person are quite low. Panel B of Table 3 reports the participation rates in public housing of new immigrant couples with children and of single immigrants. From 1982 to 1999, the participation rate of couples was approximately three times greater than the participation rate of singles.

[21] See also Boeldieu and Thave (2000), Fougère et al. (2011) and Verdugo (2011).

[22] Unfortunately, I am unable to quantify precisely the evolution of public housing participation of immigrants during the 1960s and the 1970s. The censuses before 1982 did not contain information on public housing, and as far as I know, there are no alternative sources available to study the participation rates of immigrants before 1982. The housing conditions surveys (*Enquêtes Logement*) of 1973 and 1978 collected by the French Statistical Institute gathered information on public housing participation, but they do not contain information on nationality.

Table 3: Participation rates in Public Housing

A. Participation rates per Nationality			
	1982	1990	1999
Natives	13.6%	14.0%	15.7%
Immigrants	22.9	25.8	30.6
New Immigrants	27.6	22.2	24.6
Percentage of Immigrants in Public Housing from			
Europe	16.0%	15.8%	16.3%
Africa	33.1	39.1	46.4
Algeria	34.8	42.5	49.7
Morocco	37.3	43.1	48.3
Tunisia	27.6	43.1	39.2
B. Participation Rates of Male New Immigrants per Family Status			
Couples with Children	32.9	28.6	43.6
Singles	7.4	9.8	16.3

Sources: 1999, 1990 and 1982 Censuses. Author's tabulations.

3 Estimation Framework

The empirical model is estimated using the location choice of new male immigrants who arrived in France between 1962 and 1999. Empirically, I assume that immigrants have an additive stochastic utility function.[23] By convention, t is the current census year and $t-1$ denotes the previous census year. The baseline econometric model for the location choice of a new immigrant i from ethnic group l arriving between period $t-1$ and t, and thus observed in France in the census of year t, is:

$$U_{ilkt} = Z_{lk,t-1}\theta_1 + X_{k,t-1}\theta_2 + \delta p_{k,t-1}(1-e_{i68}) + \Gamma_{68}p_{k,68}e_{i68} + \theta_k + \epsilon_{ilkt} \quad (1)$$

where U_{ilkt} is the level of utility provided by location k. The unobserved component of utility ϵ_{ikt} captures unobserved factors affecting utility. The parameter of interest is δ, which captures the effect of public housing per household $p_{k,t-1}$ on the utility of immigrants. The public housing magnet hypothesis is true if δ is positive and significantly different from zero. To rule out the possibility that the potential correlation between public housing and location decision comes from new housing that was built in response to immigrant flows, I estimate the model

[23] This follows Bartel (1989), Kaushal (2005), Bauer et al. (2005), Jaeger (2008) and Giorgi and Pellizzari (2009).

using the lagged share of public housing evaluated from the previous census in $t-1$ for a new immigrant observed in t. Using the lag values of public housing also reflects the possibility of a time lag between the construction of new public housing and the arrival of immigrants to the city. The variable e_{i68} is an indicator function equal to one if the individual is observed in the 1968 census and thus arrived between 1962 and 1968. Because historical evidences discussed above indicate that immigrants were discriminated against in public housing during the 1960s, the effect of differences in public housing supply is allowed to differ for immigrants who arrived before 1968. Data constraints also imply that for arrivals between 1962 and 1968, I am not able to use the lagged share of public housing in 1962 for this group; therefore, I use, instead, the share in 1968.[24]

To estimate (1), I follow the current approach used in literature regarding immigrant location choices and assume that ϵ_{ilkt} is an independent and identically distributed extreme value.[25] One characteristic of the conditional logit model is that the relative odds of choosing two alternatives are independent from the availability or attributes of other alternatives, a property known as the independence from irrelevant alternatives, or IIA. We will check the sensitivity of the results to this assumption.

Other covariates are also introduced in the model using their value at the time of the previous census, that is, before the arrival of new immigrants in the city. The vector $X_{k,t-1}$ contains a set of control variables that vary at the city level and that reflect changes in the underlying characteristics of the city, while variables in Z_{lkt} capture the effect of community characteristics. These variables are described below. All predictors in X and Z and the public housing supply in $p_{k,t}$ have been standardized to have an average of zero and a standard deviation of one across the choice set of individuals. This normalization is equivalent to using the relative dispersion of these variables within ethnic groups and years to estimate the model rather than using the absolute value.[26] Because of the logistic form of the model, the standardization also

[24] Using the contemporary public housing share should bias the coefficient upward if public housing constructions responded to immigrant inflows. However, in most specifications reported, I find a *negative* coefficient of differences in public housing supply in 1968.

[25] The conditional logit model is used in Bartel (1989), Kaushal (2005), Bauer et al. (2005), Jaeger (2008) and Giorgi and Pellizzari (2009).

[26] For example, the average percentage of the urban area population for immigrants from Algeria is 1%, whereas it is 0.01% for immigrants from Cameroon. Because the size of these two groups is different, normalizing is similar to assuming that a percentage increase of 1% of this variable should have a different effect on immigrants from

simplifies the quantitative interpretation of the parameter estimates. Denote P_0 as the predicted probability of the average urban area and P_k as the predicted probability of the average urban area in which the variable k is higher by one standard deviation. In the appendix, I show that the coefficient of a conditional logit in which the predictors have been standardized is equal to the log difference between these two probabilities, $\log P_k - \log P_0 = \gamma_k$, where $\gamma_k = \beta_k \sigma_{x^k}$ and σ_{x^k} are the standard deviations of the variable k with respect to initial alternatives included in the choice set.

3.1 Endogeneity Issues

The main challenge in estimating the model of Eq. 1 is that differences in the public housing supply within cities over time may be related with unobserved characteristics of the urban area that also influence the location choice of immigrants. I attempt to resolve the issue of causality by two means. First, following Jaeger (2008), the model in Eq. 1 includes urban area fixed effects θ_k that absorb the effect of unobservable or omitted urban area characteristics that are constant over time. Identification in this case relies from the within-urban area variations of the covariates over time. City characteristics such as average temperature and other various observed or unobserved amenities that are invariants over time are absorbed by the fixed effects included in the model.

Controlling for unobserved factors at the urban area level is necessary but not sufficient for causality. We follow Cascio and Lewis (2011) by using a comparison group that should react similarly to changes in other factors but which should not react similarly to changes in public housing supply to account for the effect of unobserved factors changing over time influencing the settlement patterns of both groups. We use the group of immigrants who are single and live without children as a control group and who, as emphasized before, are less likely to be eligible for public housing and therefore should not react in the same way to differences in the supply of public housing. If changes in unobserved factors correlated with changes in public housing supply also influence new immigrants who are single, the effect of public housing supply could be estimated using changes in the relative probability between singles and families to choose a

Cameroon than on immigrants from Algeria.

location with respect to changes in the public housing supply of that location. One condition for this strategy to hold is that the location choice of the control group should, on average, react on the same way to changes in other urban area attributes.

