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Immigration and the public–private school choice[☆]

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ABSTRACT

This paper empirically analyzes the effects of immigration on the schooling decisions of natives. We employ household-level data for Spain for years 2000–2015, a period characterized by high economic growth and large immigration that was halted by a long and severe recession. Our estimates reveal that increases in immigrant density at the school level triggered an important *native flight* from tuition-free, public schools toward private ones. We also find strong evidence of *cream-skimming* as more educated native households are the most likely to switch to private schools in response to immigration. Furthermore, we find that immigration leads to higher student–teacher ratios in public schools. We conclude that our results are consistent with the predictions of a political-economy model of school choice.

1. Introduction

Public education is a fundamental engine for human capital accumulation, with important consequences for income inequality and upward mobility (Glomm and Ravikumar, 1992), (Galor and Zeira, 1993), or (Fernandez and Rogerson, 1996). This is particularly so for first and second-generation immigrants, who disproportionately attend public schools and for whom socio-economic assimilation depends greatly on the quality of education they receive (Dustmann et al., 2012). However, a large concentration of immigrants in public schools may decrease the support for funding among natives and lead to a deterioration of the public education system (Epple and Romano, 1996).

This paper empirically estimates the effects of immigration on the education system of the receiving country, with an emphasis on the consequences for the public–private school choice of natives. We employ data for Spain over the period 2000–2015, a period characterized by an important economic expansion that was accompanied by a large immigration wave, and ended with a long and severe recession. To identify families with school-age children enrolled in private schools, we use information on tuition expenditures from the Spanish Family Expenditures Survey. These data also allow us to investigate changes in other consumption categories that may be triggered by changes in schooling expenditures. Lastly, we also examine the effect of migration on student–teacher ratios, and discuss mechanisms that can explain our results.

The important demographic and economic changes that occurred in Spain during the last decade offer an excellent scenario to investigate the impact of migration on schooling decisions. Between 1995 and 2007, Spain experienced a period of fast-paced economic growth. During these years, the employment to population ratio increased by 14 percentage points and real household income increased by more than 50% (Fig. 1). In contrast, between 2007 and 2013, Spain experienced a very severe recession, aggravated by drastic austerity policies.

Not surprisingly, these changes in economic conditions had large implications for migration flows. Between 2000 and 2008, the foreign-

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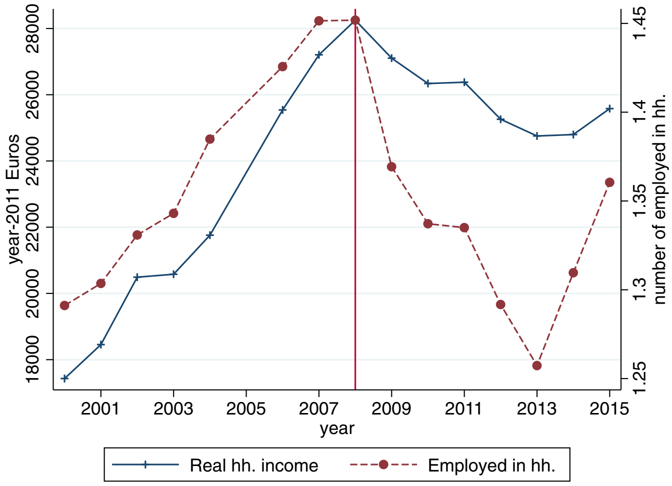

Fig. 1. Annual Household Income (in real terms) and Number of Employed in Household. *Notes:* Family Expenditure Survey.

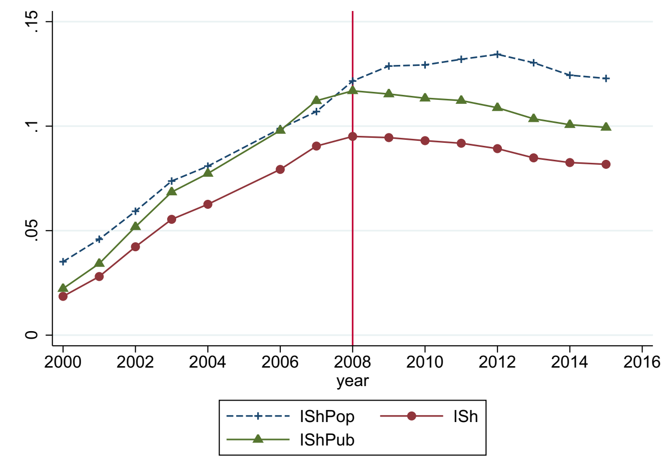

Fig. 2. Immigrant shares in the population and in enrollment. *Notes: IShPop* is the foreign-born share (for working-age population) based on data from the Population Registry. *ISh* is the share of foreign students enrolled in primary and secondary schools, including both private and public. *IShPub* refers to the share of foreign students enrolled only in public schools, including both primary and secondary. The latter two series are based on administrative enrollment data (Spanish Ministry of Education).

born share in the (working-age) population increased from 4% to 12%, as illustrated by the dashed line in Fig. 2. The inflows of workers were accompanied by a large increase in the number of immigrant children in schools. As shown in the same Figure (solid, red line), the share of immigrant students in primary and secondary schools increased in parallel to the immigrant share in the working-age population, from less than 3%in year 2000 to almost 10% in 2008. The Figure (solid, green line) also shows that immigrant children were over-represented in public schools, where their share in enrollment increased by almost 9 percentage points between years 2000 and 2008. Fig. 2 has an additional implication that plays an important role in our analysis. As soon as the economic downturn began in 2008 (Fig. 1), the immigrant share in enrollment started decreasing. In contrast, the immigrant share in the population continued rising and only peaked four years later. As we discuss in detail later, this is due to the difficulties of population registry data to accurately measure the immigrant population in periods of net outflows. A strength of our study is that we will rely more heavily on measures of immigrant density based on school enrollment data.

It is interesting to dig deeper into the impact of immigration on the Spanish education system. To do so we consider enrollment in primary and secondary schooling separately in Fig. 3. During the period 2000–2008, as immigrant households where arriving in the country, the share of immigrant students rose rapidly both in primary and secondary schools. The increase was substantial in primary schools, where in the course of these 8 years, the immigrant share rose from around 2% to about 9%. This was an important demographic shock to the school system, particularly for public schools where immigrant children are substantially over-represented. Interestingly, the trajectories for the immigrant shares across education levels differed from 2008 onward. Between 2008 and 2012, the share at primary schools declined by about 10%, while it kept rising in secondary schools up until coming to a halt in 2012. This divergence in trends reflects the vanishing of new immigration flows due to the economic downturn, at the same time as the children of the previously arrived immigrants that remained in Spain progressed through the education system. Because of the differences in the timing and magnitude of the inflows of immigrant children into primary and secondary schools, we conduct the analysis separately by education level (primary and secondary) and by time period (2000–2007 and 2008–2015). This approach can strengthen the identification of the causal effect of immigration, as exemplified by the increases in immigrant density in secondary education during the 2008–2012 period, already characterized by a severe economic contraction.

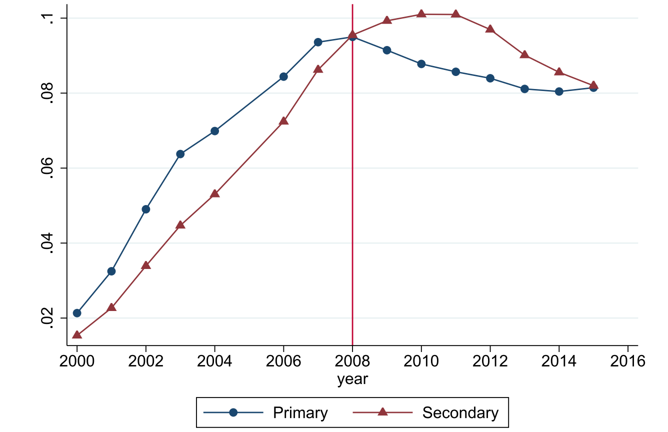

Fig. 3. Immigrant share in enrollment (primary and secondary schooling). *Notes:* The figure reports the share of foreign students in enrollment separately in primary and secondary schools. These enrollment figures include students at both public and private schools. It effectively decomposes series *ISh* in Fig. 2. Secondary schooling includes middle school (known as ESO), high school (known as 'bachillerato'), and vocational training (known as "formacion profesional"). The data are based on administrative enrollment data (Spanish Ministry of Education).

As has been widely recognized, the co-movement of immigration flows and the economic cycle poses a challenge to identify the role of migration on the schooling choices of households. We address this problem by using detailed household-level data on employment and income, combined with regional variation in immigration flows. We account for the classic endogeneity problem of the location choices of immigrants by adopting an instrumental-variables approach based on ethnic networks (Card, 2001). However, we depart from the usual renditions of the instrument by focusing on predicting changes in the immigrant share in *enrollment* at different education levels, which turns out to be much more informative than predicting immigrant shares in the *population*, as is usually done in the literature.

Our analysis delivers several interesting findings. Our two-stage least-squares estimates show that increases in immigrant density in public schools led to an increase in household educational expenditures, largely driven by *native flight* toward private schools. The intensity of this response varied across education levels (primary and secondary) and as a function of the education level of the household head. In primary schools, immigration led to a shift toward private schools both during the economic expansion (2000–2007) and contraction (2008–

2015), but only among college-educated household heads. The same happened in secondary schools during the first period. However, our estimates suggest that in the period 2008–2015 immigration displaced also households with lower educational attainment by a similar magnitude. The effects of immigration that we uncover are large, about half of the size of the effect of household income on private–public school choices. We also provide evidence of increases in student–teacher ratios in public schools, which suggest that immigration may have led to crowding and possibly a deterioration of the quality in public schools.

Our work is related to several strands of literature. Previous studies have analyzed empirically the question of the effects of immigration on the schooling decisions of natives, and found evidence of a displacement of natives away from tuition-free public schools and toward private schools. Betts and Fairlie (2003) documented that increases in the share of immigrants in a metropolitan area were associated to increases in the probability to attend private school among native households in California. Complementing the previous study, Cascio and Lewis (2012) found evidence of native migration to nearby school districts in response to inflows of Hispanic students with low English proficiency. Several authors have also found evidence of native displacement or immigrant segregation in schools in Europe. Kristen (2008) provides evidence for Germany, Gerdes (2013) for Denmark, and Schneeweis (2015) for Austria. From a theoretical perspective, Albornoz-Crespo et al. (2017) build a model that endogenizes the quality of education through the effort exerted by native and immigrant students.

This paper is also related to the wider education literature on the role of parental education and socio-economic background as a determinant of children's outcomes. These studies emphasize that more educated parents may be better informed regarding schooling options and may also put a stronger emphasis on the education of their children. This suggests that immigration may trigger heterogeneous reactions on natives of different socio-economic status, as in the analysis of cream-skimming in terms of school choices in Altonji et al. (2015). Naturally, the role of parental background is also a fundamental determinant of the educational outcomes of first and second-generation immigrants. In their comparative analysis of OECD countries, Dustmann et al. (2012) stress the important role played by the quality of the schools attended by the children of immigrants.

Our study also complements the work by Hunt (2016) and Llull (2017) on the effects of immigration on the educational attainment of natives. While they focus on the consequences for native high-school completion rates and college attainment, we focus on educational investments in compulsory education. Our use of household expenditures data to identify school choices was inspired by Arellano and Zamarro (2007) who study the determinants of the public-private school choice in Spain in year 1990, and by Dustmann et al. (2017) who use consumption expenditure data on their analysis of the effects of immigrant legal status. Finally, Anghel and Cabrales (2014) document the role of school type and parental background in explaining performance in standardized exams in Spain.

The structure of the paper is as follows. Section 2 presents a concise description of the schooling system in Spain. Section 3 discusses our empirical strategy. Section 4 describes our main data sources and provides summary statistics. Section 5 presents our main estimation results and a discussion of their quantitative implications. Section 6 provides robustness checks, followed by a discussion of the mechanisms behind our empirical findings in Section 7. Section 8 concludes. Tables and Figures are gathered at the end of the paper.

2. Background

2.1. The spanish school system

Since the early 2000s pre-university education in Spain can be described as follows. Compulsory schooling is composed of two stages: primary (elementary) school, consisting of six grades, and four years of secondary schooling (known as E.S.O, or Compulsory Secondary Schooling in its Spanish acronym). Compulsory schooling starts at age 6, however it is very common to be enrolled from age 3 – over 96% of 3-year olds were enrolled in school in academic year 2010–2011 – in order to guarantee admission to higher grades at the same school. Schooling is compulsory up to age 16 but we also include in our sample households with 17–18 year-old children because the Family Expenditures Survey did not differentiate educational expenditures in compulsory and non-compulsory secondary education in years 2000–2004.

