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How does immigration affect housing costs in Switzerland?

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Abstract

This paper examines the short-run immigration effects on prices for owner-occupied housing and rents in Switzerland, exploiting regional variation at the level of 106 local labor markets ("Mobilité Spatiale" regions) and 26 cantons, respectively. We propose two empirical strategies that exploit the Agreement on the Free Movement of Persons (AFMP) with the European Union (EU), enacted in 2002, as an exogenous shock to immigration. The first approach uses the AFMP reform within an instrumental variable approach, instrumenting current regional inflows of immigrants based on the historical distribution of immigrants across regions. The second conducts an event study of housing price changes before and after the reform, distinguishing between regions with historically high, medium, and low immigration from EU-15 countries. The analysis based on data at the level of local labor markets for the years 1985–2016 suggests that immigration triggered off by the AFMP reform has substantially raised prices of single-family homes and of owner-occupied apartments. Before the reform, immigration has not affected house prices. Estimates based on cantonal data for the years 1998–2016 suggest that immigration has raised rental prices even more than prices of owner-occupied housing.

Keywords Agreement on the Free Movement of Persons, Immigration, Shift-share instrument, Event study, House prices, Rental rates

JEL Classification F22, O18, R31

1 Introduction

Housing costs are the largest component of household spending and housing wealth is the largest private wealth component (e.g., Piketty & Zucman, 2014; Jordà et al., 2016). Most advanced countries experienced a strong upward trend in housing costs since the mid 20th century (Knoll et al., 2017). In Switzerland, according to Fig. 1, prices of both single-family homes and owner-occupied apartments have roughly doubled in the period 1985–2016. The price increases were most pronounced from the second half of the 1980 s until the early 1990 s and from the early 2000 s onwards. Rental prices (for new lettings) show a clear upward trend from 1999 onwards. Surging house prices and their associated increases in rental prices have first-order distributional consequences, as the expenditure share for housing is sharply decreasing in income and wealth (Dustmann et al., 2018). Consequently, rising housing costs imply that disposable income net of housing costs decreases relatively more for low-income households than for high-income

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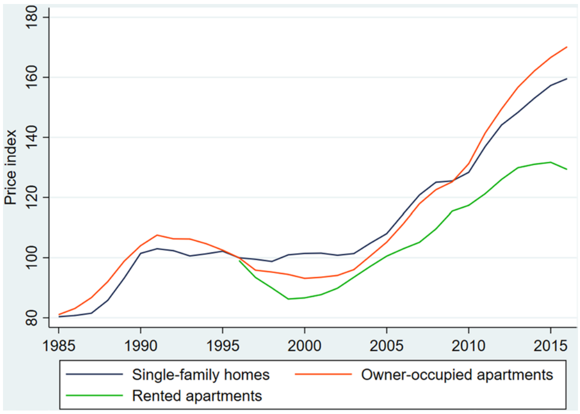

Fig. 1 Price indices for single-family homes, owner-occupied apartments and rented apartments in Switzerland, 1985–2016. *Notes:* Base year = 1996. Aggregate data for entire Switzerland. *Source:* Own calculations based on data from Wüest Partner

households, raising welfare inequality (Grossmann et al., 2021a).[1]

One potential cause of increasing housing prices is rising demand for housing associated with population growth. Significant increases in population size are typically associated with immigration waves. For instance, Grossmann et al. (2017) develop an overlapping generations model with a housing sector and show that an endogenous immigration wave in response to international labor market integration associated with high productivity in the destination leads to an increase in rental prices for housing.[2]

However, there may be offsetting factors such as negative income effects or native out-migration at the local level in response to immigration.[3] Thus, *a priori*, the sign and magnitude of the average local impact of immigration on housing prices are not obvious.

This paper examines the extent to which immigration affects average housing prices across Swiss regions in the short-run (where supply can be viewed as approximately fixed), focusing on prices for single-family homes, owner-occupied apartments, and rental prices for dwellings. Specifically, we use regional data on the prices of owner-occupied homes for the period 1985–2016 at the level of 106 local labor markets, i.e., MS-regions[4], as well as on rental rates for the period 1998–2016 at the cantonal level to relate the annual growth rate of these housing price indicators to the annual change in the stock of foreigners relative to the initial total population.