Such empirical strategy is difficult to implement directly with a conditional logit model, thus we use an alternative method executed in two steps. First, we estimate separate conditional logit models for each year t for the two groups of singles and families to estimate the probability of choosing a given urban area after the effect of the variables specific to each ethnic group $Z_{ik,t-1}$ have been taken into account. Specifically, we estimate separately using new immigrants who arrived in census year t with a conditional logit the following model:

$$U_{ilkt} = Z_{lk,t-1}\theta_1 + \gamma_{kt} + \epsilon_{ilkt} \tag{2}$$

where γ_{kt} are specific urban areas by year fixed effects. Because covariates included in Z are standardized, the parameter γ_{kt} can be interpreted as the log of the odd ratio of choosing urban area k with respect to the reference urban area. We estimate two sets of parameters using separately individuals living in couples with children in the sample ($\hat{\gamma}^c_{kt}$) and individuals who are single ($\hat{\gamma}^s_{kt}$). The difference between the two groups probability of choosing an urban area becomes the dependant variable in the second step. The second step regression is the following first difference estimate:

$$\Delta\hat{\gamma}^c_{kt} - \Delta\hat{\gamma}^s_{kt} = \Delta X_{k,t-1}\tilde{\theta} + \tilde{\delta}\Delta p_{k,t-1} + u_{kt} \tag{3}$$

where $\Delta\hat{\gamma}^l_{kt} = \hat{\gamma}^l_{kt} - \hat{\gamma}^l_{kt-1}$ for $l = c, s$. The dependent variable $\Delta\hat{\gamma}^c_{kt} - \Delta\hat{\gamma}^s_{kt}$ represents the change over time in the probability to locate in urban area k for immigrant couples with children (group c) relative to the comparison group of singles (group s) between t and $t-1$.[27] The second-step regression is weighted by the precision with which we estimate the two sets

[27]Denoted by P_{kt} and P_{0t}, the probability of an individual with average characteristics Z_{lkt} to choose, respectively, urban area k and the reference urban area indexed by 0. We have (see Appendix) $\ln\frac{P_{kt}}{P_{0t}} = \gamma_{kt}$. This implies that $\Delta\gamma_{kt} - \Delta\gamma_{kt-1} = \Delta\ln\frac{P_{kt}}{P_{kt-1}} - \Delta\ln\frac{P_{0t}}{P_{0,t-1}}$. The second term is invariant across cities and it is thus absorbed by the intercept.

of parameters in the first step.[28] Notice that one additional advantage of using a two-step method is that it produces conservative standard error estimates as the number of observations corresponds to the number of locations in the choice set (Donald and Lang, 2007).

To have a causal interpretation of the model estimates, it is also necessary to use changes in public housing supply that are credibly exogenous. The fact that we use singles as a control group and estimate the differential response of families with respect to singles attenuate the risk of bias. However, previous estimates would still be biased if public housing is constructed in response to flows of families of immigrants that are attracted by an unobserved factor that does not attract singles. A successful empirical design involves one or several variables that influence the variations in public housing per household but have no relationship with the unobserved determinants of changes of public housing supply in the municipality conditional on the other covariates. I use two different instrumental variable strategies that are based on two unrelated sources of historical variations in the public housing supply.

First, I use population changes from 1975 to 1999 unrelated to public housing construction to construct an instrument for changes in public housing per household over the period. Glaeser and Gyourko (2005) note that once built, the housing stock does not depreciate quickly. When demand declines, the quantity of housing cannot decline, at least in the short run. I use the fact that housing is durable to predict a counterfactual share of public housing per households using population changes that are orthogonal to public housing constructions and the initial stock of public housing in 1968. As long as the supply of public housing units is not perfectly elastic with population changes, the supply of public housing per household across cities will be influenced by population changes unrelated to public housing constructions. These variations in the population come from changes in labor demand at the national level and are, by definition, not correlated with the unobservable factors influencing both the location choice and the public housing supply conditional on the inclusion of the other covariates in the model. In practice, I predict a counterfactual population growth using standard shift-share models introduced by Bartik (1991) and recently used by Cutler et al. (2008), Saks (2008) and Saiz (2010) among others. The instrument uses the distribution of workers across 11 industries and national growth

[28] The appropriate weight is the inverse of the standard deviation for an observation. Assuming independent sampling, it is the inverse of the sum of the standard errors of the parameter estimates.

rates of each industry to forecast municipality population growth due to the composition effect. The counterfactual population in urban area k and period t is thus computed using:

$$\tilde{N}_{kt} = \sum_i \left(\frac{N_{it}}{N_{i68}}\right) * N_{ik,68} = \sum_i \lambda_i N_{ik,68} \tag{4}$$

where $\frac{N_{it}}{N_{i68}} = \lambda_i$ is the ratio between the number of workers in industry i in t and in 1968 at the national level. Because the variable to be instrumented p_{ijt} is the ratio between public housing units and the number of households, we use the initial stock of public housing in 1968 to predict the evolution of the number of housing units using national construction rates in the stock of public housing. More specifically, the counterfactual stock of housing units in urban area k and period t is computed using:

$$\tilde{PH}_{kt} = PH_{k,68}(1 + g_{t,68}) \tag{5}$$

where $g_{t,68}$ is the observed national growth rate of the number of public housing units in France between t and 1968 reported in panel B of Table 1. We use this counterfactual public housing stock divided by the counterfactual population to define our final instrument as

$$\tilde{p}_{kt} = \frac{\tilde{PH}_{kt}}{\tilde{N}_{kt}} \tag{6}$$

There may be several concerns for the exclusion restrictions of these instruments. Variations from shift-share population growth may capture changes in city characteristics that directly influence changes in the utility of choosing a given city. To address this, I control separately in the regression for the observed distribution of workers across 10 industries used to create the shift-share. Conditional on the inclusion of these controls in the regression, our shift-share variables should only capture shocks related to national changes in these factors.

However, the shift-share counterfactual public housing share may not be an ideal instrument. If unobserved shocks are serially correlated, the initial industrial composition in 1968 used to construct the instrument may be related to the evolution of local unobserved factors in the subsequent decades. To alleviate these concerns, I use a second instrumental variable strategy. I estimate long difference regressions based on the model from Eq. 3 combining only arrivals of new immigrants from 1962 to 1968 and 1990 to 1999 in the sample. As previously discussed, public housing authorities during the 1960s required immigrants to have at least 10 years of residency in France. As a result, we should not expect immigrants who arrived during the 1960s to react to differences in public housing supply across locations (we empirically test this assumption below). Therefore, in this specification, we use the level of public housing in 1990 as a dependant variable. More specifically, we estimate:

$$\Delta\hat{\gamma}^{c}_{k,1999-1968} - \Delta\hat{\gamma}^{s}_{k,1999-1968} = \Delta X_{k,1999-1968}\tilde{\theta} + \tilde{\delta} p_{k,1990} + u_{kt} \tag{7}$$

We still need to account for the potential endogeneity of the public housing share in 1990. To start, we use a long lag as an instrument for $p_{k,1990}$. In particular, we use the share of public housing per household in 1968, the earliest year in which public housing supply stock can be computed after 1945. While it is plausible that the initial public housing share in 1968 is not correlated with unobserved determinants of the attractiveness of the urban area for immigrants who arrived during the 1990s, it may still be the case that there is an unobserved factor highly autocorrelated over time determining the location choice of immigrants. Our second instrument relies on the fact that large differences in public housing share in 1990 are the result of war destruction that has been more substantial in some cities than in others. As described by Voldman (1997), significant destruction was seen as an opportunity by urban planners responsible for the reconstruction of radical changes in some cities. In practice, the large destruction rates in many cities provided enough space to construct large housing projects in the center of the urban area. As a result, we report below that there is a strong correlation between the share of public housing per inhabitants in 1990 and the intensity of the destruction during the Second World War. Note that cities affected by war destruction were located all over France. Many of the destroyed cities were in Normandy, Brittany or the north of France, but destruction was also substantial in the east (Saint-Etienne, Strasbourg, Colmar) and the south of France (Toulon, Marseille) and in several cities in the west (Saint-Nazaire, Royan). To implement this empirical strategy, we construct a 'destruction index' with differences in bombing intensity during the Second World War reported in Florentin (1997) that we have adjusted to take into account the fact that some cities where destroyed during the first part of the war in 1940.[29] In practice, we use the sum of the number of bombings across municipalities of the urban area weighted by the share of the population in the municipality in 1968. We take the square root of the number of bombings to allow for a nonlinear effect of the destruction index.