Primary and secondary education in public schools is free of tuition and is fully financed by taxes.[1] Besides private schools that do not receive government funding, the Spanish education system is characterized by the widespread use of publicly subsidized private schools, known as "concerted" schools, which account for about one third of all students. Concerted schools were introduced in 1985 to accommodate the increasing demand for education that resulted from the baby boom, and the majority are Catholic. In exchange for government funding that supposedly covers the school's whole salary bill, concerted schools agree to the curriculum and admission policies of public schools.

While, in theory concerted schools are not allowed to charge for tuition, in practice there are quasi-compulsory payments required from parents in terms of donations to the parents' association, building maintenance, or extracurricular activities. According to a 2012 study by the Association of Spanish Consumers (OCU, 2012), over 90% of concerted schools require payments that are perceived by households as compulsory. Nationally, the average annual payment reported in this study was 501 Euros, roughly 2% of the average total household income over the period of analysis. However, there are important regional differences, with mean values in 2012 ranging from 105 Euros to about 1000 Euros. Part of these expenses are incurred at the time of registration and the rest are paid monthly. Out-of-pocket disbursements at fully private schools are substantially higher, often amounting to several thousand Euros per year. As we show later, the tuition expenditures in our data set are highly consistent with the estimates reported in this study. In addition, a 2007 supplement to the Spanish Family Expenditure Survey (FES) reported average annual household expenditures per student disaggregated by type of school. Focusing on expenditures in tuition (upfront or as monthly fees) and extracurricular activities taking place within the schools, the average primary-school expenses for students in public, concerted, and private schools were, respectively, 46 Euros, 341 Euros, and 1765 Euros. The corresponding figures for compulsory secondary education were 10 Euros, 260 Euros, and 2223 Euros. Thus, there are systematic differences in tuition expenditures by school type, which motivates our strategy to identify the use of public versus private or concerted schools.

Previous studies have reported important differences between public and non-public schools in Spain. For example, Arellano and Zamarro (2007) and Trillo del Pozo et al. (2006) report evidence that points to higher mean test scores for students at private schools, followed by concerted, and then public schools. However, it is less clear whether the source of this differences is due to sorting in ability and family background, or to the value-added provided by the school (Calero and Escardibul, 2007) and Anghel and Cabrales (2014).

2.2. Immigration and schools

Between year 2000 and 2010 the population of students with foreign nationality increased by a factor of 5.4, from 141,916 to over 770,384 students, and accounted for about 60% of the overall growth in enroll-

[1] Not surprisingly, public spending in education fluctuates with the economic cycle. The share of public spending in education (including universities) over GDP increased from 4.27% in 2001 to 4.98% in 2009, and fell with the recession (4.74% in 2011). Public spending in universities accounts for about 20% of total public spending in education.

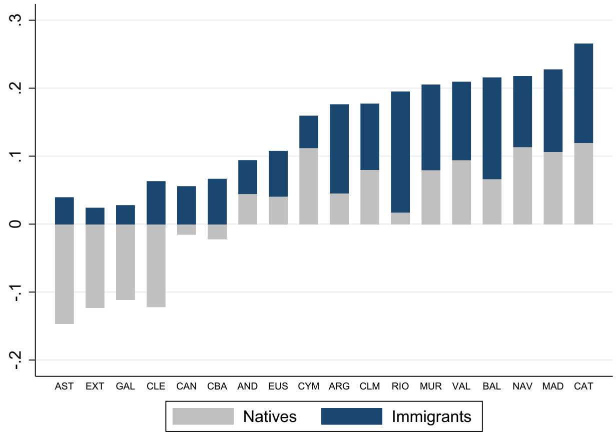

Fig. 4. Growth in enrollment by region, 2000–2010. *Notes:* The data correspond to academic years 2000–2001 and 2010–2011. We report the change in the native and immigrant student population (for all pre-university levels) over the total (native plus immigrant) value in 2000, combining public and private schools. Immigrants are defined as children with foreign-nationality. Thus second-generation immigrants or Spanish children with an immigrant parent are considered natives. Each bar corresponds to one autonomous community (sorted in increasing order): Asturias (AST), Extremadura (EXT), Galicia (GAL), Castilla y Leon (CLE), Canarias (CAN), Cantabria (CBA), AndalucÃa (AND), Euskadi (EUS), Castilla La Mancha (CYM), Aragon (ARG), Ceuta and Melilla (CLM), Rioja (RIO), Murcia (MUR), Comunitat Valenciana (VAL), Balearic Islands (BAL), Navarra (NAV), Madrid (MAD), Catalunya (CAT). The data are based on administrative enrollment data (Spanish Ministry of Education).

ment over the period.[2] Immigration is by far the main demographic factor behind the regional variation in enrollment levels since year 2000. Fig. 4 reports the 2000–2010 changes in the enrollment of children with and without Spanish nationality for each of the 17 Spanish regions (autonomous communities). While overall enrollment rose by over 20% in some regions (Catalonia, Madrid and Navarra), it fell by over 10% in others (Asturias, Extremadura and Galicia). Importantly, in all regions with a net increase in enrollment, the main driving force was the growth in students with foreign nationality. These wide regional disparities will be at the core of our empirical strategy.

The impact of immigration on public and non-public schools has been very uneven. According to the Spanish Ministry of Education, in year 2000 the shares of foreign students in public and private schools were similar (2.7% and 1.2%, respectively). By 2008 the corresponding figures were 11.9% and 4.8%. That is, a 9 percentage-point increase in public schools compared to barely 3.5 in private schools. It is also worth pointing out that the immigrant population in Spain is very diverse in terms of origin. In 2010 the breakdown of the foreign student population by origin is as follows: 40% originated from South and Central America, 29% from the rest of Europe, 23% from Africa, and about 6% from Asia and Oceania. The vast majority of students with foreign nationality are enrolled in public schools (82%), compared to 14% in concerted schools, and only 4% in fully private schools. In comparison, the breakdown for the overall student population, including natives, is approximately 68%, 27%, and 5%, respectively. The larger use of public schools among immigrant households may be due to the larger out-of-pocket household expenses documented above, although it is also possible that the emphasis on Catholic education plays a role as well. de la Rica and Ortega (2009) report that 11% of the foreign-born population in Spain (in year 2008) originated in Morocco, and there is also a sizable number of immigrants from other Muslim countries.

To accommodate the net increase in the demand for education, the supply of schools and teachers was expanded importantly over the 2001–2010 period. The numbers of public and private schools increased by 20% and 43%, respectively, and the numbers of teachers by 24% and 30%.[3] Later in the paper we analyze whether immigration had an effect on student–teacher ratios.

3. Empirical approach

In our analysis we employ annual cross-sections of household-level data restricted to a sample of households with school-age children (age 3–18). Our main goal is to identify the effect of immigration on the public-private school choice. In our first empirical model the dependent variable is the inverse hyperbolic sine of education (tuition) expenditures (*edux*) *per child* for household *i* in region *r* and year *t*.[4] The right-hand-side contains year (α_t) and region (λ_r) fixed effects, a vector of household-level characteristics ($\mathbf{X}_{i,r,t}$), a measure of the immigrant share in region *r* and year *t*, ($ISh_{r,t}$), and a disturbance term ($u_{i,r,t}$). Specifically,

$$edux_{i,r,t} = \alpha_t + \lambda_r + \mathbf{X}'_{i,r,t}\beta + \gamma ISh_{r,t} + u_{i,r,t}. \tag{1}$$

Importantly, vector **X** contains the log of real household income and the number of employed household members, together with the educational attainment of the household head and the number of school-age children in the household.

In order to isolate responses along the extensive margin, defined as the discrete choice between public and private schools, we consider a linear-probability model analogous to Eq. (1) but with a private-school indicator as dependent variable:

$$priv_{i,r,t} = \alpha_t + \lambda_r + \mathbf{X}'_{i,r,t}\beta + \gamma ISh_{r,t} + u_{i,r,t}. \tag{2}$$

Initially, we abstract from endogenity concerns in Eqs. (1) and (2) and focus on examining the association between school choices and our variables of interest. By virtue of the regional dummies in the empirical models, the effect of immigration is identified by exploiting the cross-regional changes in immigrant densities over time. In our main specifications, we measure the immigrant share in terms of enrollment, rather than population.

As in all empirical studies aimed at estimating the effects of immigration on the basis of spatial (cross-region) correlations, there is some concern that OLS estimates of Eqs. (1) and (2) may suffer from endogeneity bias. Typically, one worries that households in regions experiencing positive income shocks will tend to increase spending in education along with other spending categories. Since such regions are likely to attract more immigrants, one would expect an *upward* bias in the estimated coefficient for the immigrant share. However, the bias could also go in the other direction. Regions that experienced positive income shocks may have attracted more immigrants and, simultaneously, invested to increase the quality of *public* education. Improvements to public schools, relative to private ones, would amount to a *negative shock* in Eqs. (1) and (2) that could bias *downward* the OLS estimate of γ.

To address these issues, we employ a standard instrumental-variables strategy based on building predictors of the regional foreign-born share that can be considered exogenous to unobserved local demand shocks for private schooling. We mainly rely on an extension of the ethnic networks instrument developed by Altonji and Card (1991) and Card (2001), which has been shown to be useful in the case of Spain (Farre et al., 2011) and Gonzalez and Ortega (2013). However, a novelty of our approach is that we focus on predicting immigrant shares in *enrollment* rather than in the *overall population*. Immigrant enrollment is

[2] This figure underestimates the impact of immigration because it fails to account for students that have double nationality and second-generation immigrant children.

[3] According to the Spanish Ministry of Education, the number of public schools increased by 3052 and the number of private schools by 2411. In terms of teachers, the increases were 96,690 and 42,004 in public and private schools, respectively.

[4] Our sample contains a large number of observations with 0 educational expenditures. Since the logarithmic function is not defined at 0 the standard procedure is to add a constant or drop the zeros. A better alternative is to transform the variable using the inverse hyperbolic sine, $ihs(y) = ln(y + (y^2 + 1)^{1/2})$. This transformation can be interpreted as the log of expenditures but has the advantage that it is well defined at zero.

more accurately measured than the counts in the population. In addition, immigration shares in terms of enrolled students are a more relevant measure of immigrant density in the context of our analysis.[5]

4. Data

Our main dataset is the Family Expenditures Survey (FES, or "Encuesta de Presupuestos Familiares" in Spanish), for years 2000–2015.[6] The FES has a sample size of approximately 24,000 households per year, and is the main data source to quantify private aggregate household consumption in the national accounts and to compute the CPI. In addition this survey is commonly used by social scientists to study many issues, including education, housing, nutrition and healthcare use (Arellano and Zamarro, 2007) and Gonzalez (2013). Since 2006 it is possible to identify households who stay in the survey for two consecutive years, providing a short longitudinal dimension that we will also exploit in our analysis. One limitation of our data is that in period 2000–2004 the FES did not ask respondents about their nationality or country of birth. Thus we are not able to exclude immigrant households from the sample during this period.

The sample for our analysis are households with school-age children. Specifically, we consider all households with at least one child in age group 2–18. Schooling is not compulsory until age 6 but, as discussed in section Section 2, enrollment rates at age 3 are over 95%. Because some children are 2 years old when they enter school, turning 3 during the school year, we include households with 2 year-old in our sample. This is not a problem because our educational expenditure variable excludes payments to daycare and kindergarten programs. At the other end of the age distribution, ideally, we would like to focus on compulsory secondary schooling, which ends at age 16. However, in years 2000–2004 the data set does not allow us to separate tuition payments for the compulsory and non-compulsory segments of secondary education, typically involving 15–16 and 17–18 year-old, respectively. Because very few students are in pre-school at age 2, we will refer to our sample as containing all households with children *age 3–18*. Throughout the paper, the subsample of households with children in primary school also contains pre-schoolers since it is based on households with children *age 3–11*. Likewise, the subsample of households with children in secondary school includes all households with children *age 12–18*.

Following Arellano and Zamarro (2007), we use educational expenditures to identify school choices. For our purposes it does not matter whether students attend a concerted or a fully private school. We view concerted schools simply as private schools with low tuition. Our *educational expenditure* variable *(edux)* contains annual tuition, registration, and related expenses, such as extracurricular activities offered *within* the school, at constant 2011 prices. These expenses may have been incurred once, typically at the beginning of the year, or periodically (e.g. in monthly installments). School lunches and extracurricular activities that take place outside the school grounds are not included in this variable. Our key schooling expenditure measure is normalized by the number of school-age children in the household. Because our data measures separately educational expenditures for primary and secondary schools, when we restrict the estimation to one of these educational levels, variable *edux* divides the educational expenditures specific to that level by the number of children in the corresponding age group.

