Home Bias in Early Labor Location Choices: Does University Location Matter?*

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Abstract

Geographic mobility facilitates labor-market adjustment, yet home bias can impede relocation and perpetuate regional disparities, including in the provision of essential services. Using administrative records covering the universe of Spanish medical graduates who selected their first job between 1986 and 2019, we examine whether university location shapes early-career mobility. We first document substantial home attachment: more than half of graduates begin their careers in their home province. We then show that university location has a large effect on subsequent location choices. Completing a medical degree in one's home province increases the probability of starting work there by 22 percentage points. The effect of opening a medical school on aggregate home bias depends on the local labor market's capacity to absorb additional graduates. Finally, a nationwide liberalization of university admissions reduced students' propensity to study in their home university district and, through this channel, significantly weakened home bias. These findings show that the location of higher-education institutions and the design of admissions policies can shape the geographic allocation of skilled workers at labor-market entry.

Keywords: Labor mobility, home bias, university location

JEL Codes: J61, I23, J18

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1 Introduction

Geographic mobility is a central mechanism through which labor markets adjust to economic shocks and regional disparities (Beyer and Smets, 2015). When workers are willing and able to relocate, the reallocation of labor across regions can mitigate differences in wages and employment opportunities, thereby improving labor-market efficiency. However, the empirical literature shows that labor mobility is often limited because workers prefer to remain close to their homes, a tendency commonly referred to as *home bias* (Sjaastad, 1962; Kennan and Walker, 2011; Dao et al., 2017). As a result, some regions experience persistent shortages of skilled professionals, while others face an excess supply (Hsieh and Moretti, 2019), generating both efficiency losses and regional inequalities. Beyond these efficiency concerns, when workers provide essential services, such as education or health care, that are universally available in many countries, limited labor mobility and home bias also raise important questions of equity and fairness.

In this paper, we use a comprehensive dataset covering all Spanish medical graduates who begin their careers as physicians to study the existence and evolution of home bias over a 30-year period (1986–2019). We define home bias as the choice to work in an individual's home province. As the main innovation of our analysis, we examine whether studying medicine in one's home province affects the likelihood of taking one's first job as a physician in that province. The Spanish setting provides a unique opportunity to examine this question because entry into the medical profession is governed by a centralized national matching system (*Médico Interno Residente, MIR*), through which candidates choose hospitals and specialties according to their ranking. This institutional structure enables us to systematically and consistently observe graduates' home provinces, the provinces in which they attended university, and the locations of their first employment.

Studying mobility and home bias among physicians in Spain is particularly important because internal mobility has historically been low, regional disparities remain substantial, and health care is provided as a universal public service. First, the internal mobility rate—defined as the ratio of gross interregional migration flows to the working-age population—stood at only 0.3% in 2016, well below the rates observed in countries such as Germany and the United Kingdom, and has remained largely unchanged since the early 2000s (Liu, 2018). Second, labor-market adjustment occurs primarily through changes in participation during expansions and through unemployment and spatial mobility during recessions; nevertheless, substantial regional disparities persist in Spain (Sala and Trivín, 2014). Finally, because health care is universally provided, thousands of recent medical graduates begin their careers each year as resident physicians through a national assignment process known as the MIR system, which is administered centrally by the Ministry of Health (Díez-Rituerto et al., 2026).

In an ideal experiment, prospective doctors would be randomly assigned either to study at a university in their home province or to study elsewhere. Because such an experiment is infeasible, our empirical strategy combines three complementary approaches. First, we analyze the association between studying medicine in one's home province and working there as a resident physician. Second, we exploit quasi-experimental variation generated by the staggered opening of medical schools across provinces, using instrumental-variable and event-study designs. Specifically, we use the availability of a medical school in an individual's home province as an instrument for studying there. We further examine whether these effects vary systematically across province types and individual characteristics, such as gender and ability, thereby shedding light on the mechanisms through which the location of higher education shapes subsequent mobility decisions. Understanding this heterogeneity is crucial for assessing whether higher education policy can influence not only aggregate geographic attachment but also the spatial allocation of top talent and potential gender gaps in mobility. Finally, we study the liberalization of university admissions under the *unique district* reform using before-and-after comparisons of adjacent cohorts.

We document a strong attachment to the home province in graduates' initial job locations: more than 50% take their first job in their home province. This degree of home bias remained relatively stable until 2010, after which it began to decline slightly.

Second, completing a medical degree in one's home province plays a central role in shaping home bias at labor-market entry. Graduates who had a medical school in their home province and studied there are 13% more likely to begin their medical careers in that province. Our instrumental-variable (IV) estimates indicate that studying in one's home province increases the probability of working there as a resident physician by 22 percentage points. The IV estimate is notably larger than its OLS counterpart, consistent with negative selection into attending a home-province university: individuals who choose to study locally despite having the option to leave may have stronger preferences for mobility, thereby biasing the OLS estimate downward. A falsification exercise suggests a modest deviation from the exclusion restriction, so we interpret the magnitude of the estimated effects with caution. Nevertheless, a sensitivity analysis indicates that the qualitative conclusion is robust, and the results remain stable when we control for the number of province-level MIR positions, hospitals, and local student cohorts. Event-study evidence further shows that the opening of a medical school sharply increases the probability of studying in one's home province, but whether this increase anchors graduates there at labor-market entry depends on the local labor market's absorptive capacity. Home bias increases only in provinces where the number of MIR positions can accommodate the additional graduates; elsewhere, the effects are either null or reversed. These findings suggest that the IV estimate may constitute a lower bound on the true effect of studying locally on home bias. We also uncover substantial heterogeneity by gender and ability. Women exhibit greater home bias overall, but the causal effect of studying at a home-province university is smaller for women than for men. Regarding ability, the fact that higher-ranked candidates are more likely to work in their home province suggests that local employment is desirable. Moreover, the stronger response among high-ability candidates indicates that local educational opportunities influence not only individuals with limited outside options but also top-ranked graduates, with potential implications for the regional distribution of medical talent.

Finally, the nationwide liberalization of university admissions under the *unique district* reform reduced students' propensity to study in their home district. Importantly, this effect carried over into the labor market: post-reform cohorts were significantly less likely to remain in their home province when beginning their careers as resident physicians.

We contribute to several strands of the literature. First, despite extensive evidence on the existence and consequences of home bias, much less is known about its origins. Most studies examine migration after individuals have entered the labor market, effectively treating home bias as predetermined (Bayer et al., 2008; Notowidigdo, 2011). We contribute by shifting the analysis to the point of labor-market entry. Our second and most novel contribution is to examine the effects of policies that can influence home bias in the labor market. In particular, we study two policies operating in opposite directions: the opening of medical schools, which makes studying in one's home province more feasible, and the liberalization of university admissions, which expands opportunities to study away from home. Finally, the richness of our dataset allows us to examine heterogeneity not only by gender—a dimension along which previous studies find that women exhibit stronger home bias (Black et al., 2014; Le Barbanchon et al., 2021; Farre and Ortega, 2024; Bisantis, 2025)—but also by ability.

The remainder of the paper is structured as follows. Section 2 describes the institutional setting and data on medical doctors in Spain. Section 3 reports on the empirical analysis. Section 4 concludes.

2 Institutional Setting and Data

2.1 Institutional Setting

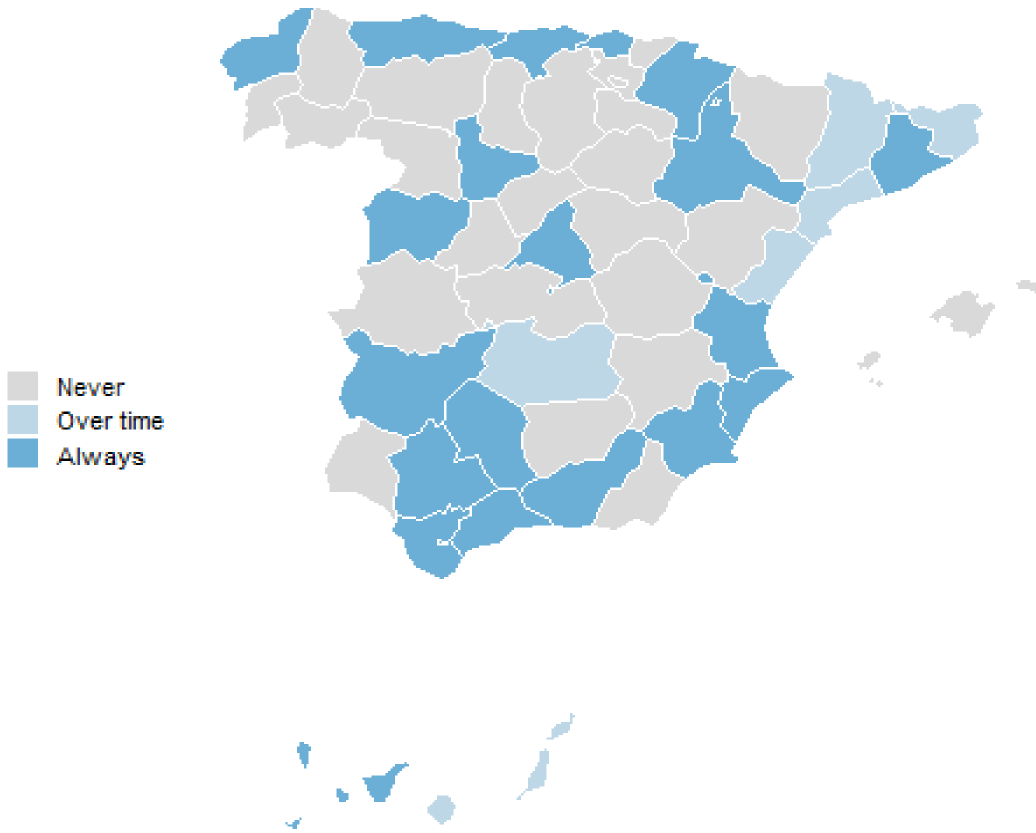
The medical profession in Spain combines a regulated undergraduate medical degree with a centralized allocation process for specialized training in hospitals, during which graduates work as resident physicians. Medicine is offered as a stand-alone undergraduate degree. Unlike in the United States, where students typically complete a general undergraduate degree before entering medical school, Spanish students enroll directly in a medical degree program after completing secondary education. The six-year program integrates theoretical coursework, laboratory training, and clinical rotations in university-affiliated hospitals. Admission has historically required exceptionally high entry grades, making medicine one of the most selective university degrees in Spain.[1] After graduation, candidates enter the *MIR* system, a centralized national procedure that ranks them based on a combination of their score on the *MIR* examination and their undergraduate grade point average. Candidates then select residency positions—each defined by a combination of hospital and specialty—in order of this ranking. Completing *MIR* training is mandatory to practice as a specialist physician in the Spanish health system. The residency period lasts four or five years, depending on the specialty, and constitutes medical graduates' first employment as physicians. For further details on the *MIR* system, see Díez-Rituerto et al. (2026).