We propose two empirical strategies to recover causal effects of immigration on housing prices. In both we exploit the Agreement on the Free Movement of Persons (AFMP) with the European Union (EU) that fully removed immigration restrictions for EU workers from 2002 onwards as an exogenous shock to the inflow of immigrants. The first is an instrumental variables (IV) approach that employs the widely used "shift-share" instrument for immigration. This instrument uses the historical distribution of immigrants (as of year 1980) across regions in Switzerland to predict current inflows into the respective regions. It exploits the tendency of newly arriving immigrants to move to areas where other immigrants of the same nationality already live (Bartel, 1989). The exclusion assumption justifying this instrument is that historical settlement patterns have no direct effect on the growth of current housing prices. Combining the IV approach with the exogenous increase in immigration in response to the reform allows us to separately study the effects of immigrant inflows before and after the reform. In some specifications, we therefore interact the immigrant inflow with a post-AFMP-reform dummy and instrument both the main variable and the interaction term. The differential effect of immigration in the post- relative to the pre-reform period can be interpreted as a difference-in-differences (DiD) estimate of the effect of the AFMP reform on house prices. This DiD effect should be robust to a violation of the exogeneity condition of the instrument stemming from heterogeneous regional trends that jointly affect the historical settlement pattern and housing price dynamics.

Our second empirical approach consists of an event study of the changes in house prices before and after the AFMP reform, where we group regions according to their historical share of immigrants from EU-15 countries. We distinguish between a low, medium, and high share of immigrants from EU-15 countries in 1980.[5] Exploiting

[1] Increases in housing prices also affect the wealth distribution. Evidence by Kuhn et al. (2020) for the U.S. (where house ownership rates are comparably high) suggests that wealth inequality tends to decline, as housing wealth is more equally distributed than non-residential wealth. Thus, wealth inequality and welfare disparities may move in opposite directions (Grossmann et al., 2021a)

[2] Notably, this holds even in the long-run after full adjustment of housing production. The result is rooted in the scarcity of land (or land-use regulations), implying a strong connection between land prices and housing prices. The importance of land prices for the evolution of house prices is emphasized in Knoll et al. (2017) and at the center of the dynamic macroeconomic model of Grossmann et al. (2021b). Glaeser et al. (2008) argue that in areas where housing supply is more elastic, higher housing demand leads to smaller increases in house prices and rents associated with fewer and shorter bubbles.

[3] See Sa (2015) for a theoretical analysis on the counteracting effects of immigration on housing prices at the local level.

[4] MS stands for "Mobilité Spatiale", i.e., spatial mobility. MS-regions are defined by the Federal Statistical Office in Switzerland. They are characterized by spatial homogeneity and a common local labor market.

[5] The approach is inspired by Beerli et al. (2021), who consider effects of the increase in the availability of cross-border workers in response to the AFMP reform on labor market outcomes, firm productivity, innovation activity, and establishment entry and exit. Beerli et al. (2021) also consider a transition phase to account for possible anticipation effects but found none.

the exogeneity of the AFMP reform, the event study approach allows us to verify that housing price dynamics are indeed unrelated to the historical share of immigrants before the reform, which supports the validity of our shift-share instrument.

We probe the robustness of our main results in further sensitivity analyses. In particular, we account for the so-called second home initiative, a referendum approved in 2012. The initiative led to a reform that restricts the construction of homes by owners who do not have first residency in those municipalities where the fraction of second homes exceeds 20% (mostly in touristic areas like those close to ski resorts), with potential effects on housing prices (e.g., Hilber & Schöni, 2020).

Our main findings are as follows. Exploiting MS-level variation for the 1985–2016 period, the IV estimates suggest a significant and positive (short-run) impact of immigration on house prices after the AFMP reform, but not in the pre-reform period. We find that an annual increase in the stock of foreigners equal to 1% of the initial population leads to an increase in single-family home prices by 4.3% and in owner-occupied apartment prices by 5.9% after the reform. Based on cantonal data, the same increase in immigration raises rents by 7.4% for the period 1998–2016 and by 8% for the period 2002–2016. Our second, event study approach suggests that switching from a region with a historically low or medium level of immigration from EU-15 countries to one with a high past stock of EU immigrants raises the annual growth rate of house prices by about one percentage point after the AFMP reform.