In our data, the average real educational expenditure per school-age child (*edux*) is 280 Euros at 2011 prices. However, the vast majority of households report zero tuition (public school users). Approximately 67% of the sample spent less than 10 Euros on tuition, despite having

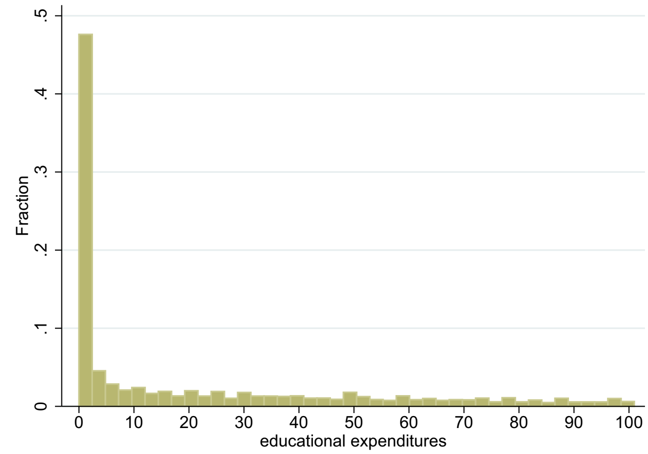

Fig. 5. Histogram tuition. *Notes:* The figure displays the fraction of households with positive educational expenditures (primary and secondary) per school-age child below 100 over the period 2000–2015. Source: Family Expenditure Survey.

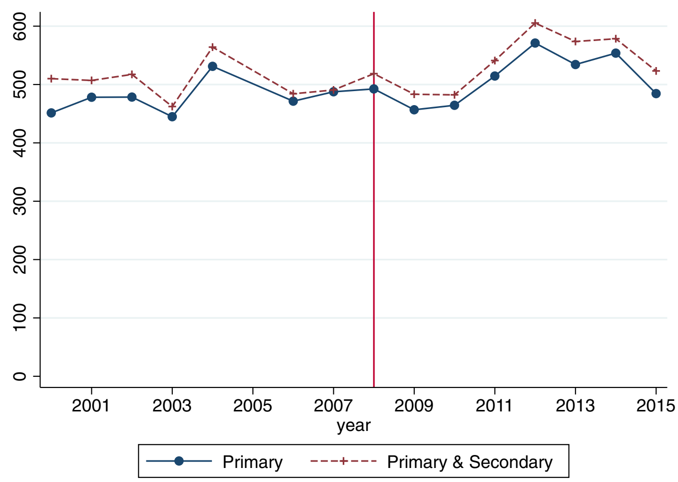

Fig. 6. Mean educational spending per child, conditional on private use. Source: Family Expenditure Survey. Educational spending per child (*(edux)*) contains annual tuition, registration, and related expenses, such as extracurricular activities offered *within* the school, at constant 2011 prices. These expenses may have been incurred once, typically at the beginning of the year, or periodically (e.g. in monthly installments). School lunches and extracurricular activities that take place outside the school grounds are not included in this variable.

school-age children. About 8% spent between 10 and 100 Euros, 21% spent between 100 and 2000 Euros, and 3.6% of the sample spent over 2000 Euros in tuition. We define a household as a *private school user* if educational expenditures per school-age child in the household are above 30 Euros, measured at 2011 constant prices. This threshold delivers a share of households using private schools of approximately 30%, which is in line with the average share of private school users in the administrative enrollment data over the period 2000–2015. Fig. 5 displays the histogram for positive educational expenditures below 100 Euros. The lack of mass points beyond the first bin suggests that the analysis of the extensive-margin (public versus private school choice) will be robust to a wide range of tuition thresholds. As an important robustness check, Section 6 considers year-specific thresholds that are calibrated to match the annual shares of students in private schools in the administrative enrollment data. When we condition on private school use (on the basis of the 30-Euro threshold), the average annual tuition in our data is around 500 Euros (Fig. 6), which is in line with the estimates by independent consumer agencies (OCU, 2012).

[5] Throughout the paper we report standard errors clustered at the region (autonomous community) level. Since there are only 17 clusters, we follow Cameron et al. (2011) and use higher than usual critical values in the significance tests pertaining to our main specifications.

[6] Up until 2004 the survey was administered quarterly. It was not administered in 2005 and since 2006 it has been conducted annually.

Table 1
Summary statistics, 2000–2015 .

Variable	Obs	Mean	Std. Dev.	Min	Max
Year	83,995	2007.843	4.701	2000	2015
Region	83,995	8.120	4.757	1	18
Children 3–18	83,995	1.525	0.674	1	10
HH. HS grad.	83,995	0.212	0.409	0	1
HH. Co. grad.	83,995	0.295	0.456	0	1
HH. Immig.	69,033	0.185	0.338	0	1
H. income	83,995	26174.66	15257.62	11.551	404734.8
H. employed	83,995	1.349	0.637	0	2
Edux	83,995	279.897	830.182	0	26,171
Edux Prim	44,018	285.079	823.968	0	21160.38
Edux Sec	44,697	246.822	798.870	0	26170.7
Private	83,995	0.293	0.455	0	1
Private Prim	44,018	0.295	0.456	0	1
Private Sec	44,697	0.254	0.435	0	1
IShPop	83,995	0.110	0.056	0.011	0.243
FB Europe / FB	83,995	0.410	0.118	0.034	0.684
FB Africa / FB	83,995	0.174	0.085	0.026	0.952
FB America / FB	83,995	0.369	0.116	0.006	0.588
FB Asia-Ocea. / FB	83,995	0.047	0.026	0	0.112
ISh	83,995	0.076	0.04	0.006	0.165
IShPub	83,995	0.095	0.054	0.007	0.198
IShPubPrim	83,995	0.099	0.054	0.008	0.215
IShPubSec	83,995	0.090	0.055	0.006	0.208

Notes: Our sample contains only households with children age 3–18. Regions are defined as autonomous communities. We report weighted means, all years pooled. HH. immig. is an indicator for a foreign-born household head. This variable is only defined for survey years 2006–2015. Children 3–18 is the number of children in that age group. H. income is the real annual household income. H. employed refers to the number of employed individuals in the household. *IShPop* refers to foreign-born share in the population. The foreign-born (FB) population is partitioned by continent of origin. We report the shares, such as FB Europe / FB is the share of foreign-born individuals born in other European countries. *IShPub* is the share of immigrant students in (primary or secondary) public schools and *IShPub* reports the analogous share but restricted to public schools only. *IShPubPrim* is the immigrant share in public primary schools, and *IShPubSec* in public secondary. *Edux* is defined for households with children in age group 3–18 and is the ratio of real educational expenditures for that household divided by the number of children age 3–18. *Edux Prim* is analogous but defined only for households with children age 3–11 and *Edux Sec* for households with children age 12–18. *Private* is an indicator for households with annual tuition expenditures per child above 30 Euros. Likewise, *Private Prim* (*Private Sec*) is an indicator for households with children age 3–11 (age 12–18) with positive annual tuition expenditures per child above the 30-Euro threshold.

Let us now provide some descriptive statistics. As noted earlier, our sample consists of all households with children age 3–18 in the FES for years 2000–2015. It contains 83,995 households and the mean number of children age 3–18 per household is 1.53, ranging from 1 to 10 (Table 1). On average 21% of the household heads are high-school graduates and 29% have a college degree. Annual household income is 26,175 Euros (at constant 2011 prices) and the mean number of employed individuals in the household is 1.35, though the latter figures fluctuated widely with the economic cycle. Real educational expenditures per child (tuition, registration and in-school extracurricular activities) average 280 Euros annually (at 2011 prices), but keep in mind that the vast majority of households have zero or negligible expenses. On average 29% of households reported tuition expenditures above the 30 Euros threshold, which we identify as private school users. Average spending per student in primary education averages 285 Euros annually (in real terms), and the analogous figure for secondary education is only 40 Euros lower. The share of households with children enrolled at private schools is also somewhat lower in secondary (25.4%) than in primary education (29.5%).

Most migration studies relying on the spatial correlation approach measure immigrant density as the foreign-born share in the population. We can construct this measure from the Spanish Population Registry.[7] However, we are concerned that this measure does not accurately capture the sharp changes in migration flows occurred over the period. Specifically, the Population Registry may have been slow in reflecting the change in net migration when it turned negative because of the onset of the Great Recession. The reason is that upon arrival to the country immigrants have strong incentives to register in order to gain access to health care and education, but there is no incentive to de-register upon departure. As a result, there may be a significant delay until reductions in the immigrant population are accurately recorded (de la Rica et al., 2013).

We propose a novel measure of immigrant density based on the Spanish Education Enrollment Registry. We employ these data to compute the share of immigrant students (defined as not being Spanish citizens) enrolled in school (*ISh*). We also define the immigrant share on the basis of enrollment in public schools only (*IShPub*), which is actually our preferred measure of immigrant density in our regression analysis. Table 1 reports the average *IShPub* to be 9.5% over the period under analysis. Fig. 2 plots the immigrant shares in the population (*IShPop*), in public schools (*IShPub*) and in terms of total enrollment (*ISh*). In year 2000 the immigrant share in the population (*IShPop*) was 4%, corresponding to the dashed line in the figure. It increased at a rapid pace until the 2008 economic downturn when it reached 13%. As we suspected, this variable keeps rising and peaks in 2012. In contrast, the share of immigrants enrolled at public schools (*IShPub*) and in total enrollment (*ISh*) peak in 2008 and fall monotonically for the next five years, displaying much higher sensitivity to changes in general economic conditions. The figure also shows that the foreign share in enrollment is higher for public schools, confirming the larger concentration of immigrants in these schools. This Figure strongly suggests that the immigrant shares based on enrollment data are more accurate measures than the share based on the population registry.

It is also interesting to compare the characteristics of households using private and public schools. Table B.1 in the Appendix provides the means for a number of variables conditional on school choice. As expected, private users are characterized by higher household income (36%) and higher educational attainment of the household head, with the share of college graduates among households using private schools at 44%, compared to only 23% among public school users.

Let us now describe the evolution of the main variables over time. Table 2 reports population-weighted means at (roughly) 5-year intervals for the key variables. Several interesting trends stand out. First, there is a sustained reduction in family size in terms of the number of school-age children for the period 2000–2010. This is more clearly seen for the age group 3–11. The mean number of children in this age group for our sample fell from 1.48 in year 2000 to 1.43 in 2010, but has recovered in the later years. The Table also shows substantial skill upgrading, with the share of high-school graduates and college graduates increasing by 5 and 20 percentage points, respectively, throughout the period 2000–2015. The share of households with a foreign household head also increased substantially from 14% in 2006 to 19% in 2015. The data also illustrate the effects of the business cycle, with household income and employment falling between 2006 and 2010 and only partially recovering by 2015 (as seen already in Fig. 1). Table 2 also reports educational expenditures per child and the share of households using private schools (on the basis of the 30-Euro threshold). Between years 2000 and 2015, the share of private users in our sample increased from 0.26 to 0.33. The increase was larger in primary schooling (from 0.24 to 0.34) than in secondary (from 0.23 to 0.29). Mean educational expenditures (in real terms) also increased over the period by more than 50%. However, after conditioning on private school use, the increase is moderate (from 752 to 962 Euros in primary education and from 811 to 891 Euros in secondary education)

[7] Previous studies using this standard measure for the Spanish case are Farre et al. (2011) and Gonzalez and Ortega (2013)

Table 2
Population-weighted means (selected years).

Year	2000	2006	2010	2015
Households count	4,434,603	4,468,639	4,789,089	4,828,842
Children 3–11	1.48	1.45	1.43	1.48
Children 12–18	1.28	1.24	1.24	1.24
HH. HS. Grad.	0.18	0.19	0.22	0.23
HH. Co. Grad.	0.19	0.29	0.32	0.39
Spouse HS. Grad.		0.21	0.21	0.22
Spouse Co. Grad.		0.24	0.29	0.34
HH. Immig.		0.14	0.19	0.19
H. income	23,700	28,619	27,181	24,784
H. employed	1.29	1.43	1.34	1.36
IShPub	0.02	0.10	0.11	0.10
IShPubPrim	0.03	0.11	0.11	0.10
IShPubSec	0.02	0.09	0.12	0.09
Private	0.26	0.28	0.31	0.33
Edux	206	284	302	328
Conditional private	798	1005	963	982
Private Prim	0.24	0.29	0.32	0.34
Edux Prim	184	299	317	327
Conditional private	752	1046	1000	962
Private Sec	0.23	0.24	0.28	0.29
Edux Sec	182	205	236	257
Conditional private	811	840	851	891

Notes: In year 2005 the Household Expenditure Survey was not conducted. *IShPop* refers to foreign-born share in the population (Population Registry). *IShPub* refers to the share of foreign students in public primary and secondary schools (Administrative Enrollment Data). *IShPubPrim* is analogous but refers only to primary schools and *IShPubSec* to secondary schools. HS. Grad. is the fraction of households with a household that graduated from high school. HS. Co. Grad. refers to the share of household heads that graduated from college. Analogous variables are also defined for spouses of the household head. Children 3–11 and children 12–18 is the average number of children in the respective age groups. Private is the share of households with annual tuition expenditures per child higher than 30 Euros. *Edux* is the average tuition paid by households divided by the number of children age 3–18. We also report these expenditures conditioning on a minimum of 30 Euros. *Private Prim* (*Private Sec*) is the share of households with children age 3–11 (age 12–18) with annual tuition expenditures higher than 30 Euros. *Edux Prim* (*Edux Sec*) per child corresponds to tuition expenditures divided by the number of children 3–11 (age 12–18) in the household. We also report the average tuition per child conditional on annual expenditures higher than 30 Euros (*Conditional private*).