Over recent decades, two major institutional changes have reshaped educational opportunities for medical students: (i) the staggered opening of new public and private medical schools, which expanded local access to medical education, and (ii) the liberalization of university admissions under the *unique district* reform. This section describes these institutional changes and explains how they generate the identifying variation used in our analysis.

A central feature of the period under study (1986–2019) was the progressive expansion of medical education across Spanish provinces. In 1986, only 28 medical schools operated in Spain, concentrated in 20 of the country's 50 provinces. Candidates from provinces without a medical school therefore had to leave their home province to study medicine. Over the subsequent decades, several additional medical schools opened—either at newly established universities or through the accreditation of schools at existing institutions—in both the public and private sectors. By 2019, 41 medical schools were operating across 26 provinces. The expansion of private medical schools increased local training capacity in some provinces

[1] According to the Spanish Ministry of Science, Innovation and Universities, medicine has the highest university entry-grade requirements, averaging approximately 12.9 out of 14, with cutoff scores above 13 at most universities. See Ministerio de Ciencia, Innovación y Universidades (2025).

Figure 1. Medical School Presence across Provinces (1986–2019)



Notes: The figure maps Spanish provinces according to medical-school availability during the study period. Dark blue indicates provinces that had a medical school throughout the period; light blue indicates provinces in which the first medical school opened during the period; and gray indicates provinces that never had a medical school.

and provided an additional avenue of access for candidates who did not meet the high cutoff grades required by public universities but could afford private tuition. Based on the timing of medical-school availability, we classify provinces into three mutually exclusive categories: those that had a medical school throughout the study period (20 provinces), those that opened their first medical school during the period (6 provinces), and those that never hosted a medical school (24 provinces), as shown in Figure 1. The analysis exploiting variation in medical-school availability across provinces is presented in Section 3.2.

The authorization of new universities offering medical degrees, whether public or private, falls under the jurisdiction of regional authorities. These openings occurred at different points in time and were driven by province-level policy and accreditation decisions rather than short-term fluctuations in local demand, substantially altering the spatial distribution of educational opportunities. The opening of a medical school represented a discrete change in the opportunity set available to prospective students: it reduced the cost of studying medicine by allowing students to remain in their home province, may have increased the attractiveness of doing so, and altered the composition of entrants by enabling students who would otherwise have been required to leave to enroll locally. Because these openings were staggered across provinces and years, they provide quasi-exogenous variation in the availability of local medical education. We exploit this variation in both our instrumental-variable strategy—in which studying at a local university is instrumented with the availability of a medical school in the home province—and our event-study design, in which the opening of a medical school constitutes the event. Section 3.3 presents this analysis.

A second institutional change expanded students' outside options at the point of university entry. Before 2001, admission to public universities in Spain was largely governed by *district*-based restrictions that confined most students to universities in their home province or in small groups of administratively and geographically connected provinces. For empirical purposes, we refer to these groupings as *districts*: clusters of provinces that historically shared coordinated admissions systems and integrated higher-education markets. Appendix Table A.1 reports the complete mapping of provinces to these historical districts.

Therefore, before 2001, although a limited number of candidates were permitted to enroll outside their district—typically corresponding to between 5 and 10 percent of available places—these opportunities were scarce and highly competitive and did not substantially relax the geographic constraints facing the vast majority of prospective medical students. Consequently, interdistrict mobility at university entry was rare, regardless of the field of study. For example, in the 1998–1999 academic year, only 4.8 percent of university students completed their degrees outside their region of origin (Rahona López and Angoitia Grijalba, 2007).

The *unique district* reform fundamentally altered this system by introducing a national admissions framework beginning in the 2001–2002 academic year.[2] The reform was phased in: 20 percent of places were opened to applicants nationwide in 2001–2002, 50 percent in 2002–2003, and 100 percent by 2003–2004, with province-level authorities retaining some discretion during the transition. By the 2003–2004 academic year, candidates could apply to any medical school in Spain, substantially expanding their geographic choice set and reducing institutional barriers to studying outside their home district.

The reform therefore provides a complementary source of variation that operates through

[2]This reform is commonly referred to in Spain as the *unique district*. It was implemented by Royal Decree 69/2000, available here.

a different channel from the opening of new medical schools. Whereas these supply-side expansions increased opportunities to study locally, the *unique district* reform expanded opportunities to study elsewhere, thereby weakening the link between students' home provinces and education locations. We present this analysis in Section 3.4.

2.2 Data and Descriptive Statistics: 1986-2019

Our empirical analysis uses an unusually rich administrative dataset, provided by the Spanish Ministry of Health, covering the universe of candidates who obtained a residency position through the *MIR* system between 1986 and 2019. The dataset contains detailed information on each candidate's gender, total *MIR* score, and rank. Most importantly, it records the complete sequence of relevant locations: the home province or province of origin, proxied by the province of the candidate's home address when registering for the *MIR* examination; the province in which the candidate obtained a medical degree; and the province of the residency position.[3] To our knowledge, this is the most comprehensive dataset assembled to date on the educational and early-career trajectories of Spanish physicians. It allows us to examine how province-level educational opportunities shape geographic mobility at labor-market entry. Table 1 reports annual summary statistics for the main variables used in the analysis over the 1986–2019 period.

Across cohorts, the probability of studying medicine in one's home province ranges between 62% and 76% but follows a clear downward trend. It is approximately 75% in the late 1980s, remains relatively stable throughout the 1990s, and begins to decline in the early 2000s, reaching approximately 62% by the end of the sample period. A similar trajectory emerges for our main measure of home bias, defined as the probability of taking up a residency position in one's home province. This probability is close to 60% for the earliest cohorts, fluctuates between 55% and 60% throughout the 1990s, and then declines gradually during the 2000s and 2010s, reaching approximately 47–49% for the most recent cohorts.

The share of female candidates increases markedly over time, from approximately 40% in 1986 to around 65% in 2019, peaking above 70% during the late 2000s and early 2010s. Appendix Table A.2 reports descriptive statistics by gender. Men attain a slightly better average *MIR* rank in almost every year. Women are only slightly more likely to study in their home province, but they exhibit substantially greater home bias in their *MIR* placement choices. These gender differences remain stable over time and point to systematic differences

[3] Martinez Miera and Sunyer (2025) also use candidates' home addresses at the time of registration for the *MIR* examination as a proxy for their province of origin.

Table 1. Descriptive Statistics by Year

Year	Prop. Female	Prob. Studying in Home Province	Prob. Doing *MIR* in Home Province (Home bias)
1986	0.406	0.761	0.613
1987	0.435	0.754	0.610
1988	0.457	0.742	0.610
1989	0.453	0.743	0.552
1990	0.451	0.741	0.541
1991	0.475	0.733	0.553
1992	0.501	0.724	0.573
1993	0.474	0.731	0.602
1994	0.520	0.723	0.604
1995	0.538	0.729	0.556
1996	0.585	0.742	0.604
1997	0.655	0.732	0.596
1998	0.588	0.733	0.591
1999	0.636	0.739	0.574
2000	0.640	0.722	0.534
2001	0.672	0.707	0.565
2002	0.667	0.689	0.583
2003	0.667	0.695	0.576
2004	0.680	0.683	0.596
2005	0.686	0.678	0.630
2006	0.683	0.676	0.617
2007	0.704	0.671	0.634
2008	0.718	0.677	0.598
2009	0.703	0.670	0.589
2010	0.706	0.647	0.597
2011	0.711	0.653	0.570
2012	0.722	0.641	0.551
2013	0.713	0.644	0.526
2014	0.684	0.625	0.486
2015	0.679	0.628	0.485
2016	0.665	0.625	0.481
2017	0.650	0.617	0.468
2018	0.655	0.622	0.471
2019	0.662	0.619	0.470
Mean Values	0.626	0.683	0.558

Notes: This table presents descriptive statistics by year. *Prop. Female* is the proportion of female candidates. *Prob. Studying in Home Province* refers to the share of candidates who studied medicine in their province of origin. *Prob. Doing MIR in Home Province* measures home bias: the share of candidates who did their intern position in their province of origin.

in examination performance and geographic mobility. We examine these patterns further in the heterogeneity analysis presented in Section 3.2.3.