While we follow most of the previous literature to approximate the immigration inflow by the change in the number of foreign nationals in our main analysis (Saiz, 2007; Gonzalez & Ortega, 2013; Sa, 2015; Degen & Fischer, 2017), alternatively, we approximate the immigration inflow as the number of foreigners entering the country minus those leaving it (net migration). We find that the estimated housing cost coefficients after the reform are then considerably lower (but still highly significant). The effect of an annual increase in net migration relative to initial population by 1% now raises rents by 2.2% (rather than 7.4%) for the 1998–2016 period (cantonal variation). For the longer period (exploiting MS-regional variation), the same immigration push raises single-family home prices by 1.8% (rather than 4.3%) and owner-occupied apartment prices by 2.4% (rather than 5.9%) after the AFMP reform.

There is a growing empirical literature on the causal effects of immigration on the housing market. Yet, none of it has used an immigration reform for identification. Other studies focussing on Swiss house prices are Degen and Fischer (2017) and Fischer (2012). Degen and Fischer (2017) employ data for 85 regions in Switzerland from 2001 to 2006. Including regional fixed effects (like we do) and employing a "shift-share" instrument based on the distribution of immigrants in 1997, they find that an immigrant inflow of 1% of a MS-region's population (excluding MS-regions with less than 25'000 inhabitants) is associated with a 2.6% increase in single-family home prices, 2.8% for multi-family home prices, and 0.7% for prices of condominiums.[6] Fischer (2012) considers the same time period and argues that the price effects are driven by immigrants who do not share a common language with the destination.

The most important difference to these contributions is that our analysis (covering a considerably longer period of 31 years) evaluates, for the first time, the effect of the AFMP reform on housing costs in Switzerland and uses the reform for identification. Moreover, the added event study analysis—based on historical stocks of immigrants from the EU-15 countries that were affected by the AFMP reform—addresses an important concern about the standard IV approach of using a "shift-share" instrument in the literature on immigration effects.[7] That is, regions may have been on different economic paths that are related to both housing price dynamics and the historical settlement pattern of immigrants that build the basis for the "shift-share" instrument. As indicated, neither our IV analysis nor the event study suggest that immigration had an impact on housing prices before the AFMP reform (but substantial effects thereafter). This strengthens the confidence in the exclusion restriction of the "shift-share" instrument.

The second contribution is that we also examine the effects of immigration on rents, in addition to prices of owner-occupied housing. From a welfare perspective, examining the effects on rents is particularly important in the Swiss context where the home ownership rate is comparatively low. The effect of immigration on rents may be different from the one on prices of owner-occupied housing. On the one hand, the effect on rents may be lower because of rent regulations. On the other hand, it may be higher because immigrants are more likely to rent than to buy homes.

Third, and relevant in the Swiss context for future research, we also gauge the impact of different data sources for house prices in further analysis. There are two

[6] Häcki (2015) examines the impact of immigration on the prices of single-family homes and owner-occupied apartments for a subsequent period (2007 to 2013) and finds positive effects as well.

[7] A recent literature prominently discusses the necessary assumptions and possible research designs for validity of the shift-share IV approach; see Adãao et al. (2019), Borusyak et al. (2022), and Goldsmith-Pinkham et al. (2020).

important data sources. Whereas our data comes from the independent Swiss consulting company *Wüest Partner*, Degen and Fischer (2017) use house price data from the *Informations- und Ausbildungszentrum für Immobilien* (IAZI). Looking at the same short time period around the AFMP reform (2001–2006) as considered in Degen and Fischer (2017) and following exactly their IV approach, we find considerably larger immigration effects on house prices in the IV estimations.

The majority of studies for other countries conclude that immigration has a positive impact on housing prices. By analyzing metropolitan areas in the US between 1984 and 1998, Saiz (2007) finds that an immigration inflow equal to 1% of a city's initial population leads to an increase in average house prices of about 3%.[8] The main difference to the Swiss context is that the AFMP reform that we use for identification has led to an inflow of skilled immigrants from advanced countries (mostly the EU-15) with relatively high earnings, on average. For Spain, Gonzalez and Ortega (2013) identify the causal effects of immigration by using data at the province level for the period 2000–2010. Their IV estimations suggest that a migration-driven increase in population of 1% leads to a rise in house prices of 1%. According to Sa (2015), by contrast, evidence for across 170 local authorities in the UK between 2003 and 2010 suggest that an immigration inflow equal to 1% of the initial population reduces house prices by 1.7%. A possible explanation is that her results capture offsetting factors such as (native) out-migration at the local level as a response to foreign immigration. Similarly, Saiz and Wachter (2011) look at neighborhoods within metropolitan areas in the U.S. from 1980 to 2000 to analyze the impact of immigration inflows on house prices and find small negative effects. Akbari and Aydede (2012) analyze the house prices of privately owned dwellings in Canada with census-data from 1996, 2001, and 2006 and find only a small (positive) effect of immigration. Overall, the evidence suggests that housing price effects of immigration tend to be positive when using data on larger geographic units and may be negative otherwise.