5. Estimates

5.1. OLS and Tobit estimates

We begin our empirical analysis by examining Table 3, which contains OLS estimates for tuition expenditures (*edux*) and for the probability to use private schools (*private*). For the former we also report the marginal effects of a Tobit model (averaged over the whole sample and for households with positive tuition expenditures only) in order to account for the large percentage of households with zero educational expenditures. As discussed in the introduction, the presence of immigrants at primary and secondary schools did not follow the same pattern over the period under analysis. Given the diverging patterns in the immigrant concentration across education levels, we examine the effects in primary and secondary education, jointly and separately. The point estimates in the table correspond to the coefficient on the immigrant share, in Eqs. (1) and (2), measured on the basis of enrollment in public schools. The estimates in columns 1–3 are for the sample of all households with school-age children (age 3–18). Accordingly, the immigrant share in public schools is based on primary and secondary schooling pooled together (*IShPub*). The estimates suggest a positive association between immigration and educational expenditures. However, we only obtain statistically significant estimates for the period 2000–2007. For years 2008–2015, the point estimate is also positive, but smaller and more imprecisely estimated. The same pattern is observed for the Tobit estimates.

Columns 4–6 restrict the analysis to households with children in primary school (age 3–11), and measure immigrant density on the basis of enrollment in primary public schools (*IShPubPrim*). The point estimates are substantially higher when considering the whole period or the economic expansion. In fact we now obtain positive and statistically significant coefficients in both of those cases for OLS and Tobit estimates. In contrast, the point estimates for the period 2008–2015 are now much smaller than when considering both education levels jointly. Columns 7–9 turn now to secondary education and measure immigrant density in terms of enrollment in secondary public schools (*IShPubSec*). In this case, we find larger (and statistically significant) coefficients for the period 2008–2012 than for the first period.

The bottom panel reports the estimates for the model where the dependent variable is an indicator for private school use. The pattern of the results is the same as before: we find evidence of a significant positive association between immigrant density in public schools and increases in private-school use in primary education during the period 2000–2007 and in secondary education for years 2008–2015. This is consistent with the differences in the timing of the increase in immigrant density at the different levels of education. As illustrated by Fig. 3, the largest inflows of immigrant children into primary schools, relative to total enrollment, took place in period 2000–2007, whereas the largest inflows into secondary schools took place during the first half of period 2008–2015. Let us now address the potential endogeneity bias that may affect these estimates due to immigrants' location choice.

5.2. Instrumental-variables estimates

Our regression models are potentially subject to the classical endogeneity problem arising from the endogenous location of immigrants. To address this point we follow the standard approach of exploiting the role played by pre-existing ethnic networks and build a shift-share predictor for immigrant density at the regional level, as originally proposed by Altonji and Card (1991) and Card (2001). However, we depart from previous literature by aiming at predicting immigrant shares in *enrollment* rather than in the *overall population*. As argued earlier, immigrant enrollment is a more accurately measured than the counts in the population. In addition, immigration shares in enrollment are a more relevant measure of immigrant density in the context of our analysis.

Specifically, our aim is to predict the immigrant share in enrollment in public schools, for primary and secondary education. Our predictor for the stock of foreign students is based on data that pools all pre-university education levels, including pre-school, elementary and secondary education.[8] We use the resulting predicted enrollment share (for all pre-education levels) as our instrument for the enrollment share in primary and/or secondary public schools at the region-year level. As usual, the key exclusion restriction is that the size of the ethnic enclaves at the regional level in the base year (1994 in our case) is uncorrelated with trends in the dependent variables over the sample period (2000–2015). Because educational expenditures are a low share of household consumption, and because we explicitly control for household income in our regressions, this assumption seems quite plausible in our context. Our predictors for the enrollment levels of immigrant students, as well as the shares in enrollment, are fairly strong. This is also the case when we consider separately enrollment in primary and secondary schools, as can be seen in Table B.4 in the Appendix together with further details on the construction of the instrument.

Table 4 presents the estimates. As before, the top panel presents results for tuition expenditures using both linear (two-stage least-squares) and non-linear (IV Tobit) estimators, whereas the dependent variable in the bottom panel is the private school indicator. As before, columns 1–3 report results pooling both education levels. The coefficients in the linear model (first row) are very similar in the three columns, although only

[8] The enrollment data made public by the Spanish Ministry of Education that provides disaggregation by country of nationality of the students does not provide a breakdown by level of education.

Table 3
Educational expenditures and probability of private school choice. OLS and Tobit estimates.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Educ.	All	All	All	Primary	Primary	Primary	Secondary	Secondary	Secondary
Years	All	2000–07	2008–15	All	2000–07	2008–15	All	2000–07	2008–15
Dep. Var.	edux	edux	edux	edux	edux	edux	edux	edux	edux
Linear model	2.79	3.67***	1.79	4.91**	6.37***	0.54	0.59	2.57	4.92*
	[2.010]	[1.087]	[2.339]	[2.210]	[1.799]	[3.837]	[2.953]	[1.986]	[2.406]
Tobit mfx	1.35	2.36**	2.46	3.19*	4.35***	0.75	0.76	2.66	5.87**
	[1.881]	[1.052]	[2.323]	[1.802]	[1.178]	[2.442]	[2.197]	[1.663]	[2.146]
Tobit mfx	0.64	1.17**	1.14	1.51*	2.12***	0.35	1.08	1.39*	2.80**
(*educ* > 0)	[0.886]	[0.525]	[1.067]	[0.828]	[0.631]	[1.128]	[0.373]	[0.831]	[0.975]
Dep. Var.	private	private	private	private	private	private	private	private	private
Linear model	0.34	.44***	0.45	.65**	.79***	0.29	0.15	0.38	.86**
	[0.275]	[0.151]	[0.408]	[0.288]	[0.220]	[0.436]	[0.378]	[0.280]	[0.341]
Obs.	83,995	28,124	55,871	44,018	13,637	30,381	44,697	16,109	28,588

Notes: The table reports only the coefficient associated to the immigrant share in public schools. In columns 1–3 the immigrant share in public schools is based on pooling primary and secondary education (*IShPub*), while in columns 4–6 it refers only to primary schools (*IShPubPrimary*), and in columns 7–9 only to secondary schools (*IShPubSecondary*). Year dummies and region dummies are included in all specifications. We also include, but do not show here, controls for education of the household head, real household income, number of individuals employed living in the household, and the number of children age 3–18 in the household. In the top panel, the dependent variable is the inverse hyperbolic sine of the tuition expenditures per child (*edux*). In columns 4–9 we only use the educational expenditures of the corresponding education level, divided by the number of children in that same age group. Columns 1–3 display the results for expenditures in primary and secondary education on children age 3–18. Columns 4–6 shows the results for expenditures in primary education only and children age 3–11 and columns 7–9 for expenditures in secondary education and children age 12–18. For the Tobit estimates we report the marginal effects computed on the whole sample (mfx) and for households with positive tuition expenditures only (mfx (*educ* > 0)). The standard errors for the marginal effects are computed using the Delta Method. In the bottom panel, the dependent variable is an indicator for using private school (based on the 30-Euro per child expenditure threshold) in the corresponding education level. Standard errors are clustered at the region (CCAA) level. Regressions are population-weighted. *** $p < .01$, ** $p < .05$, * $p < .1$

Table 4
Educational expenditures and probability of private school choice. Instrumental-variables estimates for linear and Tobit models.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Educ.	All	All	All	Prim.	Prim.	Prim.	Sec.	Sec.	Sec.
Years	All	2000–07	2008–15	All	2000–07	2008–15	All	2000–07	2008–15
Dep. Var.	edux	edux	edux	edux	edux	edux	edux	edux	edux
Linear model	4.03	3.97***	3.75	14.96**	11.42***	−2.18	−2.30	0.63	8.86***
	[5.345]	[1.350]	[2.635]	[7.570]	[3.370]	[3.872]	[4.003]	[1.854]	[2.644]
Tobit mfx	−0.26	1.17	7.86***	4.09	4.02*	2.64	−2.27	0.80	11.09***
	[3.839]	[1.180]	[3.024]	[4.740]	[2.329]	[2.464]	[2.709]	[1.733]	[2.091]
Tobit mfx	−0.13	0.58	3.65**	1.93	1.96*	1.23	−1.12	0.42	5.28***
(*educ* > 0)	[1.826]	[0.583]	[1.471]	[2.231]	[1.108]	[1.154]	[1.314]	[0.910]	[1.116]
Dep. Var.	private	private	private	private	private	private	private	private	private
Linear model	0.44	.45**	1.05**	1.62*	1.24**	0.28	−0.27	0.14	1.51***
	[0.675]	[0.203]	[0.415]	[0.897]	[0.505]	[0.506]	[0.471]	[0.260]	[0.304]
First stage									
Dep. var.	IShPub	IShPub	IShPub	IShPub	IShPub	IShPub	IShPub	IShPub	IShPub
Dep. var.	Prim&Sec	Prim&Sec	Prim&Sec	Prim.	Prim.	Prim.	Sec.	Sec.	Sec.
Z/Stu00	1.98***	3.02***	5.92***	1.43**	2.49***	5.29**	2.67***	3.52***	6.83***
	[0.473]	[0.497]	[1.019]	[0.503]	[0.574]	[1.931]	[0.506]	[0.437]	[1.080]
F-test	17.5	36.9	33.7	8.1	18.9	7.5	27.9	64.8	40.1
Obs.	83,995	28,124	55,871	44,018	13,637	30,381	44,697	16,109	28,588

Notes: The table reports only the coefficient associated to the immigrant share in public schools. In columns 1–3 the immigrant share in public schools is based on pooling primary and secondary education (*IShPub*), while in columns 4–6 it refers only to primary schools (*IShPubPrimary*), and in columns 7–9 only to secondary schools (*IShPubSecondary*). Year dummies and region dummies are included in all specifications. We also include, but do not show here, controls for education of the household head, real household income, number of individuals employed living in the household, and the number of children age 3–18 in the household. In the top panel, the dependent variable is the inverse hyperbolic sine of the tuition expenditures per child (*edux*). In columns 4–9 we only use the educational expenditures of the corresponding education level, divided by the number of children in that same age group. Columns 1–3 display the results for expenditures in primary and secondary education on children age 3–18. Columns 4–6 shows the results for expenditures in primary education only and children age 3–11 and columns 7–9 for expenditures in secondary education and children age 12–18. For the IV Tobit estimates we report the marginal effects computed on the whole sample (mfx) and for households with positive tuition expenditures only (mfx (*educ* > 0)). The standard errors for the marginal effects are computed using the Delta Method. In the middle panel, the dependent variable is an indicator for using private school (based on the 30-Euro per child expenditure threshold) in the corresponding education level. The bottom panel presents the estimated coefficient of the predicted immigrant share (*Z/Stu*00) in the first-stage regression and the associated F test. Standard errors are clustered at the region (CCAA) level. Regressions are population-weighted. *** $p < .01$, ** $p < .05$, * $p < .1$

statistically significant for the period of economic expansion.[9] When we focus on primary education (columns 4–6), we find clear evidence of a positive effect of the immigrant share in enrollment on tuition expenditures during the economic boom, both on the basis of our linear and Tobit estimates, and the private school indicator in the middle panel. When we turn to secondary schooling (columns 7–9), we again find evidence of a positive and significant effect of the immigrant share on educational expenditures but only for the period 2008–2015. Qualitatively, this pattern is the same that emerged from the estimates in Table 3. However, the point estimates and the standard errors tend to be somewhat larger.

We now turn to the discussion of the magnitudes implied by our estimates in Table 4. First of all, we note that the marginal effects derived from the Tobit estimates are much larger when computed on the basis of the whole sample than for the subsample of households using

[9] As shown in the last row of Table 4, the *F*-test associated to the first-stage regression suggests our instrument is a good predictor for the immigrant shares in enrollment, with typical values around 20.

private schools. This implies that the adjustment to increased immigration took place mainly along the extensive margin, by inducing public school users to switch toward private schools. For this reason, and because of its greater simplicity, we continue our discussion of the size of the effects on the basis of the linear probability models for the private school indicator.[10] The estimate at the bottom of column 5 implies that a 1 percentage-point increase in the immigrant share in primary public schools leads to a 1.2 percentage-point increase in the probability to attend private school. The estimated effect on secondary education is slightly larger on the basis of the estimate at the bottom of column 9 (1.51 percentage points).