2.3 Home Bias over Time

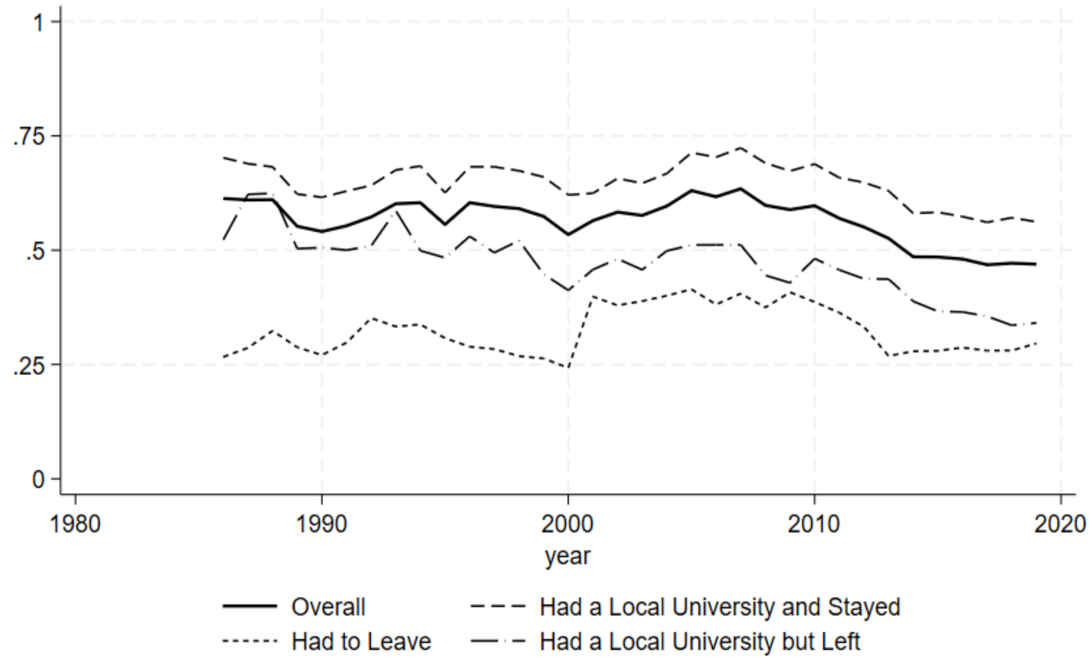
We classify candidates according to (i) whether their home province offered a medical degree and (ii) whether they studied in that province. This classification yields three groups: candidates whose home province had a medical school and who studied there (*Had a Local University and Stayed*); candidates whose home province had a medical school but who pursued their medical degree elsewhere (*Had a Local University but Left*); and candidates whose home province did not have a medical school and who therefore had to leave to study medicine (*Had to Leave*).

Two qualifications are important. First, this classification does not represent exogenous assignment, as it partly reflects individual choices. Second, the absence of a medical school in the home province may have affected earlier educational decisions, including whether to pursue medicine rather than another field of study. In the presence of strong home bias, for example, some individuals may choose not to study medicine if doing so requires leaving their province. Because our data include individuals only once they enter the national residency matching process—that is, only after they have completed medical training—we cannot directly observe or model selection into the medical profession itself. Our analysis therefore conditions on the population of medical graduates and examines how local educational opportunities shape where they study and subsequently undertake residency training, rather than whether they pursue medical education. Importantly, although this conditioning may affect the composition of the observed sample, it does not mechanically generate the within-sample differences in mobility patterns that we document across the three groups. These differences reflect observed location choices among individuals who do enter medicine.

Figure 2 plots the evolution of overall home bias—defined as the share of candidates who take up a residency position in their home province—between 1986 and 2019, together with the corresponding series for the three groups defined above. The overall series indicates a high and relatively stable degree of home bias: throughout the period, more than half of all candidates undertake their residency in their home province. However, this aggregate pattern masks substantial and persistent heterogeneity across the three groups.

Candidates whose home province had a medical school and who studied there (*Had a Local University and Stayed*) consistently exhibit the highest levels of home bias: between 60% and 75% undertake their residency in their home province. By contrast, candidates whose

Figure 2. Evolution of Home Bias over Time by Educational Opportunities



Notes: The figure shows the evolution of average home bias between 1986 and 2019 for the full sample, and separately for three groups. *Had a Local University and Stayed* refers to individuals from a province that offered a medical degree at the time of university entry who studied in their home province. *Had a Local University but Left* includes individuals from a province with a local medical school who nevertheless studied in another province. *Had to Leave* refers to individuals from a province without a medical school, who were therefore required to leave their home province in order to pursue medical studies.

home province had a medical school but who studied elsewhere (*Had a Local University but Left*) exhibit substantially lower levels of home bias, with only 40% to 50% taking up a residency position in their home province. The lowest levels of home bias are observed among candidates from provinces without a medical school (*Had to Leave*), whose rates of residency in the home province range between 25% and 35%. The figure provides no evidence of convergence among the three groups over time.

3 Empirical Analysis

3.1 Association between University Province and Home Bias

We begin by examining the association between candidates' educational opportunities, their university locations and their likelihood of selecting a residency position in their home province.

Table 2. Baseline OLS Estimates of Home Bias

	Home Bias (1)	Home Bias (2)	Home Bias (3)	Home Bias (4)
Had to Leave	-0.3159***	-0.3130***	-0.3134***	-0.1012***
	(0.0032)	(0.0032)	(0.0031)	(0.0136)
Had a Local University but Left	-0.2022***	-0.1934***	-0.1874***	-0.1342***
	(0.0039)	(0.0039)	(0.0039)	(0.0042)
Female		0.0354***	0.0162***	0.0083***
		(0.0026)	(0.0026)	(0.0024)
Rank Position		-0.1237***	-0.2532***	-0.1712***
		(0.0049)	(0.0068)	(0.0060)
Constant	0.6437***	0.6557***	0.7451***	0.6653***
	(0.0015)	(0.0025)	(0.0179)	(0.0170)
Observations	149,642	149,642	149,642	149,642
R-squared	0.0680	0.0730	0.1200	0.2678
Year and Specialty FE			Yes	Yes
Province-level FE			No	Yes

Notes: The table shows the OLS estimated coefficients of home bias (takes the value of 1 when the candidate chooses the medical intern position in the same province as the province of origin). *Female* takes the value of 1 if the candidate was female. *Rank position* is the candidates rank (1 = best score) normalized by the maximum rank in that edition. *Had a Local University and Left* takes the value of 1 if the candidate's home had a medical school and did not study there. Province-level fixed effects include home, university, and hospital province fixed effects. Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1

Table 2 presents our baseline estimates. The main regressors are indicators that capture both the educational opportunities available in a candidate's home province and the candidate's study-locations in response to those opportunities. With *Had a Local University and Stayed* as the omitted category, the coefficients on *Had a Local University but Left* and *Had to Leave* measure differences in home bias relative to candidates who had access to medical education in their home province and studied there. Column 1 reports a specification without controls, while column 2 adds individual characteristics, including gender and rank position. Column 3 adds specialty and year fixed effects, and column 4 adds province-level fixed effects for the province of origin, university province, and hospital province.

The coefficients on *Had to Leave* and *Had a Local University but Left* are consistently negative, large in absolute magnitude, and statistically significant across all specifications. The results remain highly robust to the inclusion of year and medical-specialty fixed effects in column 3 and province-level fixed effects in column 4. The estimates change most when province-level fixed effects are introduced, particularly those for the home province. This pattern suggests that candidates in these two groups are disproportionately concentrated in a small number of provinces with distinct characteristics. Focusing on the most comprehensive specification in column 4, candidates in the *Had to Leave* group are approximately 10 percentage points less likely to select a residency position in their home province than candidates in the *Had a Local University and Stayed* group. Candidates in the *Had a Local University but Left* group are approximately 13 percentage points less likely to do so. Although we reject the null hypothesis that the two coefficients are equal in magnitude (p-value = 0.01), this statistical rejection largely reflects the precision of the estimates; the difference in magnitude is economically small. Given their similar magnitudes, in the subsequent analysis we combine these two groups and use the pooled group as the omitted category.

Turning to the other variables of interest, the coefficient on *Female* is positive and statistically significant, although modest in magnitude, indicating that women are slightly more likely to select a residency position in their home province. The coefficient on *Rank Position* is negative. Because a lower numerical rank denotes a better-ranked candidate, this estimate indicates that better-ranked candidates exhibit greater home bias than their worse-ranked counterparts. These revealed-preference patterns suggest that proximity to the home province is a desirable attribute of residency positions.

3.2 Causal Effect of Studying Locally on Home Bias: IV Analysis

The descriptive evidence presented above reveals a strong and statistically significant association between studying at a university in one's home province and subsequent home bias: candidates who complete their medical degree in their home province are more likely than those who study elsewhere to remain there for their residency. Although informative, this association cannot be interpreted causally because study location may be endogenous. Specifically, where individuals study may be correlated with unobserved factors that also affect where they choose to undertake their residency. To address this endogeneity concern, we proceed as follows.

First, we exploit quasi-experimental variation in the availability of a medical school in candidates' home provinces. Some provinces had a medical school throughout the study period, others never had one, and several experienced the opening of their first medical school between 1986 and 2019. We use an instrumental-variable approach in which studying medicine in one's home province is instrumented with the availability of a medical school in that province. This strategy builds on the use of geographic proximity to educational institutions as an instrument in the returns-to-education literature, as in Card (1993), and allows us to identify the causal effect of studying in the home province on subsequent home bias.

Second, we examine heterogeneity along two dimensions: province type and individual characteristics, specifically gender and ability.

3.2.1 Instrumental Variable Approach: Provinces with Medical Schools

As a source of exogenous variation, we use the availability of a medical school in an individual's home province—an instrument that is plausibly unrelated to unobserved mobility preferences but strongly predicts studying locally.

Table 3 reports the results. To contextualize our instrumental-variable strategy, we also present three conventional estimations. The treatment is studying medicine in one's home province, and the instrument—or assignment variable—is the availability of a medical school in that province. Column 1 reports the *As-Treated (AT)* estimate. This specification compares candidates according to their actual study location, irrespective of assignment status, and is therefore likely to be biased by selection into treatment. It corresponds to the specification in column 4 of Table 2, except that the omitted category pools all candidates who did not study in their home province, either because no medical school was available there or because they studied elsewhere despite having access to one. Column 2 reports the *Per-Protocol (PP)* estimate, which restricts the sample to candidates whose study location is consistent with their assignment status. Specifically, it excludes candidates who had a medical school in their home province but studied elsewhere. Because compliance may itself be endogenous, this specification is also potentially affected by selection bias. Column 3 reports the *Intention-to-Treat (ITT)* estimate, which compares candidates solely according to assignment status—that is, whether their home province had a medical school. If assignment is exogenous, this specification identifies the causal effect of access to local medical education, although it does not recover the effect of actually studying locally. Together, these estimates illustrate the limitations of naïve comparisons and motivate our *Instrumental-Variable (IV)* strategy, which isolates the variation in study location induced by local medical-school availability and identifies a causal effect for compliers under the standard identification assumptions. Column 4 reports the first-stage estimate of the effect of having a medical school in the home province on the probability of studying medicine there. The coefficient is positive and highly statistically significant, confirming the strength of the instrument and ruling out weak-instrument concerns. Finally, column 5 reports the IV estimate for our main research question: whether studying medicine in one's home province, instrumented with the availability of a medical school there, increases the probability of selecting a first residency position in that province.