The validity of these two instruments relies on the hypothesis that they are not correlated with unobservable determinants of the location choice correlated with the public housing supply and the location choice of immigrants. The model of Eq. 7 uses differences between the location choices in 1962 to 1968 and 1990 to 1999 to potentially control for the effect of unobserved but fixed factors influencing the location choice across cities in both the 1960s and the 1990s. Similarly, the fact that we also use the difference between the location choice of families and singles eliminates the effect of unobserved factors that influence similarly both groups. As a result, the requirement for the exogeneity of our instrument is that in practice, the dispersion of Second World War destruction should not be correlated with changes in the unobserved factors in the 1990s with respect to the 1960s, affecting differentially families with respect to singles.

The linear model based on Eq. 3 is estimated using 2SLS. Direct estimates from the conditional logit model account for the endogeneity of the public housing supply with a control function to deal with the potential endogeneity of the changes in public housing supply over time (Petrin and Train, 2010; Liu et al., 2010). This procedure is detailed in the appendix. Standard errors of conditional logit models are estimated using the bootstrap with 100 replications.

3.2 Estimation Sample

We analyze the location choice of immigrants across 57 urban areas of more than 100 000 inhabitants in 1990 by their 1990 definition using constant boundaries. Immigrants admitted for family reunion may not possess the same skills as economic immigrants (Chiswick, 1986), and their location decision may depend on the location of family members already living in France. Therefore, we focus on the location choice of new male immigrants who arrived in France be-

[29] We verified the sensitivity of our results by using the estimated level of destructions calculated in 1945 by the French government and reported by Voldman (1997). The results were qualitatively similar and are available upon request.

tween 1962 and 1999 and exclude individuals reported as a child in a household.[30]

As variables included in $X_{k,t-1}$ control for differences in economic opportunities across cities, we use the differences in unemployment rates across cities and 10 variables that contain the distribution of workers across 11 industries.[31] As a proxy for the socio-demographic characteristics of the urban area, I include the percentage of university graduates, the total immigrant share of the urban area and the log of the population of the urban area. Variables included in $Z_{lk,t-1}$ vary at the ethnic group by urban area level and include the percentage of individuals for the same ethnic group in the urban area, their average number of years of education and the share of immigrants married in the group. Unlike previous studies and because of the large sample extracts available over the chosen period, I distinguish between groups with 54 different countries of birth, which are always reported separately across censuses.[32] Finally, we estimate separate models for European and non-European immigrants. Our intuition is that given the large differences in participation rates documented in Table 3, public housing might be much more attractive to non-European immigrants.

4 Results

4.1 First-Stage Results

To better understand the factors underlying the dispersion of public housing across French cities and its evolution, we provide results from first-stage regressions documenting the relationship between public housing supply and urban area characteristics between 1968 and 1999 in Table 4. Column 1 of the Table first highlights the strong correlation between the Second World War destruction and the initial stock of public housing. The coefficient is strongly significant, indicating an increase of 4% of the share of public housing per household when the destruction index increases by 1. The inclusion of additional covariates as a control in column 2 only

[30] There is no information reported in the census on the admission category of new immigrants, but according to the best available figures from Tavan et al. (2005, p.72), 80% of immigrants admitted for family reunion are female and most of the others are children.

[31] No information on wages is collected across censuses

[32] Unreported estimates show that the estimated effect of immigrant concentration is much lower when the indices of immigrant concentration are defined by region of birth instead of country of birth.

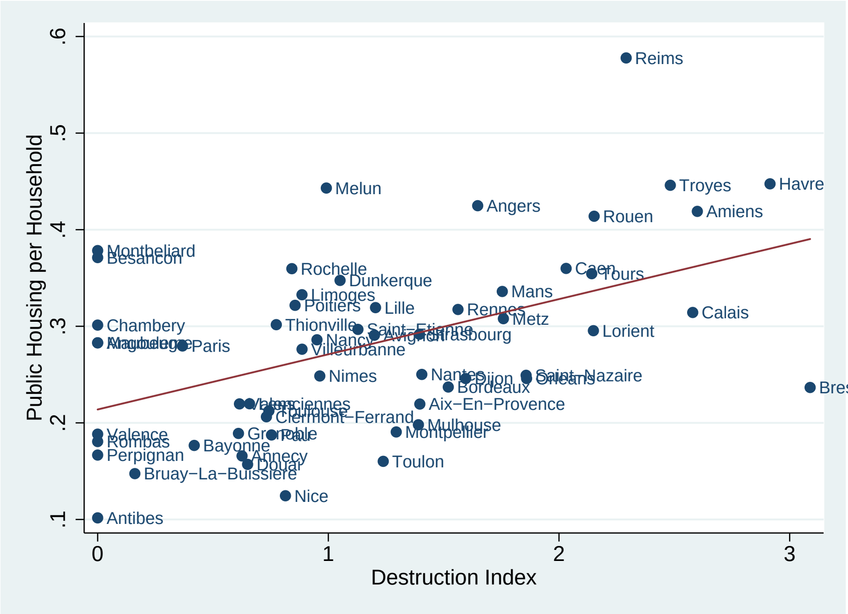

Figure 2: Public Housing Supply in 1990 and War Destructions

Notes: the Figure plots the relationship between public housing per household in 1990 and our destruction index during the Second World War. *Sources:* Public housing per household from the 1990 Census. Destruction Index computed using bombings and initial.

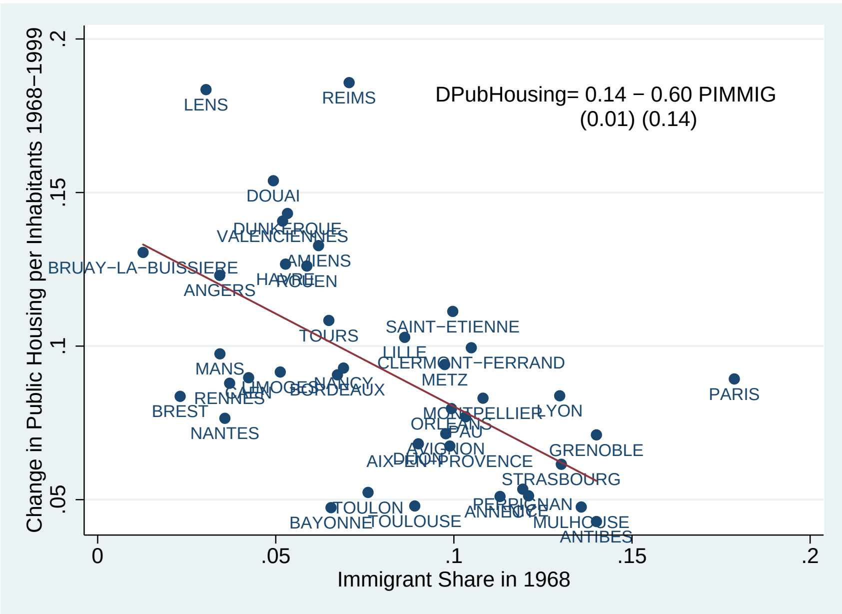

Figure 3: Public Housing Supply Increase 1968-1999 and Initial Immigrant Stock in 1968

Sources: 1999 and 1968 Census of Population and 1999 Census of Housing. The Figure plots the relationship between the change in public housing supply between 1968 and 1999 and the initial share of immigrant in the population.

decrease moderately the correlation between bombing and the initial stock of public housing, which still remains strongly significant. The results also indicate that the public housing supply in 1968 was larger in cities with a lower unemployment rate, suggesting that public housing was constructed in cities with a higher housing demand in 1968.

In columns 3 and 4, we examine the relationship between city characteristics and the dispersion of the public housing supply in 1990. We also find a significant effect of Second World War destruction in explaining the dispersion of public housing in 1990, thus indicating a lasting effect of the differences in initial constructions over time. Figure 2 illustrates the relationship between public housing per household in 1990 and our index of war destruction. The positive correlation between public housing supply in 1990 and our destruction index is strong and is not driven by a particular outlier. Excluding Rouen and Brest from the sample provides identical results. Regression results also indicate that other attributes of the urban area do not appear correlated with public housing. Column 4 introduces in the regression the counterfactual share of public housing constructed using the shift-share model. Its effect is also strongly significant, suggesting that a significant part of the dispersion of public housing per household can be explained by population changes and the persistence of the initial stock.