There exist considerably fewer studies on the immigration effects on rents and none of them has been conducted for Switzerland. Saiz (2003) exploits the inflow of about 80'000 Cuban refugees to Miami in 1980 that led to a rise in Miami's tenant population by 9%. His evidence suggests that it implied rents of lower quality units to increase by 8–11%. Saiz (2007) finds that an increase in the immigrant population share of one percentage point has increased average rents in the U.S. by about 1%, which is considerably smaller than the 3% increase found for house prices. Using census data from 1970 to 2000 on rental prices across US states and their metropolitan residents, Ottaviano and Peri (2007) obtain an even smaller effect.

We interpret our results as short-run demand effects, which hinges on the point that housing supply is fixed in the short-run.[9] To back this assumption, in the Appendix we present short-run (IV) estimates for the impact of immigration on housing supply at the level of MS-regions for the period 2009–2016. The IV estimates indeed suggest non-positive short-run effects on housing supply. Studies on immigration effects on housing supply are rare. Based on first differences IV estimates at the annual level, Gonzalez and Ortega (2013) find that a migration-driven increase in population of 1% leads to a rise in housing units of about 1.1% in Spain. By contrast, Sa (2015) presents evidence suggesting a non-positive (and small) effect for the UK, like we do for Switzerland.[10]

The rest of the paper is organized as follows. Section 2 presents characteristics of the Swiss housing market and institutional reforms that may affect it. Section 3 describes the data sources. Section 4 lays out the empirical methodology. Section 5 provides descriptive statistics. The main results and sensitivity analysis on the effects of immigration on the housing costs are reported in Section 6 and 7, respectively. The last section concludes.

[8] His evidence also suggests that the effects of immigration on house prices may be stronger than of overall population growth. One reason could be that immigrants regionally cluster (Card, 2007).

[9] Moreover, we follow the previous literature in implicitly making the 'stable unit treatment value assumption' (SUTVA) that an increase in housing costs in one region in response to immigration does not affect housing costs in other regions. As discussed, this may not always hold.

[10] In an important study, Büchler et al. (2021) estimate longer-run housing supply elasticities in response to increases in rental income and house prices, finding that the former are higher than the latter. Their research also points to an important role of land-use restrictions and geographical factors. Finally, there is a growing literature that examines other aspects of the Swiss housing market. Basten and Koch (2015) identify the effects of house prices on the mortgage demand and supply in Switzerland, using the exogenous variation of immigration to instrument house prices. They find a positive effect of house prices on the level of mortgages. Also Brown and Guin (2015) analyses the relationship between the Swiss housing market and the mortgage market. Fischer and Zachmann (2020) study the difference between the influence of self-financed property buyers, such as insurances and pension funds, and the influence of bank-financed property buyers, such as homeowners, on local house prices in Switzerland between 2008 and 2015. They find that self-financed property buyers have a strong effect on local house prices. Dambon et al. (2022) analyze spatially varying vintage effects for single-family houses in the Canton of Zurich, Switzerland.

2 Institutional background

2.1 Features of the Swiss housing market

In Switzerland, owner-occupancy is less common than in other countries, particularly in urban areas. For instance, at the turn of the millennium, the home ownership rate was 68% in the U.S., 85% in Spain, 66% in Canada, and 68% in New Zealand (Degen & Fischer, 2017). Germany has the lowest home ownership rate in the eurozone, with 44% in 2010 (Kaas et al., 2021). In Switzerland, it is even lower, with 35% in 2000 and 38% in 2017 (Federal Statistical Office, 2020a). This makes the consideration of rental prices, in addition to owner-occupied house and apartment prices, particularly important in the Swiss context. Werczberger (1997) argues that rent control regulation in Switzerland is comparatively moderate and cannot explain the differences in the homeownership to other countries.[11] It allows landlords to raise rents in response to inflation, growing maintenance costs, higher property taxes, higher mortgage interest rates, or house price increases. Nevertheless, the evolution of house prices and rents may differ because of rent control legislation.