We now turn to the estimated effect of studying at a home-province university on home

Table 3. Instrumental Variable Regression Results

	As-Treated	Per Protocol	ITT	First Stage	IV
	Home Bias	Home Bias	Home Bias	Had Local U and Stayed	Home Bias
	(1)	(2)	(3)	(4)	(5)
Had a Local	0.1322***	0.0668**			0.2202***
Univ. and Stayed	(0.0042)	(0.0263)			(0.0495)
Female	0.0083***	0.0076***	0.0085***	0.0021	0.0081***
	(0.0024)	(0.0025)	(0.0024)	(0.0016)	(0.0024)
Rank Position	-0.1712***	-0.1765***	-0.1783***	-0.0535***	-0.1665***
	(0.0060)	(0.0063)	(0.0061)	(0.0038)	(0.0066)
Prov. with University			0.0898***	0.4079***	
			(0.0203)	(0.0173)	
Constant	0.5335***	0.5882***	0.5626***	0.4903***	0.4546***
	(0.0174)	(0.0316)	(0.0265)	(0.0210)	(0.0475)
Observations	149,642	130,419	149,642	149,642	149,642
R-squared	0.2678	0.2823	0.2621	0.6264	0.2652
Year and Specialty FE	Yes	Yes	Yes	Yes	Yes
Province-level FE	Yes	Yes	Yes	Yes	Yes

Notes: The table shows the OLS and IV estimation of home bias (takes the value of 1 when the candidate chooses the medical intern position in the same province as the province of origin). *Female* takes the value of 1 if the candidate was female. *Rank position* is the candidate's rank (1 = best score) normalized by the maximum rank in that edition.*Had a Local University and Left* takes the value of 1 if the candidate's home had a medical school and did not study there. Province-level fixed effects include home, university, and hospital province fixed effects. Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

bias at labor-market entry. Across all specifications, studying locally is strongly and positively associated with the probability of undertaking residency in the home province. Column 1 reports the naïve *As-Treated* comparison: candidates who studied in their home province are 13.2 percentage points more likely to work there. Because study location reflects an individual choice, this estimate combines the causal effect of studying locally with selection into treatment. Column 2 restricts the sample to candidates whose observed study decisions are consistent with their assignment status, thereby excluding those in the *Had a Local University but Left* group. The resulting *Per-Protocol* estimate is smaller, at 6.7 percentage points. This attenuation reflects selection into compliance, as candidates who comply with their assignment may differ systematically from those who do not. Column 3 reports the *Intention-to-Treat (ITT)* estimate, which compares candidates solely according to whether their home province had a medical school, regardless of where they actually studied. The resulting reduced-form effect is 8.9 percentage points and captures the average effect of access to local medical education on home bias. Column 4 shows that the availability of a medical school in the home province increases the probability of studying locally by 40.8 percentage points, confirming that the instrument strongly predicts study location. Finally, and most importantly, column 5 reports the IV estimate. The coefficient on *Had a Local University and Stayed* indicates that studying at a home-province university increases the probability of selecting a residency position in that province by 22 percentage points. This estimate is substantially larger than the as-treated estimate in column 1, suggesting that the latter is biased downward. We interpret this difference as evidence of negative selection into studying locally. Candidates who leave their province to study medicine also appear to have stronger preferences for geographic mobility and are therefore more likely to work outside their home province. Because noncompliance is one-sided—candidates whose home province lacks a medical school cannot study locally—the treated group consists of compliers. Under the standard assumptions, the IV estimation can therefore also be interpreted as the average treatment effect on the treated.

Across all specifications, female candidates exhibit greater home bias and are more likely to select a residency position in their home province. Better-ranked candidates also display greater home bias, indicating that proximity to home is a desirable attribute of *MIR* positions.

As a first robustness check, we assess the exclusion restriction using a falsification test. If the instrument affects the outcome only through treatment, it should have no effect on the work locations of untreated candidates—those in the *Had a Local University but Left* and *Had to Leave* groups (see Appendix Table A.3). Within this sample, the estimated effect of the instrument is statistically significant but modest in magnitude, at approximately 5 percentage points, indicating that the exclusion restriction may not perfectly hold. The instrument may therefore affect home bias through channels other than study location, such as local ties, mobility costs, or unobserved preferences. Nevertheless, the magnitude of this deviation is small relative to the baseline IV estimate of 22 percentage points. Instrument relevance is not a concern in our setting, given the strong first-stage relationship between medical-school availability and study location documented in Table 3.

To assess the implications of this potential deviation from the exclusion restriction, we implement the plausibly exogenous instrumental-variable sensitivity analysis proposed by Conley et al. (2012) and implemented by Clarke and Matta (2018). Specifically, we allow the instrument to have a direct effect on the outcome of up to 5 percentage points in either direction. Under this calibration, the bounds on the causal effect range from approximately 0.001 to 0.441. These bounds indicate that our qualitative conclusion—namely, that the effect is non-negative—survives violations of the exclusion restriction of the magnitude suggested by the falsification subgroup. However, the lower bound's proximity to zero underscores the greater uncertainty surrounding the magnitude of the causal effect once modest violations are permitted.

Second, our main findings are robust to controlling for province-level residency-training capacity and cohort size. Both the baseline IV estimates and the conclusions from the falsification exercises remain qualitatively unchanged when we control for (i) the ratio of *MIR* positions offered in a province to the number of medical graduates from that province (*Local Assimilation Ratio*), (ii) the number of hospitals, and (iii) the number of available medical specialties (see Appendix Table A.3).

In sum, the IV analysis provides causal evidence that completing a medical degree in one's province of origin substantially increases the probability of taking up a first residency position there, reinforcing the view that the location of higher education shapes longer-term geographic attachment. Although the falsification test detects a statistically significant but modest deviation from the exclusion restriction, the estimated direct effect is small relative to the IV estimate, suggesting that any resulting bias is unlikely to overturn our central conclusion.

3.2.2 Heterogeneity by Province Type-level Context

Spanish provinces differ substantially in their educational capacity, economic structure, and labor-market opportunities. Provinces in which new medical schools opened during the study period tend to be more dynamic, with larger health systems, faster population growth, and greater fiscal capacity. These provinces may therefore also offer better employment prospects for medical graduates. This creates an important identification concern: the apparent anchoring effect of local medical education may reflect not education itself, but the underlying attractiveness of more prosperous provinces that both establish medical schools and retain physicians.

If this were the case, one would expect the causal effect of studying locally on home bias to be concentrated in provinces that opened new medical schools and experienced expanding labor-market opportunities. By contrast, in provinces without such educational expansion—often more peripheral or economically constrained areas—local study opportunities might play a weaker role in shaping geographic attachment.

To assess whether the anchoring effect of local medical education reflects province-level economic conditions rather than educational choices per se, we replicate the IV analysis separately for two groups: (i) provinces in which no new medical school opened during the study period and (ii) provinces in which at least one new medical school opened. This comparison

Table 4. Heterogeneous Effects by Provinces With and Without New Medical Schools

Panel A. Provinces that Never Opened a New University

	Home bias As treated (1)	Home bias Per Protocol (2)	Home bias ITT (3)	Home bias IV (4)
Had a Local University and Stayed	0.1216***	0.3186***		0.2146***
	(0.0058)	(0.0260)		(0.0050)
Female	0.0032	0.0020	0.0034	0.0026
	(0.0029)	(0.0031)	(0.0029)	(0.0029)
Rank Position	-0.1778***	-0.1831***	-0.1874***	-0.1772***
	(0.0075)	(0.0079)	(0.0076)	(0.0076)
Prov. with University			0.1624***	
			(0.0038)	
Constant	0.5927***	0.3893***	0.5389***	0.5089***
	(0.0212)	(0.0330)	(0.0211)	(0.0210)
Observations	88,765	77,446	88,765	88,765
R-squared	0.3734	0.3754	0.3481	0.3543

Panel B. Provinces that Opened at Least One New University

	Home bias As treated (1)	Home bias Per Protocol (2)	Home bias ITT (3)	Home bias IV (4)
Had a Local University and Stayed	0.1335***	-0.0615		0.2897***
	(0.0077)	(0.0480)		(0.0788)
Female	0.0038	0.0054*	0.0039	0.0036
	(0.0026)	(0.0028)	(0.0026)	(0.0027)
Rank position	-0.0586***	-0.0574***	-0.0635***	-0.0530***
	(0.0066)	(0.0067)	(0.0066)	(0.0071)
Prov. with University			0.0760***	
			(0.0208)	
Constant	0.0575**	0.4006***	0.0224	-0.2760***
	(0.0272)	(0.0878)	(0.0345)	(0.0278)
Observations	60,877	52,973	60,877	60,877
R-squared	0.5882	0.5951	0.5841	0.5822

Notes: The table shows the OLS and IV estimation of home bias (takes the value of 1 when the candidate chooses the medical intern position in the same province as the province of origin). *Female* takes the value of 1 if the candidate was female. *Rankposition* is the candidate?s rank normalized by the maximum rank in that edition. *Had a Local University and Stayed* takes the value of 1 if the candidate's home had a medical school and studied there. All regressions include year and speciality fixed effects as well as province-level fixed effects which include home, university, and hospital province fixed effects. Robust standard errors in parentheses. *** p<0.01, **p<0.05, * p<0.1.

allows us to determine whether the estimated effect is concentrated in economically dynamic provinces or also arises in provinces without similar characteristics.