Interestingly, in all cross-section models, we find no effect of the share of immigrants in the population either in 1968 or in 1999, confirming that the dispersion of the public housing stock in 1968 and in 1990 cannot be explained by the fact that cities with a higher immigrant share constructed more public housing. Additional evidences on the relationship between the initial immigrant share of the urban area and changes in public housing supply are provided in Figure 3, which reports the correlation between changes in public housing per inhabitants between 1968 and 1999 and the initial immigrant stock. The figure indicates that the public housing stock per household increased more rapidly in cities with a lower initial share of immigrants.[33]

Thus far, we have documented the relationship between the cross-section dispersion of public housing and the characteristics of cities. We now investigate the correlations between changes in public housing supply from 1975 to 1990 and city characteristics. Columns 5 and 6 use the variations of public housing per households from 1975 to 1990 as a dependant variable

[33] This relationship remains robust when non-European or new immigrants are chosen rather than the share of immigrants.

and include urban area and year fixed effects in the model and also 10 variables for the distribution of workers across 11 industry controls. The table confirms that within cities, increases in public housing per household are correlated with an increase in the unemployment rate, in the manufacturing share and in a decrease in the graduate share. Interestingly, the coefficient of the immigrant share is positive, which indicates that the share of immigrants tend to increase in cities in which public housing increases, which is consistent with the hypothesis put forth in this paper. Column 6 includes the counterfactual public housing share constructed with the shift-share model based on the initial distribution of industries in 1968 and thus is, in practice, the first-stage equation used to estimate the control function of the conditional logit model. The estimated parameter of this counterfactual share is strongly significant. This indicates that population shocks combined with the fact that public housing is durable can explain a large share of the variations in public housing supply between 1975 and 1990.

4.2 Second Stage results

In this section, I present estimates of the effect of differences in public housing supply across French urban areas on the initial location choice of immigrants.

4.2.1 Conditional Logit Results

We first present estimates of models in which the effect of public housing is estimated directly in the conditional logit model in Table 5. Separate models were estimated for Europeans and non-Europeans. The results strongly indicate that after 1968, public housing influenced the location choice of new non-European immigrant couples with children. The estimated parameters imply that for a non-European immigrant couples, an increase of one standard deviation in the public housing supply (approximately 5% over the period) increases the probability of choosing the "average" urban area by approximately 23% for non-Europeans. Not surprisingly, given that European immigrants are not particularly over-represented in public housing with respect to natives, we do not find strong evidences that differences in public housing supply had an effect on European immigrants. Inclusion of the control function in column 2 does not substantially change the results for non-European immigrants, and the residuals are not signifi-

Table 4: First Stage Results

	Dependant variable: Public Housing per Household					
	1968 share		1990 share		1975-1990	
	(1)	(2)	(3)	(4)	(5)	(6)
Unemployment		-0.033***	-0.008	0.000	0.010***	0.009***
Rate		(0.007)	(0.007)	(0.003)	(0.002)	(0.002)
Immigrant		0.001	-0.002	-0.005***	0.005**	0.004**
Share		(0.001)	(0.003)	(0.001)	(0.002)	(0.002)
Av. Age		-0.004	-0.023	0.002	-0.004	-0.006
		(0.008)	(0.014)	(0.007)	(0.008)	(0.008)
Manufacturing		-0.002**	0.000	0.000	0.006*	0.006*
Share		(0.001)	(0.002)	(0.001)	(0.004)	(0.003)
Graduate		-0.024**	-0.001	-0.003	-0.007***	-0.006***
Share		(0.010)	(0.005)	(0.003)	(0.002)	(0.002)
Log(pop.)		0.007	-0.016	-0.004	0.063	0.057
		(0.008)	(0.015)	(0.007)	(0.043)	(0.042)
War destruction	0.044***	0.030***	0.057***			
Index	(0.011)	(0.011)	(0.019)			
Predicted				0.647***		0.371***
Public Housing				(0.047)		(0.114)
N	57	57	57	57	171	171
R^2	0.23	0.57	0.37	0.851	0.97	0.97
Urban area FE					Yes	Yes
Year FE					Yes	Yes
Industries Controls					Yes	Yes

Notes: Each column in the table represents the coefficient of a regression of the public housing per households on the indicated variables. Column 1 and 2 uses the share in 1968 while column 3 and 4 uses the share in 1990. Column 5 and 6 use public housing per households in 1975, 1982 and 1990 as a dependant variable. Column 5 and 6 include indicator variables for urban area and year fixed effects. Column 5 and 6 also includes the distribution of workers across 10 industries. The variables *War destruction Index* is the square root of the estimated number of bombing campaign in the urban area during the Second World War. The variable *Predicted Public Housing* is a counterfactual evolution of public housing constructed using a shift-share model of the evolution of the population.

Table 5: Determinants of Location Choice : 57 Cities 1962-1999, Male living as a couple with Children

	Non Europeans		Europeans	
Variable varying at the ethnic group level				
	(1)	(2)	(1)	(2)
Share Similar	0.280***	0.280***	0.456***	0.456***
Immigrant in City	(0.009)	(0.009)	(0.008)	(0.008)
Av Educ. Similar	-0.034**	-0.034**	-0.080***	-0.079***
Immigrant	(0.013)	(0.013)	(0.016)	(0.016)
Share Couples	0.140***	0.140***	0.007	0.007
Similar Immigrants	(0.012)	(0.012)	(0.013)	(0.013)
Av Age Similar	-0.123***	-0.123***	-0.181***	-0.180***
Immigrants	(0.013)	(0.013)	(0.014)	(0.014)
Variables varying at the urban area level				
Log(Pop)	1.440***	1.441***	-0.137	-0.125
	(0.324)	(0.324)	(0.334)	(0.335)
Immigrant	0.253***	0.254***	0.049	0.055
Share	(0.063)	(0.063)	(0.078)	(0.078)
Graduate	-0.379***	-0.376***	-0.507***	-0.468***
Share	(0.092)	(0.098)	(0.112)	(0.114)
Manufacturing	-1.080***	-1.081***	0.291	0.266
Share	(0.194)	(0.195)	(0.254)	(0.254)
Av. Age	-0.043	-0.043	0.038	0.047
City	(0.071)	(0.071)	(0.077)	(0.077)
Unemployment	-0.049**	-0.050**	-0.217***	-0.229***
Rate	(0.020)	(0.024)	(0.024)	(0.025)
Public Housing	0.224***	0.233**	0.050	0.152*
per Household	(0.068)	(0.110)	(0.071)	(0.092)
Public Housing	-0.241***	-0.233**	-0.104	-0.004
in 1968	(0.072)	(0.110)	(0.073)	(0.092)
Control		0.137		1.934*
Function		(1.420)		(1.091)
N Individuals	26 294	26 294	28 287	28 287

Notes: The table presents estimates from a conditional logit of the location choice of male new immigrants who arrived in France between 1962 and 1999. The location choice of immigrants observed in census in year t who arrived between census year t and year $t-1$ is regressed on characteristics of the urban area and group in year $t-1$. Standard errors of models with a control function have been obtained using bootstrap using 100 replications. The model also includes 10 variables for the distribution of workers of the urban area across 11 industries. The sample only includes individuals who declared to live in couple with children at the time of the census.
Source: Censuses of Population 1968-1999 and Censuses of Housing 1999.

cant. For European immigrants in column 4, the residuals enter significantly, and the coefficient is positive, while the coefficient of the effect of public housing supply increases. The results are also consistent with the historical evidences that public housing was not accessible to immigrants during the 1960s. Parameter estimates of the differences in public housing supply in 1968 are even negative and strongly significant in that year for non-European immigrants, while they are not significantly different from zero for European immigrants.