In 2000, about 54% of the rental stock was owned by private landlords, about 32% by private businesses, and about 13% by the government, cooperatives, or non-profit organizations (Werczberger, 1997). Also noteworthy, the Swiss housing market is characterized by a low vacancy rate. In 1995, the nationwide vacancy rate was 1.4%, below 1% between 2002 and 2014, and 1.7% by 2020. Comparing it across cantons, Geneva (0.5%), Zug (0.7%), Zurich (0.9%), Obwalden (0.9%) and Basel (1%) had the lowest vacancy rates in 2020, while Thurgau (2.5%), Jura (2.5%), Aargau (2.7%), Ticino (2.7%) and Solothurn (3.2%) had the highest ones (Federal Statistical Office, 2020b). Moreover, occupancy turnover rates are low in Switzerland. The average length of stay is 5–6 years for rental units, 12–14 years for owner-occupied apartments, and 20 years for single-family homes in Switzerland (Degen & Fischer, 2017).

2.2 The Agreement on the Free Movement of Persons (AFMP)

As outlined in the introduction, we use the AFMP between Switzerland and the EU that came into force in June 2002 for identification of immigration effects on house prices. The AFMP was signed in June 1999 along with six other bilateral agreements (on trade, transportation, and scientific collaboration). The Swiss electorate approved the bilateral agreements in a national referendum in May 2000 with an approval rate of 67.2%.[12] The AFMP stipulates that nationals of the old EU Member States (EU-15) are free to move to Switzerland from 2002 onwards, provided they have a job.[13] In 2006, the agreement was extended to the ten countries that joined the EU in 2004 (EU-10).[14] Since June 2009, the agreement covers also Bulgaria and Romania. Lastly, Croatia joined the agreement in January 2017 (Directorate for European Affairs, 2020).

Evidence on aggregate immigration suggests that the AFMP was an important determinant of immigration dynamics from the EU. The number of EU-15 citizens living in Switzerland increased, on average, by only 0.5% per year in the period 1985–2001, compared to an annual increase of the total foreign population size by 2.4%. In the period 2002–2016, by contrast, the number of EU-15 citizens grew, on average, by 2.9% per year (to 1.25 million), which was slightly above the annual growth rate of the total foreign population of 2.5% (to 2.07 million). Notably, the number of Germans grew, on average, by 6.3% per year in the period 2002–2016 (it has more than doubled to 303’525 inhabitants), while growing only by 2.0% per year in the period 1985–2001.[15] EU-immigrants tend to settle in larger agglomerations, such as Zurich, Basel, Geneva, Lausanne, Bern, Lucerne, and St. Gallen. They also tend to be well-qualified, mostly young, without children, and do not own residential property (Graf et al., 2010). The changing composition of immigrants residing in Switzerland toward younger high-income earners from the EU has thus likely had substantial effects on housing demand, in particular on the rental markets in larger agglomerations.

As shown in Fig. 1, housing prices increased substantially in the second half of the 1980 s, culminating in a real estate market crisis and recession in the early 1990s (Borowiecki, 2009). The subsequent period up to 2001 is characterized by a modest growth of housing prices, while, since 2002, when the AFMP came into effect, housing prices have grown sharply.

[11] Werczberger (1997) rather points to the taxation of capital gains and of the imputed rent in Switzerland. Also high construction costs as driven by building standards play a role. The possibly most important factor is the scarcity of land associated with the Swiss geography and land-use regulations that drive land prices. See Kaas et al. (2021) for a calibrated model to analyze the determinants of homeownership rates.

[12] See swiss votes (2022a). Approval was necessary after a so-called facultative referendum was initiated, requiring 50’000 valid signatures from opponents of parliamentary decisions in Switzerland within 100 days. The majority of voters suffices, irrespective of the regional distribution of votes.

[13] The EU-15 consists of Austria, Belgium, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, Netherlands, Portugal, Spain, Sweden, and the United Kingdom.

[14] These were Estonia, Latvia, Lithuania, Malta, Poland, Slovakia, Slovenia, Czech Republic, Hungary, and Cyprus.

[15] The figures are based on the same data sources for foreign nationals as employed for our estimates. The data sources are described in Sect. 3.

2.3 The second home initiative

In addition to the AFMP, the second home initiative was a policy initiative that led to another important reform that potentially could have affected housing prices. The reform banned the construction of new second homes in municipalities with more than 20% of homes that are not inhabited by those with the first residency in the municipality. The Swiss electorate approved the initiative (against the will of both the national government and the clear majority of the two parliamentary chambers) in March 2012 by a small margin (50.6% of the votes and a small majority of cantons).[16] In January 2013, the government enacted provisional construction bans based on estimates for second homes. Both a Federal Law passed by the national parliament and an updated regulation came into force in January 2016. The implemented law still allows the construction of second homes for business purposes (i.e., renting them out to tourists).