5.3. Cream-skimming: heterogeneous effects by education of household head

We further unpack the relationship between immigration and school choices by investigating if the responses vary on the basis of the educational attainment of household heads. Evidence of heterogeneous effects along this dimension can be found in Altonji et al. (2015), who document cream-skimming effects in school choices. Their findings suggest that better students (in terms of ability or parental background) may react more strongly by changing schools in response to variations in the environment, such as the introduction of vouchers.

In order to investigate this issue we modify our baseline model in two ways. First, to account for differences in trends by education of the household head, we include education-year dummy variables. Second, we include interactions of the immigrant share with an indicator for whether the household head has a college degree. Naturally, these interactions are only appropriate if immigration does not have a direct effect on the educational attainment of the household head, which does not seem to be the case (see Appendix Table B.2). Because the results above suggest that most of the adjustment takes place along the extensive margin, namely, by switching from public to private schools, rather than by switching from low to high tuition private schools, we focus on the regression models for the private school indicator.

Table 5 presents the estimates of the augmented model, where we stress that we are controlling for household income, as well as number of employed persons in the household and the number of children in the household. The top panel presents two-stage least-squares estimates based on the immigrant share in public schools pooling primary and secondary education. Column 1 presents results for the whole period 2000–2015. As before, we cannot reject the null of a zero coefficient in either case. However, when we consider the two sub-periods separately, a clearer picture emerges. The estimates for the period 2000–2007 (column 2) show that immigration into public schools had a large displacement effect on college-graduate households (1.49 percentage points), but did not affect the school choices of other households. In contrast, in period 2008–2015 immigration displaced households of all education levels (0.99 percentage points), and we do not find evidence of a differential effect on college-educated households. This estimates suggest that the more educated households may have been quicker to react to immigration flows by switching toward private schools. Then, in the second period, households at all education levels may have imitated that behavior.

The middle panel restricts to the sample of households with children in primary schools and, accordingly, we measure immigrant density at the level of public primary schools. The estimates here reveal a response to immigration stemming exclusively from college-graduate households in both periods. Last, the bottom panel focuses on households with children in secondary schooling. The results echo again the pattern observed in the top panel. We find evidence of a response among college-graduate households in the first period that seems to diffuse to all households, regardless of education, in the second period.[11]

Table 5
Linear probability model on the probability to use private school. Heterogeneous effects by education of the household head. 2SLS estimates.

	1	2	3
Period	All years	2000–2007	2008–2015
Dep. Var.	Private	Private	Private
Primary and secondary pooled			
IShPub	0.11	−0.06	.99**
	[0.441]	[0.261]	[0.385]
IShPub × *Cograd*	0.72	1.49***	0.41
	[0.504]	[0.366]	[0.520]
Log hh. income	.06***	.09***	.05***
	[0.012]	[0.011]	[0.012]
Obs.	83,995	28,124	55,871
Primary schooling			
IShPubPrim	0.90	0.45	0.16
	[0.568]	[0.656]	[0.479]
IShPubPrim × *Cograd*	1.12**	1.74***	.84**
	[0.462]	[0.511]	[0.425]
Log hh. income	.06***	.09***	.05***
	[0.012]	[0.020]	[0.012]
Obs.	44,018	13,637	30,381
Secondary schooling			
IShPubSec	−0.32	−0.23	1.50***
	[0.360]	[0.324]	[0.267]
IShPubSec × *Cograd*	0.31	1.40***	-0.07
	[0.559]	[0.285]	[0.699]
Log hh. income	.05***	.08***	.04***
	[0.011]	[0.005]	[0.013]
Obs.	44,697	16,109	28,588

Notes: The top panel displays the IV estimates of the linear probability model for using private school (based on the 30-Euros per child expenditure threshold) pooling primary and secondary education. The middle panel restricts the analysis to primary school only (age 3–11), and the bottom panel to secondary school (age 12–18). The coefficients shown in the table correspond to the model estimated at bottom of Table 4 extended with an indicator for whether the household head graduated from college and its interaction with the share of students at public school for each educational level (*IShPub* × *Cograd*). Year dummies and region dummies are included in all specifications. We also include, but do not show here, controls for real household income, number of individuals employed living in the household, and the number of children age 3–18 in the household. Standard errors are clustered at the region (CCAA) level. Regressions are population-weighted. In all cases the instrument is the shift-share instrument described in the Appendix and its interaction with household-head college graduate indicator. *** $p < .01$, ** $p < .05$, * $p < .1$

Let us now provide a discussion of the magnitudes of these effects and the relative contribution of immigration and household income to changes in the share of private-school users. Let us begin with the role of household income. Real household income increased by 44 log points in period 2000–2007 and fell by 7 log points in period 2008–2015. On the basis of the coefficients reported in Table 5, this led to a 4 percentage-point increase in the probability to use private *primary* schools in the first period, and to a 0.4 percentage-point reduction in the second period.[12]

Quantifying the effects of immigration is a bit more subtle because the effects are heterogenous according to the education level of the household head. Regarding primary schooling, the data show that immigrant density increased by 7 percentage points in the first period and fell by 1 percentage point in the second period. On the basis of our estimates in Table 5 (middle panel), we find that immigration led to a 12 percentage-point increase in the probability of using private *primary*

[10] The coefficients corresponding to regression models for the inverse hyperbolic sine of tuition expenditures can be roughly interpreted as one typically does with models where the dependent variable is in logs. Accordingly, a 1 percentage-point increase in the immigrant share in primary and secondary public schools leads to an increase in educational expenditures of about 11% (based on column 5) and 9% (based on column 9), respectively.

[11] The lowest t-statistic corresponding to the coefficients marked with stars in Table 5 is 2.42, substantially above the typical 1.96.

[12] Regarding primary education, in the first period the point estimate for the household income variable is 0.09 and the mean change in household income over this period is 0.44. The product of these two numbers results in the 0.04 effect (4 percentage points). Similarly, the point estimate for household income in the second period is 0.05 and the mean change in the data is −0.07, resulting in a −0.004 effect (a 0.4 percentage-point reduction). Analogous calculations produce our assessment of the effects of immigration.

Table 6
Robustness (1). 2SLS estimates .

Period	1	2	3	4	5	6	7	8	9	10
Dependent variable	2000–2007	2008–2015	2000–2007	2008–2015	2000–2007	2008–2015	2006–2015	2006–2015	2000–2007	2008–2015
	Private	Private	Private	Private	Private	Private	Private	Private	Private	Private
Primary schooling										
IShPubPrim	0.45	0.05			0.71	0.23	0.16	0.67	0.36	0.08
	[0.609]	[0.490]			[0.674]	[0.795]	[0.479]	[0.709]	[0.402]	[3.798]
IShPubPrim × Cograd	1.76***	.91**			1.75***	.84**	.84**	0.76	1.18***	.82**
	[0.550]	[0.446]			[0.514]	[0.425]	[0.425]	[0.469]	[0.380]	[0.415]
IShPop			0.96	3.62						
			[1.314]	[5.332]						
IShPop × Cograd			1.92**	1.08***						
			[0.834]	[0.408]						
Secondary schooling										
IShPubSec	−0.18	1.38***			−0.23	1.64***	1.50***	1.84***	−0.53	0.55
	[0.367]	[0.275]			[0.345]	[0.398]	[0.267]	[0.393]	[0.740]	[0.440]
IShPubSec × Cograd	1.39***	−0.10			1.40***	−0.06	−0.07	−0.01	.75***	−0.06
	[0.317]	[0.690]			[0.282]	[0.701]	[0.699]	[0.634]	[0.186]	[0.695]
IShPop			−0.17	1.39						
			[0.963]	[1.764]						
IShPop × Cograd			.92*	0.45						
			[0.535]	[0.455]						
Observations	16,109	28,588	16,109	28,588	16,109	28,588	28,588	25,069	16,109	28,588
Households	all	all	all	all	all	all	nat+immg	natives	all	all
hold. controls	no	no	yes	yes	yes	yes	yes	yes	yes	yes
Controls immig. origin	no	no	no	no	yes	yes	no	no	no	no
Region linear trends	no	no	no	no	no	no	no	no	yes	yes

Notes: The top panel displays the IV estimates of the private use linear probability model for primary school and the bottom panel for secondary school (in both cases based on the 30-Euros per child expenditure threshold). Year dummies and region dummies are included in all specifications. Columns 3–10 also include as additional controls real household income, number of individuals employed living in the household, and the number of children age 3–18 in the household. Columns 5 and 6 include as controls the share of the immigrant population originating from Africa, America, and Asia/Oceania. The share of European immigrants is excluded to avoid perfect collinearity with the immigrant share in the population. Standard errors are clustered at the region (CCAA) level. Regressions are population-weighted. The instrument is the shift-share instrument described in the Appendix and its interaction with household-head college graduate indicator. *** $p < .01$, ** $p < .05$, * $p < .1$.

schools among college-educated households during the period 2000–2007, while having no effect on less educated households.[13] During the second period, immigration reduced the probability that college-educated households use private schools by 0.8 percentage points, while again having no effect on less educated households. Weighting by the size of each group, the estimates imply that immigration was responsible for a 2.3 percentage-point increase in the probability of using private schools during the period 2000–2007 and for a 0.6 percentage-point reduction in period 2008–2015. Hence, the effect of immigration is quite sizable, contributing roughly half as much as fluctuations in household income. Our estimates imply that the combined effects of immigration and household income on the probability to use private *primary* schools were a 6.3 percentage-point increase in the first period, and a 0.6 percentage-point reduction in the second.

Let us turn now to the analysis of the effects of household income and immigration on the probability to use private *secondary* schools. The role of household income is almost identical to the case of primary schooling. Fluctuations in household income were responsible for a 3.5 percentage-point increase in the private school probability in the first period, and a 0.3 percentage-point reduction in period 2008–2015. The combined effects of income and immigration are also very similar to the previous case. However, there is an important difference in the responses of households as a function of their education level. During the period 2000–2007, immigration led to an 11 percentage-point increase in the private school probability for college-graduate households, but no effect on less educated households. In contrast, in period 2008–2015, immigration was responsible for a 2 percentage-point reduction in the private school choice probability of all households, regardless of their education level. As before, the relative size of the effects due to immigration are roughly half as large as those of household income.

In conclusion, our results show that understanding the relationship between immigration and public-private school choices requires allowing for heterogeneous responses at different education levels (primary and secondary) and across households with different educational attainment. We have also shown that immigration was an important factor behind the changes in the prevalence of private schools use over our period of interest, with an effect that was roughly half as large as that of household income.

6. Robustness

This section conducts extensive sensitivity analysis on our main findings. First, we present variations of our earlier models that maintain the definition of the private school indicator used throughout the paper. Second, we experiment with a more flexible expenditures threshold to identify private school attendance.[14]

Table 6 presents estimates on a collection of checks where the expenditure threshold associated to private school attendance is fixed at 30 Euros in real terms. Throughout the table we focus on the linear probability model for private school attendance as the main outcome variable. Column 1 and 2 show that our results are robust to the exclusion of household controls (household income, employment in the household and number of school-age children). The estimated coefficients in these two columns are of the same sign, similar magnitude and significance as those in Table 5. For primary education, we find evidence of displacement among college-educate households in both periods, though the effect is much larger in period 2000–2007. In secondary schools, the estimates also point to displacement but in the first period this is only observed among college-educated households while in the second period it affects all households regardless of education. In sum, our findings are fairly robust to the set of covariates included in our

[13] In this prediction exercise we consider that non-significant coefficients have a zero effect.

[14] Additionally, we have also verified that our results are robust to an alternative definition for educational expenditures per child, where we count all children age 2–18 in the denominator, regardless of the schooling level that we are referring to.

baseline specification. Columns 3 and 4 consider using the more standard measure of immigrant density based on the foreign-born share in the population. The pattern of the point estimates is similar to that obtained measuring immigrant density in terms of enrollment. However, the standard errors are larger on average due to the lower accuracy of the population-based measure of immigrant density.

The composition of immigrants in Spain is quite diverse. By 2010, about 40% originated from South and Central America, 29% from the rest of Europe, 23% from Africa, and 6% from Asia and Oceania. Immigrants from Latin America generally speak Spanish, which may facilitate the integration and assimilation process of those children. Columns 5 and 6 explore whether immigrants from different origins have different effects on the public–private school choice. We do so by including an additional set of regressors that describe the composition of the immigrant population in each year and region. Specifically, we distinguish immigrants on the basis of continent of origin and include the corresponding shares as controls, excluding the share of European immigrants to avoid perfect collinearity. We treat these regressors as exogenous, which is a relatively mild assumption that requires the origin *composition* of the foreign-born population to be unaffected by unobserved shocks to educational expenditures in Spain. The coefficients associated to the immigrant shares in public schools and their interactions with the college-educated dummy are practically identical to those reported in columns 1 and 2. Furthermore, the variables describing the composition of the immigrant population are seldom significant. Our interpretation is that there is no evidence of differential effects by continent of origin of the immigrant population.