Assuming immigrants who are single are not attracted by differences in public housing, direct estimates of the effect of public housing on this group offer a first falsification test for whether the estimates are biased by potential unobserved confounding factors. If the impact of public housing that I capture in the previous regressions is not the result of spurious correlations, I should find much less impact of public housing on immigrants who are less likely to be eligible in the short run, namely, immigrants without children. Panel A of Table 6 presents estimates of the same model using new immigrants who are single. The results strongly indicate that public housing does not have a magnetic effect on immigrants who are singles. For both groups, Europeans and non-Europeans, the effect of differences on public housing supply across urban areas is negative, which suggests that variations in the public housing supply may be correlated to undesirable characteristics of the urban area for this group. The negative coefficient for the residuals in columns 2 and 4 is actually consistent with this interpretation.

Panels B and C of Table 6 report estimates of immigrants with children for several alternative specifications of the model. To probe the possibility of reverse causality, panel B of Table 6 reports estimates of models that include lagged and future measures of the presence of public housing supply using a dynamic version of Eq. (1):

$$U_{ikt} = Z_{ikt-1}\theta_1 + X_{kt-1}\theta_2 + \delta_{lag}p_{k,t-1} + \delta_{lead}p_{k,t+l} + \gamma_k + \epsilon_{ikt}$$

Including the lead terms in the model enables a direct test for endogenous shifts of public housing supply following immigration. If the estimation of the lead coefficient δ_{lead} is significant and positive, this can be interpreted as evidence of reverse causality, whereas an estimate close to zero is consistent with the absence of such an effect. Reassuringly, the results indicate that lagged public housing supply appears to have the highest impact on location choice.

Table 6: Alternative Location Choice Models

	Non Europeans		Europeans	
A. Sample Contains only Singles				
Public Housing	-0.066	-0.144**	-0.248***	-0.478***
per Household	(0.043)	(0.059)	(0.067)	(0.083)
Public Housing	-0.272***	-0.349***	-0.218***	-0.446***
in 1968	(0.044)	(0.059)	(0.068)	(0.083)
Control		-1.365*		-5.030***
Function		(0.706)		(1.055)
N Indiv	86 185	86 185	42 374	42 374
B. Lead Test & heterogeneity; Couples with Children				
Public Housing	0.191**		0.115	
(t-1)	(0.079)		(0.080)	
Public Housing	0.048		-0.087*	
(t+1)	(0.060)		(0.048)	
Public Housing		0.179**		-0.053
Supply * 1990		(0.072)		(0.077)
Public Housing		0.246***		-0.062
Supply * 1982		(0.071)		(0.077)
Public Housing		0.244***		0.231***
Supply * 1975		(0.072)		(0.076)
Public Housing		-0.226***		0.016
Supply * 1968		(0.073)		(0.076)
N Indiv	26 294	26 294	28 287	28 287
C. Interaction Life in Public Housing x PH supply ; Sample contains Couples				
Public Housing	0.496***	0.451***	0.333***	0.332***
x Life in PH	(0.070)	(0.072)	(0.066)	(0.066)
Public Housing	-0.108	-0.153**	-0.314***	-0.312***
x Life in PV	(0.069)	(0.071)	(0.058)	(0.058)
Control		-2.325**		-2.517***
Function		(0.962)		(0.943)
N Indiv	26 294	26 294	28 287	28 287
D. Additional Robustness Checks on Non-Europeans Immigrants in Couples				
	Only Couple	Excluding from the Choice Set:		
	with Old Child	Paris	5th biggest ua	20th biggest ua
Public Housing	0.419***	0.258***	0.226***	0.235**
Supply	(0.098)	(0.069)	(0.075)	(0.100)
N indiv	12 036	28 786	7 852	3 761

Sources and Notes: Within each panel, each column reports selected parameter estimates of a conditional logit of the location choice of immigrants on the indicated variables. The sample in panel A includes male new immigrants singles while the sample in panel B and C includes male new immigrants living as a couple with children. The sample in panel D includes only male non Europeans new immigrants. Models in column 1 and 2 in each panel except D use non-European immigrants from these groups, while models in column 3 and 4 use non-European immigrants. All additional control variables included in models of Table 5 are also included. The first column in panel D use male non-European living as a couple with a child born before the previous census. Columns 2, 3 and 4 exclude respectively from the choice set Paris, the 5th biggest cities and the 20th biggest cities.

Interestingly, the coefficients on the leads are even negative for European immigrants, though they are statistically insignificant for non-European immigrants. While these tests are far from definitive, they do provide some evidence that the relationship between changes in public housing supply per household and immigrant inflows is not driven by serious problems of reverse causality.

It might also be possible that our results reflect the impact of public housing in a given period of time. If this is the case, this may be troublesome as this might reflect the effect of a particular unobserved event attracting immigrants in some urban areas. To investigate this, columns 2 and 4 in panel B allow for a heterogenous effect of differences in the public housing supply from 1968 to 1990 so as to investigate the potential differences in the effect of public housing across years. For non-European immigrants, parameter estimates are remarkably similar for arrivals from 1968 to 1975 and 1975 to 1982, while the coefficient decreases slightly for arrivals between 1982to 1990. For European immigrants, only arrivals from 1968 to 1975 appear to have been influenced by differences in public housing supply.

In panel C, we report estimates of models that allow the impact of differences in public housing supply to differ between new immigrants that are observed living in public housing and those in private housing after their arrival in France. We estimate the following model:

$$U_{ikt} = Z_{ikt-1}\theta_1 + X_{kt-1}\theta_2 + \delta_{PH}p_{k,t-1}PH_i + \delta_{PV}p_{k,t-1}(1 - PH_i) + \gamma_k + \epsilon_{ikt}$$

where the variable PH_i is equal to one if a new immigrant i lives in public housing in period t. We estimate this model using data on public housing participation that are available in the 1982, 1990 and 1999 censuses. We expect the effect of differences in public housing supply to be higher for those living in public housing. If we also find a strong effect of public housing supply on those observed in private housing during the census year, this may suggest that another factor correlated with public housing attracted those immigrants. However, the results in panel C indicate a strong effect of differences in public housing supply on immigrants observed as living in public housing but not on immigrants living in private housing. This also suggests that new immigrants living in public housing during the census year live in cities with a higher public housing supply. On the other hand, differences in the public housing supply appear to be negatively correlated with the location choice of new immigrants observed as living in private housing.

Finally, the parameters of other covariates included in the model and reported in Table 4 for immigrant couples have the expected sign and are consistent with previous evidences from the literature. First, the results indicate that immigrants prefer cities in which similar immigrants make up a larger percentage of the population. Interestingly, this effect is stronger for European immigrants than for non-European immigrants. Surprisingly, the share of educated individuals from the ethnic group has a negative sign as does the average age of the group. Non-European immigrants also appear to be attracted by cities in which a large share of the members of their ethnic group live as couples. The results also indicate that immigrants strongly prefer cities with a lower unemployment rate, but the effect of the difference in unemployment rates is five times larger for European immigrants.[34] Across cities, an increase of one standard deviation of the unemployment rate in a given year decreases the probability to choose the urban area by approximately 22% for immigrants from Europe but only by 5% for immigrants from Maghreb.

4.2.2 Long-Difference Regressions

There are several potential problems with the previous results. First, if unobserved shocks are serially correlated, the destination choice of migrants may be correlated across decades, which may bias the panel data estimates reported in the previous section. Second, estimated standard errors for the conditional logit may be underestimated as the error term related to the choice of a given urban area may be correlated between cities across individuals. To account for these issues, Table 7 provides results using the two-step model described in the previous section where the dependant variable captures the relative probability to choose an urban area once the effects of differences in the characteristic of the ethnic group across cities have been taken into account. To save space, we focus on non-European immigrants in this section.[35] We also attempted to account for heteroscedasticity when evaluating the standard errors with

[34]This result has an important implication for the research on the impact of immigration on the labor market; as emphasized by Borjas et al. (1997), if migrants locate in cities with booming economies, estimates of the impact of immigration from methods using correlations across cities will be biased upward.