The results in Table 4 show that the estimated effect is present and similar in magnitude across both types of provinces. In provinces in which no new medical school opened, studying locally increases the probability of selecting a residency position in the home province by approximately 21.5 percentage points. In provinces in which a new medical school opened, the effect remains positive and is slightly larger, at 29 percentage points.

This heterogeneity analysis is particularly informative in light of the falsification results. Because we detect a small but statistically significant association between the instrument and residency location even within a subgroup for which the *study locally* channel is muted, one could argue that the instrument partly captures province-level attractiveness—such as economic dynamism, expanding health systems, or broader labor-market opportunities—that retains residents regardless of where they studied. However, the estimated effect is present and similar in magnitude across both types of provinces, a pattern that is difficult to reconcile with a purely prosperity-based pull mechanism. Such a mechanism would predict a substantially weaker or nonexistent effect in nonexpanding, more peripheral provinces. Instead, the results point to a genuine anchoring effect of local medical education that is not solely explained by province-level conditions and is reinforced, rather than weakened, when new medical schools create locally trained cohorts where none previously existed.

Taken together, the results support the interpretation that studying at a university in one's home province has an anchoring effect across diverse regional contexts, while remaining consistent with the possibility—highlighted by the falsification test—that province-level attractiveness also influences initial residency-location choices at the margin..

3.2.3 Heterogeneity by Candidates' Characteristics: Gender and Ability

This section examines whether the causal effect of studying in one's home province on home bias at labor-market entry varies with candidates' characteristics. We focus on two dimensions that are central to labor-market allocation and mobility decisions: gender and ability, with the latter proxied by the candidate's standardized rank on the *MIR* examination.

Examining heterogeneity by gender is particularly relevant in this setting. A large literature documents systematic gender differences in geographic mobility and labor-market outcomes, even among highly educated workers. Women are less likely to migrate in response to job opportunities, particularly when relocation decisions require household coordination or depend on a partner's circumstances (Black et al., 2014). More recent evidence shows that these mobility frictions affect job-search behavior and matching outcomes, leading women to sort differently across locations and employers (Le Barbanchon et al., 2021; Bisantis, 2025). These mechanisms are likely to be especially salient in medicine, where training paths are long, geographically rigid, and closely intertwined with family location. This spatial rigidity may interact with gender in consequential ways: although women account for 62% of physicians in the Spanish National Health System, only 42% of heads of service are women (WOMEDS, 2026). From this perspective, studying in one's home province may have different effects for men and women. On the one hand, local study opportunities may disproportionately benefit women by relaxing mobility constraints and enabling them to pursue medical careers without relocating. On the other hand, if women's location choices are driven primarily by noneducational factors—such as family ties or a partner's employment—studying locally may play a more limited causal role in determining where they eventually work. Distinguishing between these channels is important for assessing whether education policies can meaningfully affect gender gaps in the spatial allocation of labor.

We also examine heterogeneity by ability because mobility and geographic sorting depend strongly on candidates' outside options. High-ability candidates typically face better opportunities nationwide and may therefore be less geographically constrained, whereas lower-ability candidates may benefit more from local educational and professional pathways. Estimating how the effect of studying in one's home province varies across the ability distribution allows us to determine whether local study primarily anchors candidates with weaker outside options or also affects top performers.

Table 5 reports the results. Both analyses use the baseline IV specification, in which studying medicine in one's home province (*Had a Local University and Stayed*) is instrumented with the availability of a medical school in the candidate's province of origin. In the heterogeneity specifications, the endogenous interaction terms are instrumented with the corresponding interactions between medical-school availability and the relevant candidate characteristic. The estimates reveal substantial heterogeneity in the causal effect of studying locally on home bias along both dimensions: gender and ability.

Column 1 examines heterogeneity by gender. For male candidates, studying in the home province increases home bias by approximately 25 percentage points. Although female candidates exhibit greater home bias overall, the causal effect of studying locally remains large but is somewhat smaller than for men. The interaction coefficient indicates that the effect is approximately 5 percentage points smaller for women, implying a net effect of about 20 percentage points.

Column 2 examines heterogeneity by ability, proxied by the standardized *MIR* rank po-

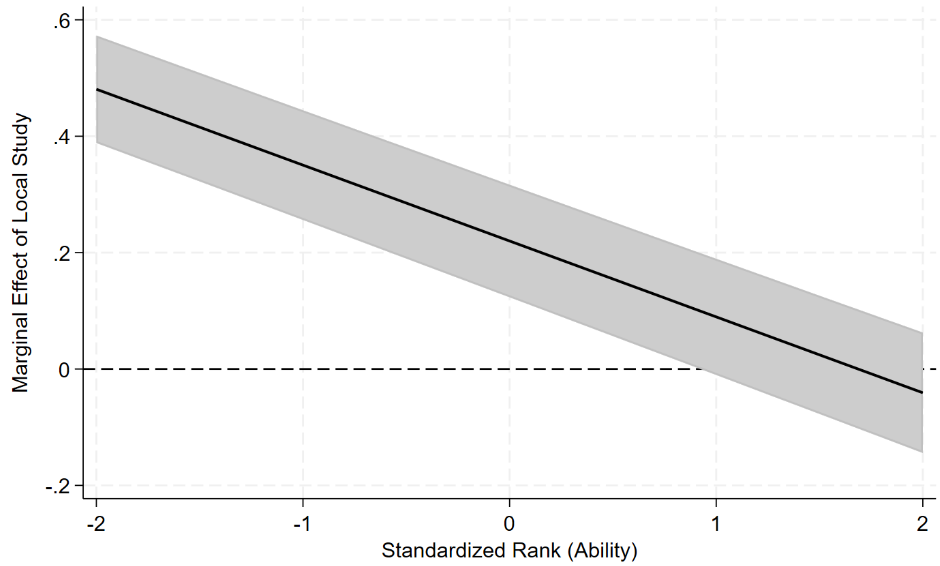
Table 5. Heterogeneity in IV Estimates on Home Bias: Gender and Ability

	Home Bias	
	Gender	Ability
	(1)	(2)
Had a Local University and Stayed	0.2477***	0.3344***
	(0.0499)	(0.0477)
Female	0.0412***	0.0067***
	(0.0058)	(0.0024)
Had a Local University and Stayed × Female	-0.0485***	
	(0.0075)	
Rank position	-0.1681***	0.1828***
	(0.0066)	(0.0121)
Had a Local University and Stayed × Rank position		-0.5064***
		(0.0156)
Constant	0.4375***	0.3911***
	(0.0477)	(0.0464)
Observations	149,642	149,642
R-squared	0.2654	0.2712
Year and Specialty FE	Yes	Yes
Province-level FE	Yes	Yes

Notes: The table shows IV estimates of home bias (takes the value of 1 when the candidate chooses the medical intern position in the same province as the province of origin). *Female* is standardized by year. *Rank position* is the candidate's rank (1 = best score) normalized by the maximum rank in that edition. Province-level fixed effects include home, university, and hospital province.Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

sition. The interaction between studying in the home province and rank position is large, negative, and statistically significant, indicating that the effect of studying locally declines as the numerical rank increases—that is, as candidate ability decreases. To illustrate this pattern, Figure 3 plots the marginal effect of studying locally across the distribution of standardized rank values. The marginal effect is largest and statistically significant among high-performing candidates, who have low *MIR* rank values, and gradually declines as ability decreases. Among candidates with below-average *MIR* performance—that is, those with higher rank values—the effect of studying locally becomes statistically indistinguishable from zero.

Figure 3. Marginal Effect of Studying Locally on Home Bias, by Standardized *MIR* Rank



Notes: The figure plots the marginal effect of studying in the home province on home bias across the distribution of rank position, obtained from the IV specification that interacts studying locally with rank position (see Column 2 of Table 5). The solid line reports the point estimate and the shaded area the 95% confidence interval.

This pattern highlights the importance of individual ability in shaping geographic responses to educational opportunities. High-ability candidates typically have stronger outside options and broader geographic choice sets in the national matching process. For these candidates, studying locally appears to play a decisive role in anchoring them to their province of origin, suggesting that local universities help retain top talent that might otherwise relocate. By contrast, lower-performing candidates with worse numerical ranks may face tighter constraints in the *MIR* allocation process. Their residency locations may therefore respond less to where they trained and more to the availability of unfilled positions across provinces.

These findings qualify the interpretation of local education policy. Rather than merely retaining candidates who would have remained in their home province regardless, expansions in local training capacity appear to have the greatest influence on the location choices of the highest-performing candidates. The opening of new medical schools may therefore not only expand access to medical education but also improve the local retention of top-ranked graduates, with important implications for workforce planning and geographic equity in access to medical talent.

3.3 Causal Effect of Medical School Openings on Home Bias: Event study

We extend our causal analysis using an event-study design that traces the dynamic effects of medical-school openings on home bias. The question addressed by this analysis differs slightly from that examined using IV. Whereas the IV estimates identify the average causal effect of studying locally on home bias, the event-study framework examines whether opening a medical school causally increases the local retention of medical residents and through which mechanisms. It also allows us to test for differential pre-trends between provinces that eventually open a medical school and those that do not, trace how the effects evolve over time, and assess whether they persist across subsequent cohorts.

Our approach examines two sequential location decisions along the medical career path. First, we study home bias in the *educational location* by estimating how the opening of a medical school in a candidate's home province affects the probability of completing a medical degree there. Second, we study home bias in the residency-location decision by estimating how the opening of a medical school in the home province affects the probability of taking up a *MIR* residency position in that province.