Whereas proponents stressed the goal of protecting the natural landscape, opponents were worried about detrimental economic effects in the affected regions.[17] Hilber and Schöni (2020) show that the construction ban indeed had substantial housing price effects, lowering price growth on average by 15% for primary homes in the treated municipalities and raising it by 26% for secondary homes, according to their preferred specification. Also the unemployment rate increased (on average by 12%) in the affected municipalities.

We account for the second home initiative in a sensitivity analysis by restricting the sample to MS-regions and cantons with less than 20% of second homes in 2012. Accordingly, we exclude 31 out of 106 MS-regions (among them nine in the canton of Grisons and eight in the canton of Valais) and the cantons of Uri, Obwalden, Grisons, Ticino, and Valais.

3 Data

We obtain Swiss housing price data by region and year from the independent Swiss consulting company *Wüest Partner* that focusses on construction and real estate markets. The data distinguishes between prices for single-family homes, for owner-occupied apartments, and for rented apartments. Prices for single-family homes and owner-occupied apartments are provided annually for the period from 1985 to 2016 and for the 106 MS-regions. They are available in the form of a transaction price index with 1985 as the base year. The transaction price index, developed by *Wüest Partner*, is a quality-adjusted price index based on a hedonic valuation model. A hedonic valuation model unbundles an object into separate components, for which people are willing to pay. Relevant factors in the hedonic valuation model by *Wüest Partner* are, for example, living space, condition of the object, location in the municipality, accessibility, and the type of the municipality (Wüest Partner, 2017).

Rental prices for apartments are available by canton and year from 1996 to 2016. *Wüest Partner* offers the rental prices as a quarterly asking price index, i.e., a price index based on advertising prices, with the first quarter of 1996 as the base period. For our analysis we use the second quarter indices in order to analyze rental prices jointly with the annually available variables in our data set. The asking price index is based upon around 500’000 real estate offers per year. These offers include information on prices, living space, condition of the object, and municipality. The price indices are weighted averages of homogenous groups of apartments (Wüest Partner, 2017).

Moreover, we use data from the Federal Statistical Office (FSO) on the population by region and year from the annual population statistics ESPOP which is available from 1981 to 2010 and from the population and household statistics STATPOP which is available since 2011 (Federal Statistical Office, 2017a, d). The main explanatory variable employs the permanent resident population and the permanent foreign resident population of a region for the years from 1985 to 2016.[18] In addition, ESPOP and STATPOP also provide the international net migration of foreigners, which corresponds to immigration minus emigration of foreigners. We use that data as an alternative measure of the immigration flow in our robustness analysis. The construction of the instrument as well as the event study relies, in addition to information on the permanent resident population in Switzerland by nationality from ESPOP and STATPOP, on the 1980 federal population census which offers the resident population by nationality and municipality (Federal Statistical Office, 2017a, d, 1980).[19] Continuing annual information

[16] See swiss votes (2022b). The second home initiative was a so-called popular initiative. Popular initiatives can be initiated by citizens, political parties, or interest groups. They require 100’000 valid signatures in favor of the initiative within 18 months. Both the majority of voters and cantons as defined for the second chamber of the parliament (23 cantons and 6 half-cantons) have to approve such initiative for it to come into effect.

[17] Resistance thus came mainly from the affected areas. For instance, in Valais, a canton with many municipalities exceeding the 20% threshold, the share of yes-votes was the lowest among all cantons (26.2%), whereas in the unaffected canton of Basel-city it was highest (62.2%) (swiss votes, 2022b).

[18] The permanent foreign resident population includes all foreign nationals who resided in Switzerland for a minimum of 12 months or who have a residence permit for a minimum of 12 months.

[19] The number of nationalities is limited to 39. Nationalities with very few people living in Switzerland are summarized on a continental level. Therefore, the dataset lists most of the European countries, some major countries from the other continents, as well as one number for the remaining countries of each continent.

on the permanent resident population by nationality and municipality is available only from 1990 onward.[20] To identify the MS-regions and cantons with more than 20% of second homes in 2012, that we exclude in the sensitivity analysis, we use the data from the Federal Office for Spatial Development ARE (2022).