Columns 7 and 8 examine whether excluding immigrants from our sample has an impact on the estimates. As mentioned in Section 4, the nationality of the household head is only known in the FES from year 2006 onward. Thus the sample in these two columns corresponds to the period 2006–2015, which mostly comprises the recession years. Nevertheless, the estimated coefficients in the two columns (natives and immigrants in column 7 and only natives in column 8) are very close to each other, suggesting that removing immigrants from the sample has little effect on our previous estimates. Lastly, columns 9 and 10 include region linear trends in our specification. The coefficients associated to the immigrant shares in public schools are reduced only slightly in the case of primary education relative to columns 1 and 2. For secondary schools, the reduction in the coefficients is more significant, but the qualitative results also remain largely unchanged.

Since 2006 a fraction of households are interviewed in two consecutive years in the FES, providing an opportunity to better account for individual heterogeneity.[15] Specifically, we are able to estimate within-household fixed-effects models. The results are presented in columns 1 and 2 of Table B.3. The point estimates are positive both for the primary school and secondary school sub-samples in the top and bottom panels, respectively. But we are only able to reject the null of a zero effect in the latter case. These fixed-effects estimates suggest that increases in immigrant density in (secondary level) public schools are associated to increases in the probability to send children to private schools. In columns 3 and 4 we present two-stage least-squares estimates that also include household fixed-effects. Again, the point estimates are positive at both levels of education. However, standard errors increase a lot, rending the estimates not statistically significant. Columns 5 and 6 present two-stage least-squares estimates for the sub-sample of households that used public schools in the first period they are interviewed, and estimate our models using reported educational expenditures in their second year in the survey. A positive and significant effect on this sub-sample would provide additional evidence of an extensive-margin response to immigration. Unfortunately, the estimates are very noisy and we cannot reject the zero null hypothesis in any of the two columns. In sum, household

[15] The households that are interviewed in two consecutive years are less than half of all the FES respondents.

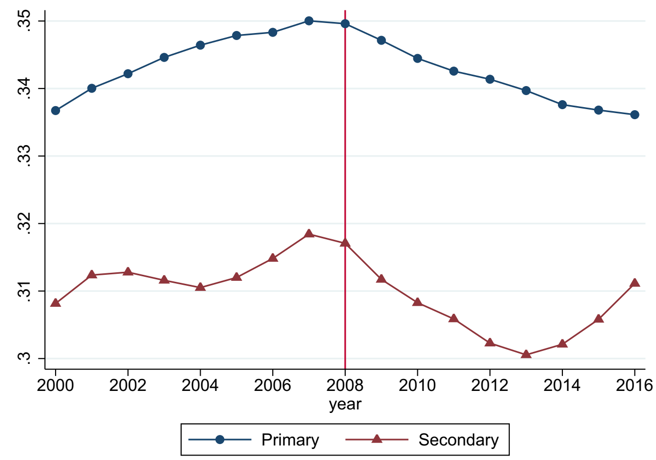

Fig. 7. Share of native students in private schools (primary and secondary). *Notes:* The figure displays the share of native students enrolled at private schools separately for primary and secondary education. Source: Administrative enrollment data, Spanish Ministry of Education.

fixed-effects models are not very informative, probably due to the short longitudinal dimension of the data.

Next, we relax the assumption of the fixed 30-Euro threshold (in real terms) used to identify private school attendance. Specifically, we calculate year-specific cutoffs that are calibrated to match the annual shares of students enrolled at private schools obtained from the Spanish Education Enrollment Registry. The private shares in primary and secondary grades in the Enrollment Registry are depicted in Fig. 7. The private share in primary education (blue line with circles) is about 3 points higher than for secondary education (red line with triangles). The private share in primary school is fairly constant at around 0.34, although it displays a slight increase for the period 2000–2008, from 0.34 to 0.35, and then falls back to 0.336 in 2015. The private share in enrollment at secondary schools also increases over the period 2000–2008 from 0.31 to 0.32, and decreases to 0.30 by 2012. However, in contrast to the share of enrollment in private primary schools, the private share in secondary education rebounds after 2012 to reach a level similar to that in 2008, 0.31.

Next, we turn to our household-level data and calibrate the annual expenditure cutoffs to try to match the private shares in enrollment (as observed in the administrative data) for all years. The resulting shares are represented by the dashed lines in Fig. 8 (for primary schools) and in Fig. 9 (for secondary schools). We are able to match very well the private shares from year 2006 onward. The cutoffs for the 2007–2015 period are set at 2 Euros for years 2007–2011, 5 Euros for years 2012 and 2013 and 25 and 35 Euros, respectively, for years 2014 and 2015. For the years prior to 2007, the cutoff is set at zero Euros and we are still unable to generate a high enough private share in the household-level data.[16] The main reason is that over these years only about 1 in 4 households in the Family Expenditures Survey report positive educational expenditures. We suspect that the reason for this is the presence of a sizable number of concerted schools that did not charge tuition. While these schools are identified as private in the Enrollment Registry data, we are considering them as public on the basis of tuition expenditures. Over the decade, most of these concerted schools started charging positive tuition, which allowed us to identify them as private. Admittedly, this exercise would have been more convincing if we had been able to reproduce the private shares for all years in our sample. Nonetheless,

[16] For secondary education (Fig. 8), the cutoffs for the 2007–2015 are set at 2 Euros for years 2007–2009, 5 Euros for year 2010, 10 Euros for year 2011 and 15 Euros from 2012 to 2015. For the years prior to 2007, the cutoff is also set at zero Euros.

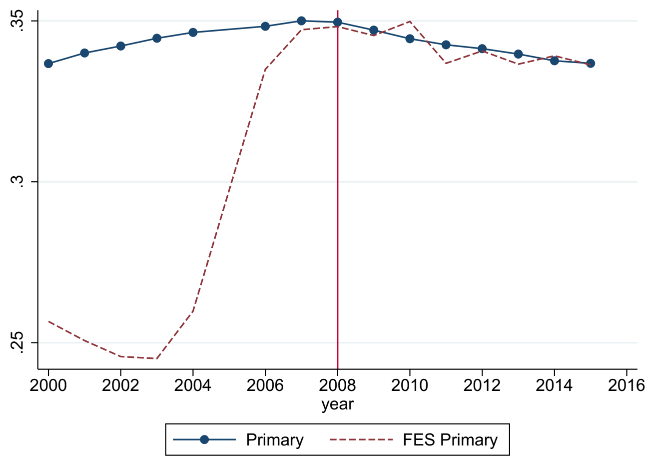

Fig. 8. Share of students in private schools. Primary education, native plus immigrant students. *Notes:* The figure displays the share of students enrolled in private schools in the administrative enrollment data (*Primary*) and in our data when the year-specific thresholds are employed in estimation (*FESPrimary*). Source: Administrative enrollment data, Spanish Ministry of Education and Family Expenditures Survey.

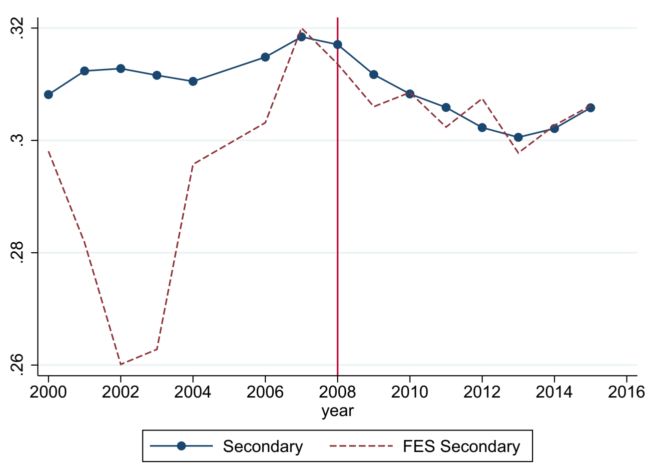

Fig. 9. Share of students in private schools. Secondary education, natives plus immigrant students. *Notes:* The figure displays the share of students enrolled in private schools in the administrative enrollment data (*Secondary*) and in our data when the year-specific thresholds are employed in estimation (*FESSecondary*). Source: Administrative enrollment data, Spanish Ministry of Education and Family Expenditures Survey.

the results from this additional analysis reported in Table 7 do not reveal substantial differences with respect to the constant cutoff estimates presented earlier the paper. In particular, this is the case for the findings related to the period 2008–2015, for which we are able to replicate the administrative shares of enrollment in private schools for all years. All in all, we conclude that the results on the effects of immigration on the public-private school choice are not sensitive to the assumption of a constant tuition cutoff.

7. Discussion

In the previous sections we have documented that household income is, not surprisingly, an important determinant of tuition expenditures, largely connected to the private-public school choice. In good economic times, households are more likely to send their children to private schools. This decision is partly reversed in times of falling income and employment. Secondly, increases in immigrant density in public schools also lead to displacement of natives toward private schools, and this response is more pronounced among households with high educational attainment. We now discuss the mechanisms that can account for these findings.

7.1. Consumption expenditures

Prior to turning to the theoretical mechanisms that can account for the empirical findings, it is helpful to consider the budget constraint of households. Specifically, we ask what are the adjustments in savings or consumption that took place to accommodate the reported immigration-induced increase in tuition expenditures. To address this question we exploit the detailed information on consumption expenditures contained in the FES.

We follow the approach in Dustmann et al. (2017) and estimate a series of models with varying measures of household expenditures that resemble our baseline specification. The results are presented in Table 8, where we distinguish by education level. The dependent variable in column 1 is the consumption income ratio and in column 2 the inverse hyperbolic sine of consumption including all categories. In both cases we cannot reject the zero null suggesting that immigration did not have an effect on household income or savings. As a result, the increase in tuition expenditures must have triggered offsetting reductions in other consumption categories.

The remaining columns in Table 8 investigate the effect of immigration on alternative consumption categories. Interestingly, the estimates in column 3 suggest no changes in the overall measure of educational expenditures. This variable includes tuition fees and in-school extracurricular activities in primary and secondary education (as in the educational expenditure measure employed throughout the paper). It also contains expenditures in tertiary education and after-school activities such as courses on foreign languages or technology, that take place outside the school. Column 4 shows the estimates for educational expenditures in tertiary education and after-school activities, whereas column 5 restricts to spending in after-school activities only. The estimated coefficients in these two columns are negative when pooling primary and secondary schooling. When we separately consider immigrant density in primary and secondary schooling, the point estimates gain precision. Particularly in column 5, when we restrict to spending in after-school activities only, we obtain negative point estimates that are significantly different from zero at the conventional level. Last, columns 6–8 consider spending in non-education consumption categories, which do not seem affected by immigration, further supporting our interpretation of offsetting effects within the educational expenditure categories.[17]

These results suggest that the increase in the use of private schools documented earlier in the paper was funded by reducing expenditures in after-school activities taking place outside the school. Thus, increases in immigrant density in public schools seem to have led Spanish families to switch from a scenario where they sent their children to public school and complemented their education with after-school activities that took place outside schools, to a scenario where these families sent their children to private schools. To the extent that private schools offer similar activities at no extra cost, this switch may not have consequences in terms of human capital accumulation. However, it suggests an increase in immigrant segregation in public schools.

7.2. Immigration and school assignment rules

In Spain, as well as in many other countries, households submit school preferences and local governments match students to public schools on the basis of some pre-specified assignment rule. Most Spanish cities follow the so-called Boston Mechanism. Accordingly, households

[17] In Table 8 column 6 includes food, drink and clothing; column 7 considers expenditures in housing and column 8 all the remaining consumption categories (i.e. leisure, communication, transportation, health, hotels and other consumption).

Table 7
Robustness (2). 2SLS with calibrated annual expenditure cutoffs .