A key clarification concerns the timing of the event. Our data do not record when candidates begin medical school. Instead, we observe them when they take the *MIR* examination, together with the university from which they obtained their medical degree. We therefore define event time relative to the first year in which the *MIR* data include a candidate who graduated from a newly opened medical school in a given province. Event time should thus be interpreted as indexing cohorts relative to the first observable exposure to medical education in our data, rather than the number of calendar years since the institution formally opened. Importantly, we observe candidates' outcomes only after their exposure to treatment.[4]

We estimate the following specification:

$$\begin{aligned} y_{irt} = \alpha + \beta_1 \text{Female}_i + \beta_2 \text{RankPosition}_i + \sum_{k \neq -1} \delta_k \mathbf{1}(t - T_r = k) \\ + \gamma_r + \lambda_t + \mu_u + \sigma_s + \rho_h + \varepsilon_{irt}, \end{aligned} \tag{1}$$

[4]In this regard, we classify Lleida, Tarragona, and Las Palmas as having access to local medical education throughout the study period. Although independent medical schools were formally established in these provinces at different points in time, our data include *MIR* candidates who completed their medical degrees there from the beginning of the sample period. This reflects the presence of medical schools or satellite campuses affiliated with universities headquartered in other provinces, which effectively allowed students to complete their medical studies locally. From the perspective of students' educational choices, these provinces therefore offered a local option throughout the period under study.

where y_{irt} is an indicator for the home bias in either (i) completing the medical degree in the home province (educational location) or (ii) selecting an intern position in the home province (residency-location) for individual i from home province r in year t. The indicators $\mathbf{1}(t - T_r = k)$ capture event time relative to T_r, defined as the first year in which graduates from the newly opened medical school in province r appear in the *MIR* data. We normalize the coefficient for $k = -1$ to zero. All specifications include year fixed effects (λ_t). Home-province fixed effects (γ_r), and university-province fixed effects (μ_u), medical specialty fixed effects (σ_s) and hospital fixed effects (ρ_h) are included only when studying residency-location home bias.

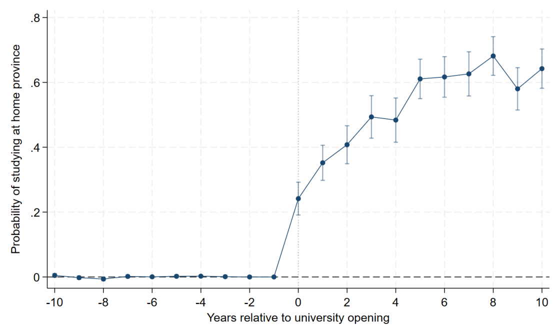
We estimate equation (1) separately for two groups of provinces: those without a medical school that subsequently opened their first one (*first-school*) and those that already had at least one medical school and subsequently opened another (*additional-school*).

Figure 4 presents the event-study estimates of home bias in educational and residency-locations, separately for the two types of provinces. Focusing first on the educational-location, the opening of a medical school produces a clear and sustained increase in the probability that candidates complete their medical degree in their home province. For both types of provinces, the pre-event coefficients are flat, providing no evidence of differential pre-trends. The response is particularly pronounced in first-school provinces: the coefficients remain close to zero throughout the pre-event window and increase sharply at $t - T_r = 0$, by approximately 25 percentage points. They continue to rise thereafter and stabilize at around 60–65 percentage points in later years. In additional-school provinces, the post-event increase is also positive and statistically significant, although more gradual because a local option was already available to some candidates before the new school opened. These results provide a useful validation of the event-study design: opening the first medical school in a province enables candidates from that province to graduate locally, whereas opening an additional school increases the share of candidates who do so.

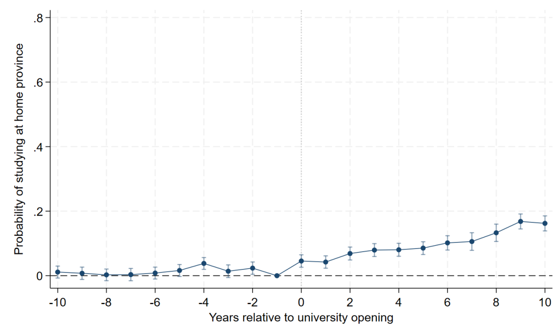
More importantly, and central to this section, the work-location estimates differ markedly across the two types of provinces. In *first-school* provinces, the estimated effect on home bias is statistically indistinguishable from zero throughout the pre- and post-event periods. Thus, despite the large and clear response in educational-locations, the opening of a province's first medical school does not translate into a detectable increase in the local retention of *MIR* residents. The results for *additional-school* provinces are more nuanced. Panel (d) shows a transitory positive effect on home bias that peaks at approximately $t - T_r = 0$, dissipates thereafter, and becomes negative for later cohorts. Importantly, however, the pre-event coefficients are not flat and exhibit a declining trend before the opening. This pattern raises a natural concern: the positive coefficients between $t - T_r = 0$ and $t - T_r = 2$ may partly capture a rebound from the pre-event decline rather than a purely causal response to the opening of the additional school. We therefore interpret the post-event estimates with this caveat in mind.

Figure 4. Event Study Estimates of the Effect of a Medical School Opening on Home Bias

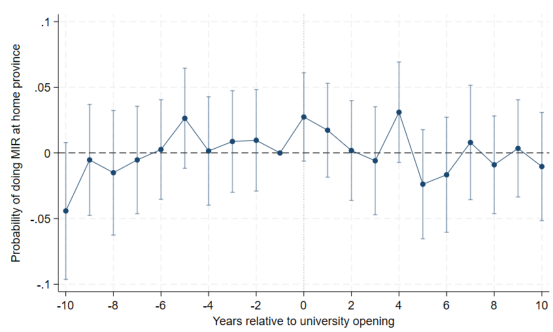
(a) Educational location: first school



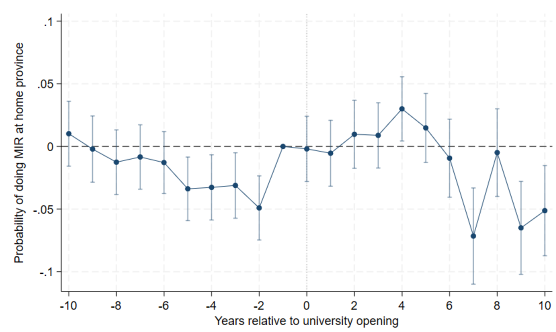
(b) Educational location: additional school



(c) Residency-location: first school



(d) Residency-location: additional school

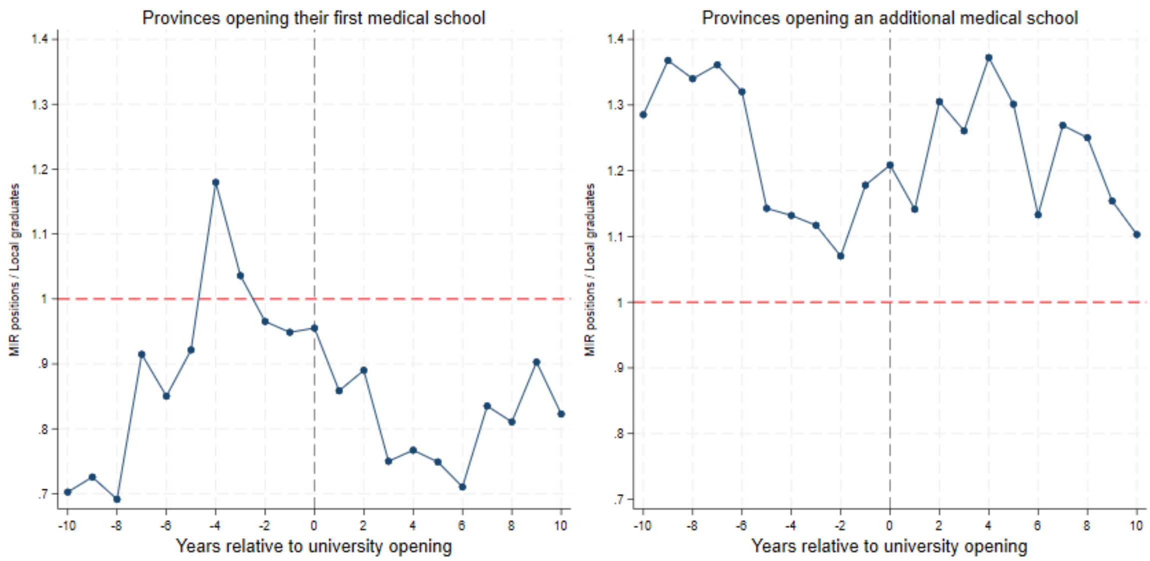


Notes: The figure reports event-study estimates of the effect of a medical school opening on the home bias. The top row shows the impact on the probability of completing the medical degree in the home province (educational location); the bottom row shows the impact on the probability of selecting an intern position in the home province (residency-location). The left column corresponds to provinces opening their first medical school (*no-school* vs. *first-school*); the right column to provinces opening an additional one (*existing-school* vs. *additional-school*). Event time is defined relative to the first year in which a MIR candidate who graduated from the newly opened medical school in a province appears in the data. Points denote coefficient estimates and vertical bars represent 95% confidence intervals. Standard errors are robust.

To further investigate the residency-location pattern in *additional-school* provinces, we split the sample between candidates who studied locally—the treated group in the IV framework—and those who did not. The results are reported in Appendix Figure A.1. Among candidates who did *not* study locally, the coefficients are scattered around zero throughout the pre- and post-event windows, with no systematic pattern. This finding is consistent with a placebo test: opening an additional medical school has no discernible effect on local *MIR* retention among candidates who did not study locally. Among treated candidates, however, the pattern is more nuanced. The pre-event coefficients exhibit a declining trend beginning around $t-T_r=-5$, warranting caution when interpreting the post-event estimates as purely causal. During the early post-event period, the coefficients are modestly positive, peaking at approximately 3–4 percentage points around $t-T_r=4$. From $t-T_r=6$ onward, however, the estimates turn sharply negative, reaching approximately -10 percentage points by $t-T_r=7$ and remaining negative through the end of the event window. Although the pre-event trend precludes a clean causal interpretation of the short-run increase, the pronounced long-run reversal is consistent with labor-market saturation operating through the educational channel: as the supply of locally trained graduates expands without a commensurate increase in local *MIR* positions, even candidates who studied locally are increasingly displaced to other provinces. We examine this potential constraint on local assimilation capacity—measured by the ratio of *MIR* positions to medical graduates—in the next section.