Additional control variables with data sources are as follows. First, we use the unemployment rate provided by the State Secretariat for Economic Affairs (SECO) which is available from 1985 to 2016 at the cantonal level. The SECO unemployment statistics include all unemployed persons registered with a regional employment center (RAV) at the end of a month (State Secretariat for Economic Affairs, 2017). Second, we employ data on the monthly gross wage by major region as provided by the Swiss earnings structure survey conducted by the FSO.[21] It reports average earnings per worker (not differentiated at the occupational level) as full-time equivalent in Swiss Francs (CHF). The earnings information is available for the years of 1998 to 2016 (Federal Statistical Office, 2017f). Third, we include the construction price index by major region and year from the Swiss construction price statistics provided by the FSO. The index is published half-yearly and is normalized to 100 for the second half of 1998. We use the indices from the second half of each year for the available period from 1998 to 2016. The construction price statistics measure the effective market price development in the construction sector by mainly recording contract prices (Federal Statistical Office, 2017e).[22] Finally, we control for the number of vacant apartments by region and year as measured at the first of June by the empty dwellings census of the FSO. The data is available for the years form 1995 to 2016 (Federal Statistical Office, 2017c).

After combining all variables, the final dataset used for the regression analysis of owner-occupied house prices includes the 106 MS-regions in Switzerland for the period from 1985 to 2016 and the dataset used for the analysis of rental prices covers the 26 cantons for the period from 1998 to 2016. The analysis at the level of MS-regions required us to take into account municipal mergers that are documented in the Swiss official municipality register (Federal Statistical Office, 2017g).[23] Table 1 gives an overview of the variables employed in the main analyses with data sources. Figure 9 in the Appendix shows the 106 MS-regions in Switzerland, whereas Fig. 10 displays the seven major regions and the 26 cantons.

4 Methodology

We propose two different empirical strategies to recover the causal effect of immigration on housing costs. The first one uses the AFMP reform in combination with an IV approach, while the second one employs the reform in the context of an event study of house price changes.

4.1 Model 1: IV approach

Our estimated effects of immigration on prices for housing are based on equation

$$\Delta(\log P_{i,t}) = \beta_1 \cdot \frac{\Delta I_{i,t}}{Pop_{i,t-1}} + \beta_2 \cdot \frac{\Delta I_{i,t}}{Pop_{i,t-1}} \times \mathbf{1}(t \geq 2002) + \beta_3 \cdot \Delta X_{i,t-1} + \alpha_t + \varepsilon_{i,t}, \tag{1}$$

where $P_{i,t}$ is the employed housing price variable (for single-family homes, for owner-occupied apartments, or for rented apartments) in region i and year t. The price indices are quality adjusted, so any price changes do not reflect changes in the quality of houses. Throughout the paper, $\Delta x_{i,t}$ denotes the change in a variable x in region i over time, typically between years $t-1$ and t (in sensitivity analyses sometimes between years $t-3$ and t). Thus, typically, $\Delta(\log P_{i,t})$ is (approximately) the annual percentage change of $P_{i,t}$. $\Delta I_{i,t}/Pop_{i,t-1}$ is the annual change in the stock of foreigners relative to the initial population and $\mathbf{1}(t \geq 2002)$ is a dummy variable equal to one from year 2002 onwards, i.e., since the AFMP is in place. Considering changes over time on both sides of the estimated equation avoids omitted variable bias from time-invariant factors specific to each region like the type of a region (urban versus rural), proximity to a city, etc. $X_{i,t-1}$ denotes a vector of control variables capturing lagged time-varying region-specific characteristics. In the MS-level regressions, $\Delta X_{i,t-1}$ is the (lagged) change in the unemployment rate to control for changes in regional economic conditions that affect housing demand. α_t denotes a set of year dummies capturing national trends,[24] and ε_{it} is the error term.

Because housing rental rates are available for a shorter period (1998–2016) than house prices and at the cantonal (rather than MS) level only, we restrict $\beta_2 = 0$ for all estimates with cantonal level data. In fact, in the period

[20] The respective data sources are the federal population census for 1990, PETRA (statistics of foreign resident population) for 1991–2009 and STATPOP for 2010–2016 (Federal Statistical Office, 1980, 2022, 2017d).

[21] There are seven major regions is Switzerland: Lake Geneva region, Espace Mittelland, Northwestern Switzerland, Zurich, Eastern Switzerland, Central Switzerland, Ticino.

[22] Building construction includes both new construction and renovation of above-ground buildings (single-family houses, apartment buildings, office buildings).