	1	2	3	4	5	6	7	8	9	10
Period	2000–2007	2008–2015	2000–2007	2008–2015	2000–2007	2008–2015	2006–2015	2006–2015	2000–2007	2008–2015
Dependent variable	Private	Private	Private	Private	Private	Private	Private	Private	Private	Private
Primary schooling										
IShPubPrim	−0.03	0.16			0.14	0.66	0.27	0.59	−0.60	−1.62
	[0.591]	[0.634]			[0.657]	[1.068]	[0.618]	[0.675]	[0.401]	[4.481]
IShPubPrim × Cograd	1.77***	.83*			1.76***	.76*	.76*	0.72	1.20***	.74*
	[0.618]	[0.452]			[0.584]	[0.433]	[0.433]	[0.467]	[0.447]	[0.422]
IShPop			0.16	3.31						
			[1.173]	[5.360]						
IShPop × Cograd			1.85**	1.03**						
			[0.933]	[0.401]						
Secondary schooling										
IShPubSec	0.40	2.29***			0.43	3.06***	2.41***	2.52***	−1.55	.94**
	[0.451]	[0.428]			[0.359]	[0.623]	[0.404]	[0.488]	[1.526]	[0.386]
IShPubSec × Cograd	1.00***	−0.21			1.01***	−0.17	−0.17	−0.12	.52***	−0.18
	[0.311]	[0.724]			[0.284]	[0.737]	[0.731]	[0.667]	[0.180]	[0.732]
IShPop			1.00	3.36						
			[0.844]	[3.128]						
IShPop × Cograd			0.56	0.38						
			[0.525]	[0.473]						
Observations	16,109	28,588	16,109	28,588	16,109	28,588	28,588	25,069	16,109	28,588
Households	all	all	all	all	all	all	nat+immg	natives	all	all
Hhold. controls	no	no	yes	yes	yes	yes	yes	yes	yes	yes
Controls immig. origin	no	no	no	no	yes	yes	no	no	no	no
region linear trends	no	no	yes	yes	yes	yes	yes	yes	yes	yes

Notes: The top panel displays the IV estimates of the private use linear probability model for primary school and the bottom panel for secondary school. Private use is defined by calibrated annual cutoffs to match the private share in the administrative enrollment data (see Section 6). Year dummies and region dummies are included in all specifications. Columns 3–10 also include as additional controls real household income, number of individuals employed living in the household, and the number of children age 3–18 in the household. Standard errors are clustered at the region (CCAA) level. Regressions are population-weighted. The instrument is the shift-share instrument described in the Appendix and its interaction with household-head college graduate indicator. *** $p < .01$, ** $p < .05$, * $p < .1$

Table 8
Consumption Expenditures. 2SLS estimates. Years 2000–2015 .

Spending Category	Obs.	1 C/Inc.	2 All C	3 Education All	4 Education Tertiary & After-school	5 After-school	6 Food, drink, clothing	7 Housing	8 Other
IShPub	83,995	−1.22	−2.76	−6.70	−67.88	−57.01	3.92	−11.08	−2.31
		[1.781]	[4.884]	[26.205]	[94.524]	[78.441]	[7.633]	[16.920]	[6.783]
IShPubPrimary	44,018	−1.50	−0.25	7.15	−14.28**	−10.06**	0.99	−1.22	−0.94
		[1.781]	[0.816]	[6.990]	[5.824]	[4.489]	[1.213]	[1.418]	[1.972]
IShPubSecondary	28.588	−0.74	−0.16	−1.88	−4.04	−4.99**	0.55	−1.32	−0.02
		[1.597]	[0.433]	[4.062]	[2.667]	[2.330]	[0.771]	[1.142]	[0.769]

Notes: The dependent variable in column 1 is the ratio of expenditures in all consumption categories over income. The dependent variables in all remaining columns are the inverse hyperbolic sine of different categories of consumption expenditures. Column 2 focuses on all consumption categories. Column 3 refers to all education expenditures: tuition and in-school extracurricular activities in primary and secondary education, tuition and fees in tertiary education and after-school activities such as language lessons, computer courses and supplementary lectures to reinforce regular education that do not take place at school. The dependent variable in column 4 includes educational expenditures in tertiary education and after-school activities only, and column 5 restricts further to only after-school activities. Last, columns 6, 7 and 8 follow the consumption expenditures classification in Dustmann et al. (2017). Column 6 focuses on food, drink and clothing; column 7 on housing, and column 8 on the rest of consumption expenditures: leisure, communication, transportation, health, hotels and other consumption. Year dummies and region dummies are included in all specifications. All columns include the number of individuals employed living in the household, and the number of children age 3–18 in the household. From column 2 to 8 real household income is also included in estimation. Standard errors are clustered at the region (CCAA) level. Regressions are population-weighted. The instrument is the shift-share instrument described in the Appendix and its interaction with household-head college graduate indicator. *** $p < .01$, ** $p < .05$, * $p < .1$

submit applications where they rank public and concerted schools according to their preferences. Whenever possible, children are allocated to their first choice. Naturally, some schools are over-subscribed because they are viewed as more desirable due to their location, teacher quality, peer effects, and so on. In these cases, applications are ranked using some priority rules, which award points on the basis of family characteristics.[18] The school is then filled with the applicants with the highest priority. Students who do not get their first choice school are allocated to non-oversubscribed schools that still have vacancies.

Priority rules typically favor low-income households and large families. As a result, on average, applications from immigrant households tend to obtain higher scores, displacing higher income native households to schools with lower demand. In general, concerted schools require tuition payments and they may be viewed as less desirable among lower-income immigrant families. In the presence of a large immigration boom higher-income native families may rank concerted schools first to avoid non-desired public school. This change in natives' school preferences is consistent with our earlier findings.

[18] For example, in the case of Barcelona, having a sibling already in the school awards 40 points, living in the schools catchment area awards 30 points, and families from disadvantaged economic backgrounds have 10 additional points. See Calsamiglia and Guell (2014) for a detailed description of the allocation rules.

Table 9
Student-teacher ratios.

Dep. var.	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Years	All	All	2000,2010	2000,2010	All	All	2000,2010	2000,2010
Estimation	OLS	2SLS	OLS	2SLS	OLS	2SLS	OLS	2SLS
Students/Teacher	Public	Public	Public	Public	Private	Private	Private	Private
IShPub	5.50*	14.75***	8.09	15.98***	3.59	3.28	2.98	2.68
	[2.90]	[4.55]	[5.88]	[3.12]	[5.54]	[4.24]	[3.79]	[4.72]
Log Real GDP	−4.27	−0.85	−13.56	−7.94*	−12.35	−33.69***	−33.87***	−34.13***
	[3.50]	[3.59]	[9.32]	[4.45]	[8.96]	[5.06]	[5.04]	[5.25]
Observations	204	204	34	34	204	34	34	34
R-squared	0.95	0.94	0.97	0.96	0.94	0.98	0.98	0.98
IShPop	5.96	20.51*	6.96	16.50***	4.72	−0.32	3.22	−2.07
	[3.44]	[11.25]	[4.98]	[5.78]	[5.08]	[9.92]	[4.04]	[5.34]
Log Real GDP	−4.53	−0.19	−15.91*	−11.25***	−12.26	−13.77	−34.46***	−37.04***
	[3.30]	[4.49]	[7.73]	[4.07]	[8.48]	[8.40]	[4.54]	[4.23]
Observations	204	204	34	34	204	204	34	34
R-squared	0.95	0.92	0.97	0.95	0.94	0.94	0.98	0.97

Notes: IShPub. is the share of immigrant students in total enrollment in primary and secondary public schools. *IShPop* refers to foreign-born share in the population. Year dummies and region dummies included in all specifications. Standard errors are clustered at the regional level. Observations are weighted by the year-2000 population. Instruments are predictors of immigrant density using existing ethnic networks (initialized in 1990). *** $p<.01$, ** $p<.05$, * $p<.1$

7.3. Immigration and the political economy of education

The interpretation above based on the school assignment rules is only satisfactory in the short run since it takes the quantity and quality of schools as given. Presumably, a large immigration shock like the one analyzed here will have the potential to alter the political-economic equilibrium that determines the supply and the quality of public schools.

To examine this issue we rely on the predictions of the dual provision system of public services developed by Epple and Romano (1996) and extended by Coen-Pirani (2011) to analyze the effects of immigration on public education. In this framework households are heterogeneous in (lifetime) income and on their taste for private education, and choose between public school (at zero tuition) and a menu of private schools that vary in tuition and quality. Public education is financed by a proportional income tax, determined by majority vote. In this model, richer households opt out of public schools and send their children to private schools. Conditional on income this is also the case for households with fewer children.

In a recent study, Tanaka et al. (2017) calibrate a version of this model to quantify the effects of Spain's 2000–2008 immigration wave on public school funding and private–public school choices. They argue that immigration led to a 3% reduction in public spending per student, leading to an increase in the share of native households using private schools. In their model, at a given income tax rate, an immigrant inflow that lowers average household income leads to lower tax revenue per household (and per student). As a result, educational expenditure per student in public schools falls, inducing some native households to opt out to private schools. The natives that are displaced from public to private school by immigration are, *ceteris paribus* the wealthiest households among the public school users. Naturally, when these voters stop using public schools, their preferred level of funding for public schools falls, leading to a new equilibrium with a lower quality for public schools.

These predictions are consistent with the empirical findings in this paper. Namely, an increase in immigration leads to an increase in tuition expenditures, which is largely linked to increases in private school use. Furthermore, we find that this increase occurs mainly among higher-education households. In our empirical model we control for *current* household income. Lacking information on households' assets, one can interpret the education level of household heads as a proxy for wealth or, perhaps more appropriately, a measure of expected *lifetime* income. From this viewpoint households with higher education will also tend to have higher wealth.

In the model by Tanaka et al. (2017), native flight toward private schools is driven by a deterioration in the funding and quality of public schools triggered by immigration. Thus we now turn to investigate whether this has been the case in Spain. We proxy the qualify of education by the number of students per teacher.[19] Unfortunately, we do not have data on student-teacher ratios at the school level but we do know the aggregate student enrollment and the number of teachers at the regional level, separately for private and public schools. Hence, our results on this question are more vulnerable to omitted-variable bias and should be taken as less definitive.

To formally test for this association, we propose the following regression model:

$$\left(\frac{Students_{r,t}}{Teachers_{r,t}}\right)^{Pub} = \alpha_r + \lambda_t + \beta ISh_{r,t} + \gamma lnRGDP_{r,t} + u_{r,t}, \tag{3}$$

where the dependent variable is the student–teacher ratio in public schools in region r in year t, the main regressors of interest is the immigrant share (*ISh*), measured either as the foreign-born share in the population or the share of foreign students in total enrollment in public (primary or secondary) schools. The regression also controls for the log of real GDP in the region. The latter is meant to proxy for tax revenue, which is an important determinant of the resources (including the number of teachers) that regional governments devote to the funding of public schools. The model also includes a non-parametric time trend in the form of year dummy variables, and region dummy variables. We will also estimate a similar model for the student–teacher ratio in private schools.

Table 9 presents the OLS and IV estimates of the model in Eq. (3). In columns 1–4 the dependent variable is the student–teacher ratio in public schools.

The top panel measures immigrant density on the basis of enrollment in public schools, and the bottom panel employs the foreign-born share in the population. Column 1 presents OLS estimates based on annual data. The point estimates are positive but only marginally significant for the enrollment-based measures of immigrant density. The estimates in column 1 also hint at a negative association between student–teacher ratios and real GDP, as expected. Naturally, these OLS estimates may be biased if, for example, regions that have experienced a positive income shock are also more able to invest in their public schools (reducing student–teacher ratios). To address this point, column 2 presents 2SLS estimates. The point estimates for the effect of the immigrant density are substantially higher than those reported in column 1 and statisti-

[19] Hanushek (2003) reviews the literature on the effects of class size on student performance. While many studies have used this measure, there is an open debate regarding whether smaller class sizes are cost-effective policy to improve student performance.

cally significant. Columns 3 and 4 report analogous long-difference estimates that rely only on data for years 2000 and 2010. These estimates largely confirm the findings in column 2 but display gains in precision. Columns 5–6 conduct a similar analysis but using as a dependent variable the student–teacher ratio in *private* schools. In this case we are not able to reject the null hypothesis of a zero effect of migration. It is also worth noting that the pattern is qualitatively the same when measuring immigration density in terms of population (bottom panel). However, the standard errors associated to the immigrant density variable are now much larger. This highlights the gains from using a more accurate measure of immigrant density.

In sum, we find evidence that increases in the regional immigrant share led to increases in student-teacher ratios in public schools, but this was not the case in private schools. This finding suggests that regions that experienced increases in immigrant density may have suffered from an increase in congestion in their public schools, which may have had detrimental effects on the quality of their education. This evidence is also consistent with the motivation behind the "native flight" towards private schools predicted by the political-economy model of education finance.

8. Conclusions

This paper provides an empirical analysis of school choices in Spain during a period of important economic and demographic changes. Since the early 2000s the economy was growing rapidly. This economic prosperity attracted a dramatic large number of immigrants that was abruptly halted by the Great Recession and subsequent austerity policies.

Our results reveal that immigration is an important determinant of households' school choices. The increase in the share of immigrants in the population during the period 2000–2007 triggered a large shift of natives from public to private schools, that was only partially reversed when net migration became negative during the recession. We also find that this native flight toward private schools was led by college-educated households, possibly driven by concerns of crowding in public schools.