The long-run reversal points to a labor-market saturation mechanism: as the supply of locally trained graduates expands without a corresponding increase in local *MIR* positions, candidates may increasingly be compelled to take residency positions in other provinces. Figure 5 provides direct evidence on this mechanism by plotting, for each group of provinces, the ratio of locally available *MIR* positions to local medical graduates around the event. In *first-school* provinces, this ratio fluctuates around one before the event but falls sharply to approximately 0.75–0.80 from $t-T_r=0$ onward and does not recover. This pattern indicates that the new medical school produces more local graduates than the provincial *MIR* market can absorb, helping explain the null estimates in panel (c) of Figure 4: the educational channel is activated but constrained at the point of labor-market entry. In *additional-school* provinces, by contrast, the ratio remains above 1.1 throughout the event window, indicating structural excess capacity in local residency positions. This capacity helps explain both the high pre-existing level of home bias and the absence of an additional effect at $t-T_r=0$ in panel (d) of Figure 4. Importantly, however, the assimilation ratio begins to deteriorate around $t-T_r=4$ and reaches its minimum later in the post-event period. This timing coincides with the turning point in the residency-location coefficients in panel (d), where the initially positive estimates become negative at approximately the same horizon. The correspondence is consistent with a saturation mechanism: as the additional medical school expands the supply of locally trained graduates without a commensurate increase in local *MIR* positions, candidates are progressively displaced to other provinces. Whether expanded local medical education translates into greater labor-market retention therefore depends not merely on opening a medical school, but also on whether the local *MIR* market has sufficient capacity to absorb the additional locally trained graduates.

Figure 5. Assimilation Capacity around Medical School Opening



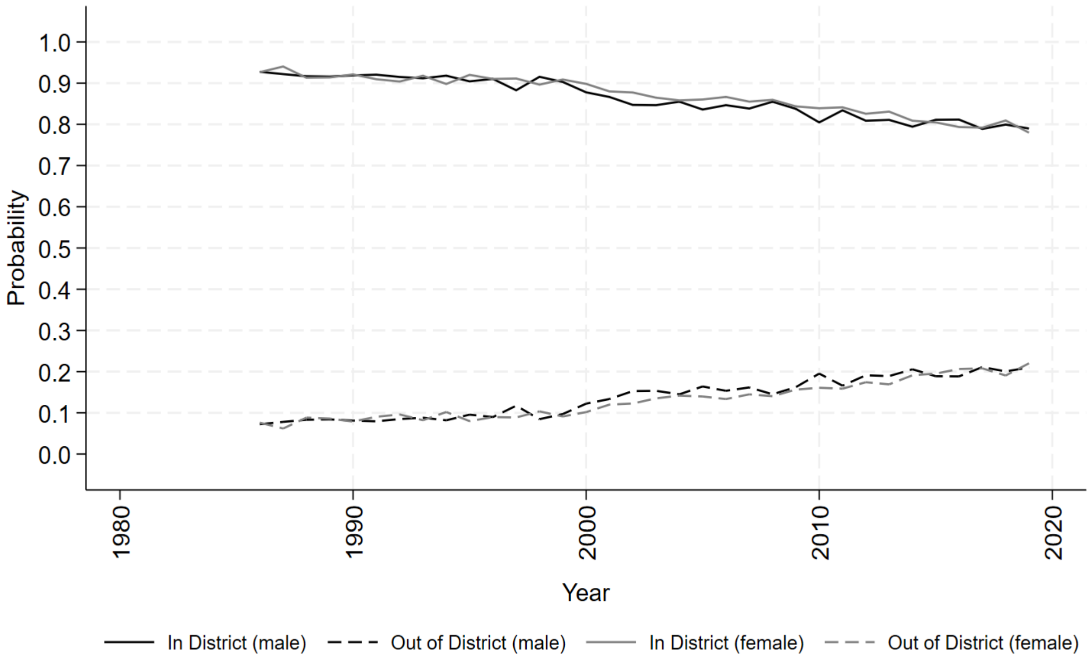
Notes: The figure plots the ratio of local *MIR* residency positions to local medical graduates around the time of a medical school opening, for provinces opening their first medical school (left panel) and provinces opening an additional one (right panel). A ratio below 1 indicates that local residency positions are insufficient to retain all locally trained graduates.

These findings help interpret the event-study results and complement, rather than undermine, the IV estimates. The event study further reveals that whether local medical education translates into greater residency retention depends critically on a province's assimilation capacity. Where sufficient *MIR* positions are available, the educational channel increases local retention; where capacity is insufficient, this channel is constrained at the point of labor-market entry. Accordingly, the IV estimate may represent a lower bound on the underlying preference-based anchoring effect: in capacity-constrained provinces, some graduates who would prefer to remain in their home province may nevertheless be compelled to take residency positions elsewhere.

3.4 The Effect of the *unique district* Reform

In addition to the opening of new medical schools, which anchored students locally, Spain also introduced a major reform that worked in the opposite direction. The *unique district* reform, implemented gradually between 2001 and 2003, liberalized access to public university places by removing the district-specific admission quotas that had previously limited

Figure 6. Evolution of the Proportion of Candidates who Studied Within and Outside the Candidate's District by Gender



Notes: The figure plots the yearly proportion of candidates who studied medicine either within or outside their province-level district, separately by gender. Studying within district refers to enrollment in a province belonging to the same district as the candidate's province of origin, while out of district refers to enrollment in a province belonging to a different district. The full mapping of provinces to districts is reported in Appendix Table A.1.

students' choices largely to their home province or nearby areas.

To analyze students' geographic choices in a consistent way before and after the reform, we rely on the official administrative grouping of provinces into twelve broader districts used in the university admissions system. These districts reflect clusters of geographically proximate provinces with strong socio-administrative ties and historically integrated higher-education markets. The classification allows us to distinguish between studying within the local educational area — where mobility costs and institutional frictions are relatively low — and studying outside that area, which more clearly captures cross-province-level mobility. Table A.1 in the Appendix reports the full mapping of provinces to districts. For instance, District 1 groups the provinces of Catalonia and the Balearic Islands (Barcelona, Girona, Lleida, Tarragona, and Baleares), while District 4 includes Madrid and several surrounding provinces from Castilla-La Mancha and Castilla y León.

Figure 6 shows the long-run evolution of study location by gender. Throughout the 1980-

s and 1990-s, almost all students -men and women alike- studied within their province-level district, with probabilities close to 90%. Beginning in the early 2000-s, coinciding with the introduction of *unique district*, this pattern slowly changes: the probability of studying outside the province-level district increases from below 10% in the 1990-s to nearly 20% by the late 2010-s.

Because medical studies last up to six years and candidates usually spend around one year preparing for the *MIR* test, the *unique district* reform (2001–2003) affected candidates who took the *MIR* test only in 2007-2009 onward. We therefore focus on time windows from 2007–2010 and 2008–2011, which capture cohorts entering university within a narrow bandwidth around the reform — specifically, students who enrolled up to two years before and up to two years after its implementation, respectively. This specification estimates the change in the probability of (i) studying within the home district and (ii) completing an intern position in the home province, comparing post-reform cohorts to pre-reform ones.

The results in Table 6 show that the *unique district* reform is associated with decline in the probability that medical students studied within their home district, on the order of 1.9-3.1 percentage points. The effect is statistically significant across all specifications and time windows, indicating that the liberalization of admissions measurably increased students' propensity to study outside their province-level district. Panel B reveals that this shift carries through to the labor-market stage. Home bias falls by 4.5 to 5.2 percentage points among post-reform cohorts, and the estimates are significant at the 1% level throughout. The reform thus operates along the full pipeline: by weakening where candidates train, it also weakens their tendency to return home for *MIR* training. In both panels the interaction term with female is small and statistically insignificant, indicating that male and female candidates responded to the reform in comparable ways.

4 Conclusion

This paper studies home bias among medical doctors in their entry into the labor market over a 30 year period of time. Using administrative data on the universe of Spanish medical graduates between 1986 and 2019, we exploit two major institutional changes that operated in opposite directions: the staggered opening of new medical schools, which expanded local educational supply, and the nationwide liberalization of university admissions through the unique district reform, which broadened access to non-local institutions. Together, these institutional changes provide a unique opportunity to examine how the geography of educa-

Table 6. Effect of the *Unique District* Reform on Studying Locally and Home Bias

Panel A. Probability of Studying within District

	2007–2010 (1)	2008–2010 (2)	2007–2010 (3)	2008–2010 (4)
Post	-0.0253***	-0.0197***	-0.0303***	-0.0310***
	(0.0070)	(0.0068)	(0.0106)	(0.0104)
Female	0.0124**	0.0117**	0.0089	0.0037
	(0.0056)	(0.0056)	(0.0075)	(0.0076)
Post × Female			0.0071	0.0156
			(0.0108)	(0.0109)
Rank position	-0.1457***	-0.1328***	-0.1457***	-0.1328***
	(0.0205)	(0.0212)	(0.0205)	(0.0212)
Constant	0.9583***	0.9709***	0.9608***	0.9764***
	(0.0391)	(0.0376)	(0.0393)	(0.0377)
Observations	17,635	17,632	17,635	17,632
R-squared	0.2334	0.2309	0.2334	0.2310

Panel B. Home Bias

	2007–2010 (5)	2008–2011 (6)	2007–2010 (7)	2008–2011 (8)
Post	-0.0521***	-0.0489***	-0.0476***	-0.0450***
	(0.0090)	(0.0091)	(0.0136)	(0.0138)
Female	0.0091	0.0171**	0.0123	0.0198*
	(0.0074)	(0.0074)	(0.0101)	(0.0104)
Post × Female			-0.0064	-0.0054
			(0.0143)	(0.0144)
Rank position	-0.4990***	-0.5459***	-0.4989***	-0.5459***
	(0.0266)	(0.0278)	(0.0266)	(0.0278)
Constant	0.9165***	0.9219***	0.9143***	0.9200***
	(0.0508)	(0.0511)	(0.0510)	(0.0513)
Observations	17,635	17,632	17,635	17,632
R-squared	0.2418	0.2460	0.2418	0.2460

Notes: Panel A estimates the probability of studying within the home district; Panel B, home bias (equal to 1 when the candidate's intern position is in their province of origin). *Post*equals 1 for the two post-reform years in each window; *Female*equals 1 for female candidates; *Post × Female*is their interaction. *Rank Position*is the candidate's standardized exam rank (lower = better). All columns include year and specialty fixed effects, as well as province level fixed effects (home, university and hospital province). Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

tional opportunities influences where highly skilled workers train and begin their professional careers.