[35]As in the previous section, results with European immigrants do not indicate a significant effect of public housing on the location choice for this group. The results are available upon request.

robust standard error estimates. However, we obtained, in practice, standard errors inferior to the classical one. Following the advice of Angrist and Pischke (2008, chapter 8), we report the maximum of the estimated standard errors, which are, in this case, the classical standard errors.

The table reports estimates of two different models: column 1 and 2 use data from 1975 to 1999 and thus includes an observation for each census from 1975 to 1999. Column 3 to 5 of the table present "long-difference" estimates using first difference between parameter estimated using arrivals of immigrants in 1962-1968 and 1990-1999, leaving 30 years between pre and post-periods. In this model, the public housing supply is introduced by using directly the 1990 value under the assumption that differences in public housing supply in 1968 had no effects on the location choice. Various instrumental variable strategies are used to account for the potential endogeneity of changes in the public housing supply across models. Column 2 uses variations in public housing supply from the shift-share model. Column 4 uses the share of public housing in 1968 to predict the share in 1990 while column 5 uses the predicted public housing share in 1990 with the shift-share model. Column 6 uses the instrument based on destructions related with the Second World War.

We first evaluate separately the impact of public housing on singles and on families in, respectively, panels A and B by reporting estimates of models in which the dependant variables are the change in the probability to choose an urban area for these two groups. The model described in Eq. 3 using differences in relative probability between families and singles is reported in panel C. Parameter estimates from panel data model in columns 1 and 2 indicate a significant effect of public housing supply on the location choice of couples with children and on the relative location choice of couples with respect to singles. Estimates for couples are slightly larger than in the conditional logit model, indicating an increase of 48 log points in the probability to choose the urban area. Estimates in panel C using the differences in settlement patterns between families and singles indicate a coefficient of approximately 0.28 in both OLS and 2SLS estimates, implying that for every standard deviation increase of public housing, the probability to choose an urban area in 1999 with respect to 1968 for an immigrant with children increases by 28 percentage point over that for singles. As previously determined, we do not find any effect on the location choice of immigrants who are singles with parameter estimates being positive, but insignificant, for this group.

Turning now to the long-difference estimates in columns 3 through 6, we find a positive effect that is strongly significant on the relative location choice both in the OLS and the IV estimates. The IV point estimates in columns 4 and 5 are never markedly different from their OLS counterparts in column 3. The parameters suggest that the log probability to choose a given urban area in 1999 with respect to its probability in 1968 increases by 52 log points to 58 log points for couples with respect to singles if the supply of public housing in that urban area increases by one standard deviation in 1990. The magnitude of the effect is much greater when war destruction is used as an instrument, but the point estimates are more imprecise though significantly different from zero. This last model indicates an increase of approximately 116 log points for the relative probability of choosing the urban area.

Figure 4 graphs the relationship between relative probability and public housing supply in 1990. The positive correlation between the relative change in the location choice of immigrants and public housing supply in 1990 is strong, and it is not driven by a particular outlier. Excluding Reims and La Rochelle from the sample does not affect, in practice, the results.

4.3 Additional Robustness Checks

I evaluate the robustness of the previous estimates in several ways. I first examine the potential source of bias related to the possibility that the family status of immigrants may be a response to differences in public housing supply, thus leading to a bias in the estimates. As previously discussed, immigrants may have been attracted to cities with a large public housing supply for another reason and could have subsequently responded to the supply of public housing by having a higher probability to have children in places with a higher public housing supply. I test for this using the subsample of immigrants with children born before the arrival of the immigrant in France.[36] Results of such models are provided in the first column of panel D in table 6. The results indicate that parameter estimates are strongly significant and higher than those obtained previously using the whole sample. This suggests that the previous results are not driven by an

[36] Because I have no information on the exact arrival year, the sample contains households where the age of the oldest child is superior to the period of time elapsed since the previous census in which the immigrant reported he was not living in France.

Table 7: Two-Step Model: Panel and Long-Difference Regressions

	(1)	(2)	(3)	(4)	(5)	(6)
	Panel Regressions 1975-1999		Long-Difference Regressions 1968 - 1999			
	Dependant Variable:					
	A. Location Choice Couples with Children (C)					
Public Housing	0.484***	0.484***	0.303*	0.259	0.294*	1.186*
Supply	(0.112)	(0.120)	(0.172)	(0.182)	(0.176)	(0.610)
	B. Location Choice Singles (S)					
Public Housing	0.187	0.180	-0.130	-0.142	-0.167	0.222
Supply	(0.102)	(0.110)	(0.162)	(0.168)	(0.163)	(0.444)
	C. Difference Location Choice (C-S)					
Public Housing	0.290***	0.275**	0.548***	0.522***	0.579***	1.161**
Supply	(0.101)	(0.107)	(0.191)	(0.201)	(0.194)	(0.588)
N	168	168	56	56	56	56
Method	OLS	2SLS	OLS	2SLS	2SLS	2SLS
Instrument		CF Public Housing		PH Share 1968	CF Public Housing	War Destruction Index

Sources and Notes: Each entry in this table represents the coefficient on public housing per household from a separate regression. Each regression is weighted using the inverse of the standard deviation of the estimates of the dependant variable. Standard error are in parentheses. The dependant variable in panel A and B is respectively for families and singles the relative probability estimated with a separate conditional logit to choose one of the 56 urban area with respect to the reference urban area. Panel C uses the difference in the relative probability between families and singles as a dependant variable. All models are estimated in first differences. Column 2 and 5 use the counterfactual share of public housing predicted by the shift share model as an instrumental variable for the public housing per household. Column 4 uses the share of public housing per household in 1968. Column 6 uses the war destruction index as an instrument.

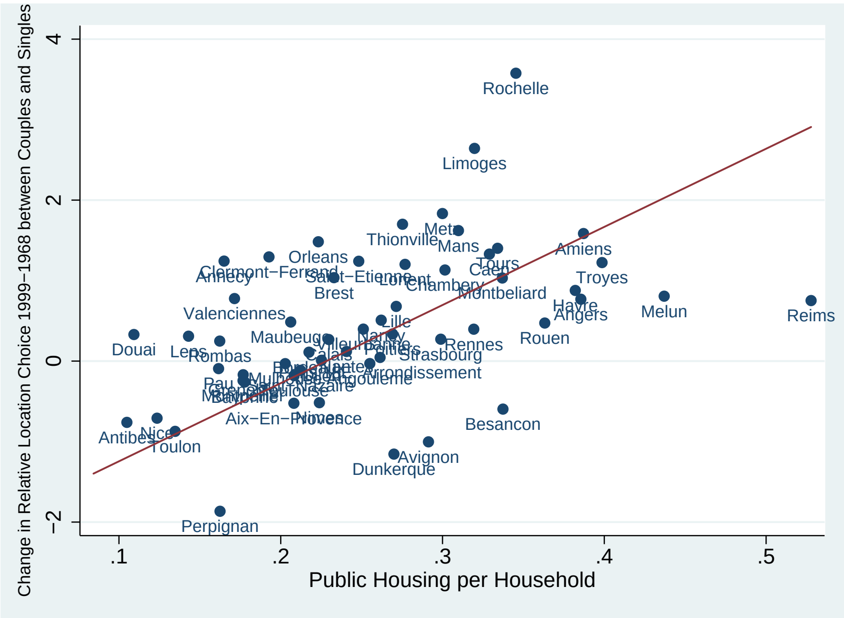

Figure 4: Changes in Relative Location Choice between Families and Singles from 1990 to 1968 and Public Housing per Household in 1990

Notes: the Figure plots the relationship between public housing per households and changes in relative location choice of immigrants in couples with respect to singles between those arrived from 1962 to 1968 and 1990 to 1999. *Sources:* Public Housing per Household from the 1990 Census.

endogenous response of the fertility decisions by immigrants to the public housing supply.