These findings suggest that immigration may have contributed to widening the gap between the educational investments of rich and poorer households. The evidence of *cream-skimming* points toward greater segregation at the school level, with an increased concentration of disadvantaged native children and immigrants in public schools, in line with the findings of Cascio and Lewis (2012) for California. This can have important long-term consequences for the educational attainment of both native and immigrant children (Gould et al., 2004; 2009), and poses an important policy challenge for Spain over the coming decades. School segregation in terms of income and ethnicity may hinder economic upward mobility and social integration. Policymakers should develop initiatives aimed at maintaining the quality of public education in order to mitigate the flight toward private schools of the children with more favorable socio-economic background.

Appendix A. Instrument definitions

Our instrumental-variables strategy is based on the use of ethnic enclaves to predict the location of recent immigrants pioneered by Altonji and Card (1991) and Card (2001). This approach has been shown to be effective in the context of Spain's immigration wave by Farre et al. (2011) and Gonzalez and Ortega (2013). Here we update their analysis to include more recent years, and discuss the performance of the predictor separately in the two phases of the business cycle.

In essence, the ethnic networks instrument is a predictor for the stock of the foreign-born population at the regional level in every year between 2000 and 2015, based on historical information on immigrant networks in year 1990. We expect current location decisions of migrants to be influenced by the location decisions of earlier migrants from the same country of origin. To the extent that this influence is orthogonal to current local demand conditions, this instrumental-variables approach will help uncover the causal effect of immigration on our outcomes of interest.

Specifically, we define the predicted current foreign-born population in province r and year $t > t_0$ by:

$$Z_{r,t} = \sum_{c} \left(\frac{FB_{c,r,t_0}}{FB_{c,t_0}} \right) FB_{c,t}, \tag{4}$$

where FB_{c,r,t_0} is the number of individuals born in foreign country c that inhabited province i in some base year t_0. The term in parenthesis is the province share of c-born individuals in the base year, which can be found to identify the most important destination provinces for each immigrant group. In our case the base year is 1990 and we build the province-level distribution of the foreign-born population (by birth country) on the basis of the 1990 Population Census. The only time-varying term in the equation is $FB_{c,\,t}$, the stock of individuals originated from country c that live in Spain in year t. Hence, an inflow of, say, Polish immigrants into Spain in 2006 will lead to a predicted contemporaneous increase in the Polish population in each province in proportion to the size of the Polish enclave in that province in the base year.

In this paper our regional unit will be autonomous communities.[20] Thus we aggregate the province-level predictor above at the autonomous community level. Because our regression models always include region fixed-effects, identification will be based on within-region changes in the stock of the foreign-born population, which we will predict on the basis of

$$\Delta Z_{r,t} = \sum_{c} \left(\frac{FB_{c,r,t_0}}{FB_{c,t_0}} \right) \Delta FB_{c,t}. \tag{5}$$

In words, the predicted increase in the foreign-born population in a region in a given year is the weighted sum of the vector of origin-specific, Spain-wide immigrant inflows in the current year. Note though that the weights vary by region. Regions with large shares to a specific immigrant group (on the basis of the 1990 enclaves) will be predicted to receive the largest shares of the current inflow of immigrants belonging to that specific group.

The previous discussion refers to the stock of immigrants, or its change over time. However, in our regression models, we focus on immigrant shares, either in the population or in the student body. In the original renditions of the instrument, many authors divide the predicted stock of immigrants, $Z_{r,\,t}$, by the current population in the region, $Pop_{r,\,t}$. This choice results in a relatively weak first stage because several Spanish regions have experienced large changes in population over our sample period that are not driven by international migration, but rather by internal migration. This led us to normalize the regional stocks of foreign-born using the region's population in a fixed year, which we chose to be year 2000. Thus we predict the current share of foreign-born in the population of region r in year t, $FB_{r,\,t}/Pop_{r,\,t}$, using $Z_{r,\,t}/Pop_{r,\,2000}$, which obviously does not change unless its numerator changes over time. We also note that this variation does not require additional exogeneity assumptions for the validity of the instrument, and does not affect the interpretation of the structural coefficient in the second-stage regression.

The previous discussion describes precisely our construction for the predicted (stock and share) of immigrants in the *population*, which is based on data from the Continuous Population Registry. The predicted stock of foreign students and its share in overall *enrollment* are constructed in a very similar manner. The only differences are as follows. Enrollment is measured using administrative data produced by the Spanish Ministry of Education for the period 1994–2015. The data distinguish between public and non-public (private) schools and between Spanish nationals and enrollment of children with foreign nationality.

[20] Spain is composed of 17 autonomous communities, further subdivided into 50 provinces.

Our predictor for the stock of foreign students is based on data that pools all pre-university education levels, including pre-school, elementary and secondary education. The reason is that these are the only enrollment data that are disaggregated by country of nationality. In building the predictor (as in Eq. 4), the base year used to measure the size of the regional ethnic enclaves is 1994. Exactly as we did earlier, we then normalize by the year-2000 foreign-born population (as measured in the Continuous Population Registry). We use the resulting predicted enrollment share as our instrument for the enrollment share in primary and secondary (public) schools in the region.

Table B.4 assesses the predictive power of our instruments. This is not strictly speaking the first-stage regression of our estimates because it is estimated at the regional level, rather than on households. However, it provides a simple way to assess the relevance of the instrument. The top panel refers to predictors of the foreign born stock and share in the population. In columns 1–3 we predict the stock of foreign-born for all years, for the 2000–2007 (boom), and for the 2008–2015 (recession) years. The predictor performs well, but we note the standard errors in column 2 (boom) are much lower than in column 3 (recession). In columns 4–6 the dependent variable is the foreign-born share in the population and we use the standard predicted foreign-born share. Clearly, the performance is much worse than in the case of stocks. In columns 7–9 we use our variation on the predicted share that normalizes by the year 2000 population. In this case the performance of the instrument is vastly increased, with lower standard errors during the boom period (columns 5 and 8 versus columns 6 and 9.

The second panel of the table reports on the performance of the predictor for the stock and share of foreign students in *enrollment* in primary and secondary schools. Columns 1–3 refer to the level (stock) of enrollment. Clearly, our predictor works well in both sub-periods, but standard errors are much lower for period 2000–2007. When aiming at enrollment shares in columns 4–9, the predictor also performs well, though standard errors are one order of magnitude larger than in the previous columns, and somewhat smaller when normalizing by enrollment in year 2000. The third and fourth panels break down enrollment in public schools by education levels (primary and secondary). It is important to note that in both cases we use the same identical shift-share predictor for overall immigrant enrollment in both education levels. The third panel focuses now on predicting immigrant enrollment in public schools for primary education. The instrument is a fairly good predictor in all periods, but we note that standard errors are three times as large for the period 2008–2015 compared to the previous period. The bottom panel uses the instrument to predict immigrant enrollment in public, secondary schools. The performance of the predictor is very similar to what we found for primary schools.

Appendix B. Additional Tables

Table B1
Comparison private users versus public users .

Users Variable	Private Mean	Public Mean	Private/Public ratio
HH. Income	32,267	23,643	1.36
HH. Employed	1.50	1.28	1.17
HH. HS. Grad.	0.24	0.20	1.19
HH. Co. Grad.	0.44	0.23	1.90
Spouse HS. Grad.	0.23	0.21	1.12
Spouse Co. Grad.	0.42	0.23	1.84
HH. Immig.	0.15	0.20	0.74
Children 3–11	1.48	1.44	1.03
Children 12–18	1.26	1.52	0.83
N. households	25447	58548	0.43

Notes: Means computed using the survey weights. See Table 1 for variable definitions.

Table B2
Auxiliary Regressions: other outcomes, 2SLS.

Dep. Var.	(1) Co. Grad. HH	(2) Emp. HH	(3) Log HH. Income
IShPub	0.10	−0.01	0.89
	[0.309]	[1.602]	[1.628]
secondary head		.18***	.22***
		[0.014]	[0.016]
tertiary head		.34***	.56***
		[0.017]	[0.027]
children 3–18	.02**	.01*	−.01***
	[0.007]	[0.005]	[0.004]
Observations	83,995	83,995	83,995
R-squared	0.0239	0.086	0.160
Years	2000–2015	2000–2015	2000–2015
Educ. Level	All	All	All

Notes: Year and region dummies, and a the number of school-age children are included in all columns. Columns 2 and 3 also control for education household head (high school and college graduate dummies). Regressions are population-weighted. In all cases the instrument is the predictor of the foreign-born share using existing ethnic networks (initialized in 1990). The sample includes households with children ages 3–18 and years 2000–2015. *** $p < .01$, ** $p < .05$, * $p < .1$

Table B3
Panel regressions .

	1	2	3	4	5	6
Period	2006–2015	2006–2015	2006–2015	2006–2015	2006–2015	2006–2015
Dependent variable	edux	private	edux	private	edux	private
Primary Schooling						
IShPubPrim	1.07	0.37	54.68	3.48	−3.58	0.29
	[5.818]	[0.950]	[83.472]	[12.993]	[16.166]	[2.351]
Observations	27,482	27,482	27,482	27,482	8,353	8,353
First-stage F-test			2.55	2.55	5.01	5.01
Secondary Schooling						
IShPubSec	5.99*	1.27**	7.89	1.12	−0.22	0.06
	[3.328]	[0.461]	[15.367]	[2.469]	[2.994]	[0.615]
Observations	26,277	26,277	26,277	26,277	9,094	9,094
Estimation	FE	FE	IV FE	IV FE	IV	IV
Sample	All	All	All	All	Public t−1	Public t−1
First-stage F-test			21.11	21.11	23.18	23.18

Notes: Columns 1 and 2 report the fixed effect (FE) estimates of the model in Eq. (1) and (2). Columns 3 and 4 reports the IV estimates of the FE models. Columns 5 and 6 reports the IV estimates of the linear model of educational expenditures and private use conditional on the sample of individuals who attended a public school when interviewed for the first time in the survey (Public $t-1$). The panel structure of the survey is available only from 2006 onwards. *** p < .01, ** p < .05, * p < 0.

Table B4
Instrument relevance. Region-level regressions.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Years	All	2000–2007	2008–2015	All	2000–2007	2008–2015	All	2000–2007	2008–2015
Dep. Var.	FB	FB	FB	IShPop	IShPop	IShPop	IShPop	IShPop	IShPop
ZFB	0.97***	1.02***	0.82***						
	[0.04]	[0.06]	[0.15]						
ZFB/Pop				0.12	0.21*	0.05			
				[0.09]	[0.12]	[0.16]			
ZFB/Pop2000							0.21***	0.30***	0.26**
							[0.06]	[0.09]	[0.12]
Dep. Var.	ImPub	ImPub	ImPub	IShPub	IShPub	IShPub	IShPub	IShPub	IShPub
ZFS	0.50***	0.49***	0.51***						
	[0.03]	[0.03]	[0.09]						
ZFS/Stu				2.80***	3.83***	7.56***			
				[0.32]	[0.35]	[0.58]			
ZFS/Stu2000							2.00***	3.12***	5.87***
							[0.25]	[0.29]	[0.59]
Dep. Var.	ImPubPrim	ImPubPrim	ImPubPrim	IShPubPrim	IShPubPrim	IShPubPrim	IShPubPrim	IShPubPrim	IShPubPrim
ZFS	0.29***	0.30***	0.22**						
	[0.02]	[0.02]	[0.10]						
ZFS/Stu				2.12***	3.27***	6.65***			
				[0.39]	[0.43]	[1.32]			
ZFS/Stu2000							1.46***	2.65***	5.11***
							[0.30]	[0.35]	[1.03]
Dep. Var.	ImPubSec	ImPubSec	ImPubSec	ImPubSec	IShPubSec	IShPubSec	IShPubSec	IShPubSec	IShPubSec
ZFS	0.21***	0.19***	0.29***						
	[0.01]	[0.01]	[0.05]						
ZFS/Stu				3.68***	4.38***	8.69***			
				[0.28]	[0.32]	[1.10]			
ZFS/Stu2000							2.71***	3.59***	6.88***
							[0.21]	[0.25]	[0.79]
Obs.	288	144	144	288	144	144	288	144	144

Notes: All regressions include region dummies and year dummies. Observations are weighted by the year-2000 population in the region. *FB* is the number of foreign-born in the region in each year. *IShPop* is the foreign-born share in the population. ZFB is the predicted stock of foreign-born individuals. Pop and Pop2000 are the year-t and the year-2000 population in the region. *ImPub* are the number of immigrant students in public schools, *ImPubPrim* (*ImPubSec*) are the immigrants enrolled in public schools for primary (secondary) education. *IShPub, IShPubPrim, IShPubSec* are the shares in enrollment for both primary and secondary schooling, for primary schooling only, and for secondary schooling only, respectively. *ZFS* is the predicted number of immigrant students in primary or secondary schools (pooled). *Stu* and *Stu*2000 are the overall enrollment in each year and in year 2000, respectively. All enrollment data is from the Spanish Enrollment Registry. The data by origin country cannot be disaggregated by education level. *** p < .01, ** p < .05, * p < .0.

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