Our results point to a clear and robust conclusion: local educational opportunities have a large causal effect on early-career geographic attachment. Access to a medical degree in the home province substantially increases the probability that individuals work in their home provinces. Instrumental-variable estimates indicate that studying medicine in the home province increases by 22 percentage points the likelihood of remaining in the same province to work. At the same time, a falsification exercise suggests a modest deviation from the exclusion restriction, so we interpret effect sizes cautiously. A sensitivity analysis indicates the qualitative conclusion is robust, and the results are stable when controlling for province-level positions, hospitals, and local student cohorts. Event-study analysis further indicates that this anchoring effect depends highly on whether the local labor market can assimilate the newly graduated students. Finally, liberalizing university admissions also affects early-career geographic attachment, and through both stages: study location and work location. The *unique district* reform reduced the probability that students studied within their home district and, unlike a purely educational-supply channel, this shift carried through to the *MIR* stage: post-reform cohorts were significantly less likely to exhibit home bias, with reductions of roughly four to six percentage points. Broadening access to non-local institutions thus loosens geographic attachment not only where candidates train but also where they begin their careers, indicating that even in a low-mobility context such as Spain, expanding students' study location sets can meaningfully weaken home bias.

Our paper speaks to the ongoing debate on place-based policies and the design of education systems in low-mobility labor markets. Although our analysis focuses on the medical profession, the mechanisms we document –early anchoring through local training, network formation, and reduced mobility costs– are likely relevant for other high-skill occupations with structured educational pathways. Understanding these dynamics is essential for designing policies that balance province-level retention with the efficient allocation of talent across space.

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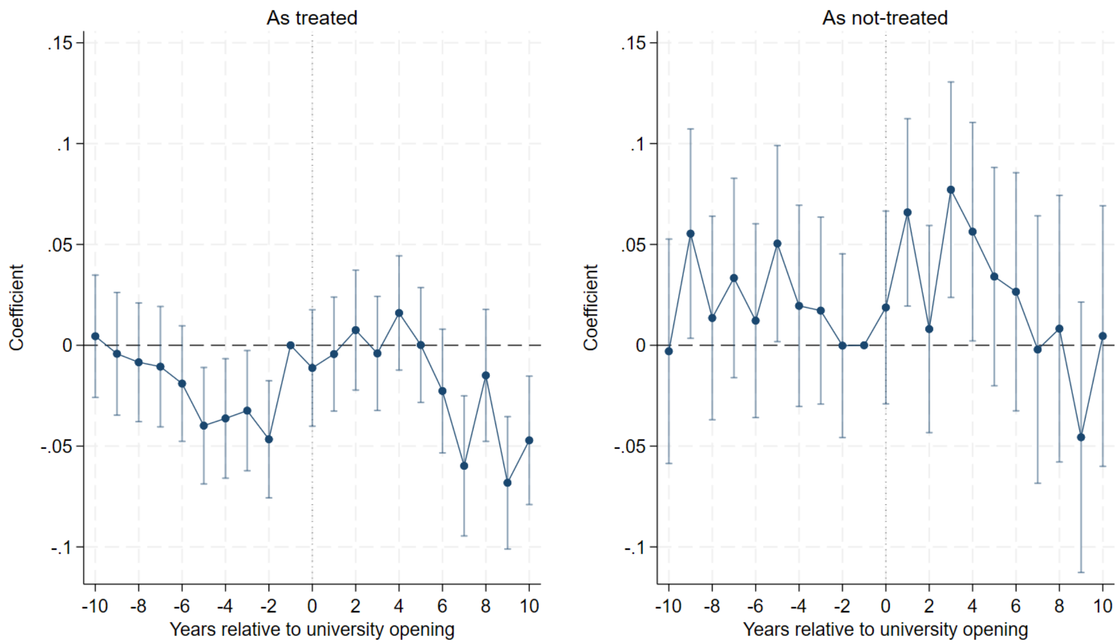
5 Figures and Tables in the Appendix

Figure A.1: Event-Study Estimates of Additional School Openings on Home Bias, by Local-Study Status



*Notes:*The figure reports event-study estimates of the effect of an additional medical school opening on home bias, splitting the sample by whether candidates studied locally. The left panel shows as treated (those who studied locally); the right panel shows those who did not. Vertical bars denote 95% confidence intervals.

Table A.1: Mapping of Provinces to Districts

District	Provinces
1	Lleida, Barcelona, Tarragona, Girona, Baleares
2	Granada, Almería, Jaén, Málaga
3	Las Palmas, Santa Cruz de Tenerife, Ceuta, Melilla
4	Madrid, Segovia, Toledo, Guadalajara, Cuenca, Ciudad Real
5	Murcia, Albacete
6	Asturias, León
7	Ávila, Salamanca, Zamora, Cáceres
8	A Coruña, Ourense, Lugo, Pontevedra
9	Sevilla, Córdoba, Cádiz, Huelva, Badajoz
10	Valencia, Alicante, Castellón
11	Valladolid, Burgos, Palencia, Vizcaya, Alava, Guipúzcoa, Cantabria
12	Zaragoza, Huesca, Teruel, Navarra, La Rioja, Soria

Table A.2: Descriptive Statistics by Year and Gender

	Rank Position		Prob. Studying in Home province		Prob. Doing *MIR* in Home Province (Home bias)	
Year	M	F	M	F	M	F
1986	926	1,006	0.746	0.784	0.585	0.654
1987	1,158	1.220	0.751	0.757	0.589	0.638
1988	1,963	1,683	0.741	0.744	0.585	0.640
1989	2,045	2,205	0.740	0.748	0.539	0.568
1990	2,109	2,321	0.736	0.747	0.515	0.572
1991	1,950	2,123	0.737	0.729	0.527	0.583
1992	2,163	2,359	0.733	0.715	0.542	0.603
1993	2,070	2,289	0.723	0.740	0.591	0.613
1994	2,227	2,475	0.720	0.726	0.572	0.633
1995	1,915	2,150	0.706	0.748	0.560	0.553
1996	1,734	1,810	0.730	0.751	0.582	0.619
1997	1,564	1,661	0.717	0.739	0.550	0.620
1998	1,593	1,729	0.739	0.728	0.560	0.612
1999	1,644	1,754	0.735	0.742	0.560	0.582
2000	1,934	2,009	0.712	0.728	0.522	0.541
2001	2,649	3,039	0.701	0.710	0.563	0.565
2002	2,737	3,066	0.675	0.696	0.558	0.596
2003	2,849	3,191	0.694	0.695	0.561	0.584
2004	2,654	2,928	0.684	0.682	0.559	0.614
2005	2,804	2,989	0.665	0.684	0.601	0.644
2006	2,825	3,037	0.663	0.682	0.584	0.632
2007	3,084	3,196	0.649	0.680	0.608	0.645
2008	3,110	3,219	0.678	0.676	0.555	0.615
2009	3,391	3,349	0.660	0.674	0.554	0.603
2010	3,566	3,709	0.640	0.649	0.572	0.608
2011	3,629	3,674	0.663	0.649	0.540	0.582
2012	3,112	3,382	0.647	0.639	0.539	0.555
2013	2,599	3,097	0.645	0.644	0.517	0.529
2014	2,887	3,324	0.644	0.617	0.462	0.497
2015	2,867	3,342	0.629	0.628	0.466	0.494
2016	3,027	3,550	0.636	0.620	0.470	0.486
2017	3,072	3,604	0.616	0.617	0.470	0.467
2018	3,194	3,782	0.623	0.621	0.458	0.479
2019	3,731	4,225	0.628	0.615	0.455	0.477
Mean values	3,410	2,911	0.685	0.687	0.546	0.583

Notes: This table reports descriptive statistics by year and gender. *Rank Position* corresponds to the position obtained in the national MIR ranking (lower values indicate better performance). *Prob. Studying in Home Province* denotes the share of candidates who completed their medical degree in their province of origin. *Prob. Doing MIR in Home Province* measures home bias and refers to the share of candidates who completed their residency in their province of origin. M and F denote male and female candidates, respectively.

Table A.3: Robustness Checks

	Home bias (1)	Home bias (2)	Home bias (3)
Had a Local University and Stayed		0.1820^{***}	
		(0.0512)	
Female	0.0023	0.0082^{***}	0.0032
	(0.0039)	(0.0024)	(0.0039)
Rank Position	-0.1710^{***}	-0.1680^{***}	-0.1680^{***}
	(0.0098)	(0.0066)	(0.0100)
Local assimilation ratio		0.1670^{***}	0.1130^{***}
		(0.0067)	(0.0081)
No. Hospitals		0.0003^{***}	0.0001
		(0.0001)	(0.0001)
No. Med. Specialities		0.0013^{***}	-0.0058^{***}
		(0.0004)	(0.0006)
Prov. M. University	0.0524^{***}		0.0445^{***}
	(0.0147)		(0.0150)
Constant	0.289^{***}	0.320^{***}	0.373^{***}
	(0.0401)	(0.0518)	(0.0451)
Observations	47,465	149,642	47,465
R-squared	0.354	0.268	0.354
Year and Specialty FE	Yes	Yes	Yes
Province-level FE	Yes	Yes	Yes

Notes: Column 1 presents a falsification test of our IV identification strategy. Column 2 reports the baseline IV estimates with additional controls, and column 3 replicates the falsification test under this extended specification. In all columns, the dependent variable is our measure of home bias, defined as an indicator equal to one if the candidate chooses a medical residency position in the same province as their province of origin. Province-level fixed effects include home, university, and hospital province fixed effects. Robust standard errors in parentheses.