One concern with the conditional logit model is that results are sensitive to the definition of the choice set. I have checked the sensitivity of the estimations to the inclusion or exclusion of several alternatives. More specifically, I have estimated models that exclude the Paris urban area from the choice set and models that include or exclude the twentieth biggest urban areas from the choice set. These estimates are reported in columns 2, 3 and 4 of panel D in table 6. The impact of public housing on location choice in these estimates is qualitatively unaffected by such changes in the choice set. As a consequence, these results do not offer evidence against the use of a conditional logit. Finally, the appropriate control function is a specification issue. While I tried other specifications, including a quadratic error term, these alternative specifications all provided similar results and are available upon request.

5 Conclusion

The past several decades have been characterized by a significant increase in public housing supply across European countries. This study finds relatively robust evidence that the availability of public housing influences the location choices of new immigrants with children, particularly non-European immigrants. Immigrant couples tend to choose cities with a relatively higher supply of public housing.

The reason that the impact of public housing on immigrants differs widely between European and non-European immigrants and between immigrants and natives across Europe is deserving of more research. For example, it remains to be determined whether the over-representation of non-European immigrants in public housing is due to specific financial constraints, discrimination in the housing market, or a low supply of cheap housing for families in the private housing market.

The fact that public housing influences the location choice of non-European immigrants is troublesome, particularly in France, where a generous welfare state provides a similar level of benefits across the country. Given that public housing individuals are less mobile, immigrants in public housing may be trapped in cities with very unfavorable labor market prospects but where the relative benefits of living on welfare negate or compensate for the higher costs of living. Such a mismatch between labor demand and the location choice of some immigrant groups may help explain part of the large gap between the unemployment rate of immigrants and natives in France (Decreuse and Schmutz, Forthcoming).

Appendix

5.1 Interpretation of parameters of conditional logit with standardized variables

In this section, I show that the parameters of a conditional logit where the predictors have been standardized such that the variables of the choice set of each individual have an average of zero and a variance of one have a simple and intuitive interpretation. See Gelman (2008) for a more general discussion on the interest of scaling predictors of regressions model. Suppose the true model is given by Eq. (1). Denote by $z_j^k = \frac{x_j^k - \bar{x}^k}{\sigma_{x^k}}$ the standardized variable of the predictor k of alternative j; $\bar{x}^k$ and σ_{x^k}, respectively, the average and the standard deviation of the predictor k over the initial choice set. Because only differences in utility matter (Train, 2003, p.23), the model described by (1) can be rewritten as $ZU_{iJ} = z_{iJ}\gamma + \epsilon_{ij}$, where the relation between β and γ is simply given by $\beta_k = \frac{\gamma_k}{\sigma_{x^k}}$ for all predictor k.

Let me consider the counterfactual probability of choosing two alternatives not included in the initial choice set. The first is the 'average' urban area for which the characteristics are equal to the average of the J preexisting alternatives. The second is identical to the 'average' urban area except that the characteristic l is equal to the average plus one standard deviation. When the predictors have been standardized, the characteristics of the average urban area are a vector of zero, whereas the vector of characteristics of the other additional alternative z is $z^l = 1$ and $z^k = 0$ for $\forall k \neq l$. The probability P of the average alternative is equal to $P = \frac{1}{1+exp(\gamma_l)+\sum_j^J exp(z_j\gamma)}$ whereas the probability P_l for the other alternative is $P_l = \frac{exp(\gamma_l)}{1+exp(\gamma_l)+\sum_j^J exp(z_j\gamma)}$. It is straightforward to derive that $\frac{P_l}{P} = exp(\gamma_l)$ which implies that $\log P_l - \log P = \gamma_l$. The previous expression indicates that the parameter γ_l is equal to the log difference between the probability of choosing the 'average' urban area and the probability of choosing the average urban area in which predictor l is higher by one standard deviation when both cities are included in the choice set. Note that the relationship between β_l and γ_l is a function of the variance σ_{x^l} and therefore γ_l depends on the initial alternatives included in the choice set and used to standardize the variables.

5.2 Control Function Approach

Let me specify the error in the utility function as a two-component error: $\epsilon_{ilkt} = \rho\xi_{kt} + \varepsilon_{ilkt}$ where ε_{ilkt} is an idiosyncratic error term, assumed to be independent across individuals and locations, while ξ_{kt} is observed by immigrants and influences their location choice but is unobserved by the researcher. If p_{kt} and ξ_{kt} are related, such that, for example, a higher level of public housing is constructed in cities in which the unobserved factor is higher, ϵ_{ilkt} and p_{kt} will be correlated, even after conditioning on other covariates. This correlation violates the weak-exogeneity requirement for conditional logit covariates and leads to inconsistent parameter estimates. Petrin and Train (2010) illustrates how the use of a control function can be used to test and correct the omitted-variables problem. Denote as H_{kt} as an n-vector of a variable varying at the urban area level not included in X_{kt} but correlated with p_{kt}. This variable does not affect utility directly but only through its relationship with p_{kt}. The linear projection of p_{kt} on the exogenous variables is $p_{kt} = X_{kt}\pi_1 + \pi_2 H + \eta_k + \mu(\xi)$ which implies $E(X'\mu) = 0$ and $E(H'\mu) = 0$ where $\mu_{kt}(\xi_{kt})$ is one to one with ξ. The method consists in first performing a linear regression of p on X, H and the fixed effects η_k to obtain a consistent estimate of π. The residual is then used to construct the control function $f(\mu, \lambda)$, where μ is the disturbance from the first stage regression and λ a vector of estimated parameters. Denoting by W the vector of all variables included in the location choice model, utility can now be written: $U_{ikt} = W_{ikt}\theta + f(\mu_{kt}, \lambda) + \eta_{ikt}$, where the new error term, $\eta_{ikt} = \beta\xi_{kt} - f(\mu_{kt}, \lambda) + \varepsilon_{ikt}$, includes the difference between the actual specific error $\beta\xi_{jt}$ and the control function plus the idiosyncratic error term. With additivity and independence assumptions, μ_{kt} are straightforward to recover using OLS. Residuals from these regressions can be thus used to estimate the control function.

References

Algan, Yann, Camille Hémet, and David Laitin, "Diversity and Public Good in a Republican Setting," January 2011.

Angrist, Joshua. D. and Jorn-Steffen Pischke, *Mostly harmless econometrics: An empiricist's companion*, Princeton Univ Pr, 2008.

Aydemir, Abdurrahman and George J. Borjas, "Attenuation Bias in Measuring the Wage Impact of Immigration," *Journal of Labor Economics*, 2011, *29* (1), 69–112.

Bartel, Ann P., "Where Do the New U.S. Immigrants Live?," *Journal of Labor Economics*, 1989, *7* (4), 371–391.

Bartik, Timothy J., *Who Benefits from State and Local Economic Development Policies?* number wbsle. In 'Books from Upjohn Press.', W.E. Upjohn Institute for Employment Research, November 1991.

Bauer, Thomas, Gil Epstein, and Ira Gang, "Enclaves, language, and the location choice of migrants," *Journal of Population Economics*, November 2005, *18* (4), 649–662.

Baum-Snow, Nathaniel and Justin Marion, "The effects of low income housing tax credit developments on neighborhoods," *Journal of Public Economics*, 2009, *93* (5-6), 654–666.

Beaudry, Paul, Mark Doms, and Ethan Lewis, "Should the Personal Computer Be Considered a Technological Revolution? Evidence from U.S. Metropolitan Areas," *Journal of Political Economy*, 2010, *118* (5), 988 – 103.

Boeldieu, Julien and Suzanne Thave, "Le logement des immigrés en 1996," *INSEE Première*, August 2000, *730.*

Boeri, Tito, "Immigration to the Land of Redistribution," *Economica*, October 2010, *77* (308), 651–687.

Borjas, George J., "Immigration and Welfare Magnets," *Journal of Labor Economics*, 1999, *17* (4), 607–637.

_ , "Does Immigration Grease the Wheels of the Labor Market?," *Brookings Papers on Economic Activity*, 2001, (1), 69–119.

_ , Richard B. Freeman, and Lawrence F. Katz, "How Much Do Immigration and Trade Affect Labor Market Outcomes?," *Brookings Papers on Economic Activity*, 1997, (1), 1–90.

Bouvard, Laurence, Pierre-Philippe Combes, Bruno Decreuse, Benoît Schmutz Morgane Laouénan, and Alain Trannoy, "Géographie du chômage des personnes d'origine africaine: Une discrimination sur le marché du logement ?," *Revue Française d'Économie*, 2009, *23* (1), 8–56.

Cascio, Elizabeth U. and Ethan G. Lewis, "Cracks in the Melting Pot: Immigration, School Choice, and Segregation," *American Economic Journal: Applied Economics*, 2011.

Chiswick, Barry R., "Is the New Immigration Less Skilled Than the Old?," *Journal of Labor Economics*, 1986, *4* (2), 168–192.

Cutler, David M., Edward L. Glaeser, and Jacob L. Vigdor, "When are ghettos bad? Lessons from immigrant segregation in the United States," *Journal of Urban Economics*, 2008, *63* (3), 759 – 774.

Decreuse, Bruno and Benoit Schmutz, "Mobilité résidentielle et emploi des immigrés d'origine africaine en France: une approche par calibration," *Annales d'Économie et de Statistique*, Forthcoming.

Desplanques, Guy and Nicole Tabard, "La localisation de la population étrangère," *Economie et Statistiques*, April 1991, *242*, 51–62.

Dettling, Lisa J. and Melissa Schettini Kearney, "House Prices and Birth Rates: The Impact of the Real Estate Market on the Decision to Have a Baby," Working Paper 17485, National Bureau of Economic Research October 2011.

Donald, Stephen G. and Kevin Lang, "Inference with Difference-in-Differences and Other Panel Data," *The Review of Economics and Statistics*, 03 2007, *89* (2), 221–233.

Durif, Pierre and Olivier Marchand, "Les locataires des HLM en 1973," *Économie et Statistique*, 1975, *73*, 3–20.

Florentin, Eddy, *Quand les Alliés bombardaient la France, 1940–1945*, Paris, Perrin, 1997.

Fougère, Denis, Francis Kramarz, Roland Rathelot, and Mirna Safi, "Social Housing and Location Choices of Immigrants in France," IZA Discussion Papers 5557, Institute for the Study of Labor (IZA) March 2011.

Gelman, Andrew, "Scaling regression inputs by dividing by two standard deviations," *Statistics In Medicine*, 2008, *27* (15), 2865–73.

Giorgi, Giacomo De and Michele Pellizzari, "Welfare migration in Europe," *Labour Economics*, 2009, *16* (4), 353 – 363.

Glaeser, Edward L. and Joseph Gyourko, "Urban Decline and Durable Housing," *Journal of Political Economy*, April 2005, *113* (2), 345–375.

Gobillon, Laurent, "Emploi, logement et mobilité résidentielle," *Economie et statistique*, 2001, *349* (1), 77–98.

Guillouet, Jean-Jacques and Philippe Pauquet, "La demande de logements locatifs sociaux en Île-de-France en 2010," *Institut d'Aménagement et d'Urbanisme d'Île de France*, May 2011.

Hunt, D. Bradford, *Blueprint for disaster: The unraveling of Chicago public housing*, University of Chicago Press, 2009.

Jaeger, David A., "Local Labor Markets, Admission Categories, and Immigrant Location Choice," 2001. Hunter College and Graduate Center.

_ , "Green Cards and the Location Choices of Immigrants in the United States, 1971-2000," *Research in Labor Economics*, 2008, *27* (1), 131–183.

Jayet, Hubert and Nadiya Ukrayinchuk, "La localisation des immigrants en France: Une première approche," *Revue d'économie régionale et urbaine*, 2007, *4* (11), 625–649.

Kaushal, Neeraj, "New Immigrants' Location Choices: Magnets without Welfare," *Journal of Labor Economics*, January 2005, *23* (1), 59–80.

Le Blanc, David, Anne Laferrère, and Rémy Pigois, "Les effets de l'existence du parc HLM sur le profil de consommation des ménages," *Économie et Statistique*, 1999, *8* (328), 37–59.

Lequin, Yves, "Immigrés en Ville," in Yves Lequin, ed., *Histoire des Étrangers et de l'Immigration en France*, Larousse 2006, pp. 408–429.

Liu, Xuepeng, Mary E. Lovelt, and Jan Ondrich, "The Location Decision of Foreign Investors in China: Untangling the Effect of Wages Using a Control Function Approach," *The Review of Economics and Statistics*, 2010, *92* (1), 160–1666.

Mincer, Jacob, "Family Migration Decisions," *Journal of Political Economy*, October 1978, *86* (5), 749–73.

Moretti, Enrico, "Chapter 14 - Local Labor Markets," in Orley Ashenfelter and David Card, eds., *Handbooks in Economics Handbook of Labor Economics, Vol 4B*, Vol. 4, Part B of *Handbook of Labor Economics*, Elsevier, 2011, pp. 1237 – 1313.

Musterd, Sako and Rinus Deurloo, "Ethnic segregation and the role of public housing in Amsterdam," *Tijdschrift voor economische en sociale geografie*, 1997, *88*, 158–168.

Petrin, Amil and Kenneth Train, "A Control Function Approach to Endogeneity in Consumer Choice Models," *Journal of Marketing Research*, 2010, *47* (1), 3–13.

Pinçon, Michel, *Les HLM: structure sociale de la population logée. Agglomération de Paris - 1968*, Centre de sociologie urbaine, Paris, 1976.

Pischke, Jorn-Steffen and Johannes Velling, "Employment Effects of Immigration to Germany: An Analysis Based on Local Labor Markets," *The Review of Economics and Statistics*, 1997, *79* (4), 594–604.

Priemus, Hugo and Frans Dieleman, "Social housing policy in the European Union: past, present and perspectives," *Urban Studies*, 2002, *39* (2), 191–200.

Rutter, Jil and Maria Latorre, *Social housing allocation and immigrant communities*, Equality and Human Rights Commission, 2008.

Saiz, Albert, "The Geographic Determinants of Housing Supply*," *Quarterly Journal of Economics*, 2010, *125* (3), 1253–1296.

Saks, Raven E., "Job creation and housing construction: Constraints on metropolitan area employment growth," *Journal of Urban Economics*, July 2008, *64* (1), 178–195.

Schor, Ralph, *Histoire de l'immigration en France de la fin du XIXe siècle a nos jours*, Armand Colin, 1996.

Selod, Harris, *La mixité économique et sociale*, Vol. Villes et Économie, La Documentation Française,

Stébé, Jean-Marc, *Le logement social en France*, PUF, 2007.

Subra, Philippe, "Heurs et malheurs d'une loi antiségrégation : les enjeux géopolitiques de la loi Solidarité et renouvellement urbain (SRU)," *Hérodote*, 2006, *122* (3), 138–171.

Tavan, Chloé, Agnès Dugué, Jean-Paul Caille, and Maryline Bèque, *Les immigrés en France*, INSEE, 2005.

Train, Kenneth, *Discrete Choice Methods with Simulation*, Cambridge University Press, 2003.

Verdugo, Gregory, "Public Housing and Residential Segregation of Immigrants in France, 1968-1999," *Population-E*, 2011, (1).

_ , "Fragmentation urbaine et chocs économiques : deux déterminants de l'offre de logement sociaux en France," *Économie et Statistiques*, Forthcoming.

Voldman, Danièle, *La reconstruction des villes françaises de 1940 à 1954: histoire d'une politique*, L'Harmattan Edition, 1997.

Weil, Patrick, *La République et sa diversité: Immigration, intégration, discriminations*, Seuil, 2005.