[23] Over the last 30 years the number of municipalities in Switzerland decreased steadily. The final dataset reflects the situation as of April 2017. At that time, there were 2240 municipalities in Switzerland.

[24] One may think about the general inflation rate, the Swiss franc exchange rate, or the long-term interest rate, which are factors affecting building costs and terms of financing.

Table 1 Data description

Variable	Explanation	Unit	Years	Region	Source
Dependent variables					
HP_{it}	Price index for single-family homes	1985 = 100	1985–2016	MS	WP
AP_{it}	Price index for owner-occupied apartments	1985 = 100	1985–2016	MS	WP
RP_{it}	Price index for rented apartments	1996 = 100	1996–2016	Canton	WP
Independent variables					
Pop_{it}	Permanent resident population	Persons	1985–2016	Municip.	FSO
I_{it}	Permanent foreign resident population	Persons	1985–2016	Municip.	FSO
Control variables					
u_{it}	Unemployment rate	[0,1]	1985–2016	Canton	SECO
$wage_{it}$	Monthly gross wage	CHF	1998–2016	Maj. reg.	FSO
CP_{it}	Construction price index	1998 = 100	1998–2016	Maj. reg.	FSO
VAC_{it}	Number of vacant apartments	Apartments	1995–2016	Municip.	FSO

This table summarizes the information on data availability. It mentions in each case the spatially smallest region and the complete time span for which the data are available. WP stands for Wüest Partner, FSO for Federal Statistical Office and SECO for State Secretariat for Economic Affairs

1998–2001, the reform was already on its way, albeit Fig. 1 suggests considerable changes in housing costs only from 2002 onwards. In order to compare the effects on rents with those on house prices, we estimate analogous regressions for house prices at the cantonal level over the shorter time period. At the cantonal level, we are able to consider a detailed set of controls for changing economic conditions ($\Delta X_{i,t-1}$): the change in the log of average gross (full-time equivalent) wage because of its potential income effect, the change in construction prices (index) to control for the changing costs of buildings, and changes in the number of vacant apartments to capture changes in local factors such as a region-specific newly introduced law or a local event that limits housing supply unexpectedly (e.g., a landslide).

With MS-regions as observational units, we cluster standard errors by MS-regions. With cantonal data, however, this could lead to biased standard errors because the number of clusters is too small. We therefore use the wild bootstrap, resampling at the cantonal level (e.g., Cameron et al., 2008, Djogbenou et al., 2019, Roodman et al., 2019, MacKinnon et al., 2022), and report p-values rather than standard errors throughout.

4.1.1 Expected immigration effects

The main coefficients of interest in Model 1 are β_1 and β_2. If correctly identified, β_1 provides the percentage change in housing prices in response to an annual increase in the stock of foreigners equal to 1% of the initial population before the AFMP reform and $\beta_1 + \beta_2$ the corresponding effect after the reform. Since the data has an annual frequency and the change in the foreign population is not lagged, our estimates can be interpreted as a short-run demand effect of immigration because possible housing supply adjustments are small (we test this hypothesis in the Appendix), if there are severe supply restrictions (land-use regulation or land scarcity) or if price effects are largely unanticipated. As discussed in the introduction, standard theory suggests that an inflow of immigrants raises housing demand, thus leading to an increase in house prices and rents. However, the magnitude and also the sign of the effects depend on factors such as income or out-migration of locals. Because the considered regions are spatially relatively large by Swiss standards and the immigration effects are likely to be higher when the size of the inflow of immigrants is larger (hypothesizing a non-linear impact of immigration), we expect $\beta_2 > 0$ and $\beta_2 > \beta_1$.

4.1.2 Potential endogeneity and the "shift-share" instrument

The estimates of the effects of immigration on the housing market potentially suffer from different endogeneity issues. First, omitted factors could lead to a correlation of immigration with the error term. This bias is likely to be mitigated by first differencing and the inclusion of time fixed effects, but potentially not fully. Second, immigrants are not randomly distributed across regions, but rather choose themselves where to settle. This raises reverse causality concerns, albeit the sign of the bias is unclear. On the one hand, immigrants may prefer to live in attractive regions that face increasing demand for housing also from internal migration, such as urban areas, where housing prices are growing fast.

This would lead to an upward bias in OLS estimates of β_1 and β_2. On the other hand, however, controlling for a region's economic condition, immigrants may also prefer to locate in areas where housing prices increase more slowly. In that case, there would be a downward bias in OLS estimates of β_1 and β_2.

To address such potential endogeneity, we employ the shift-share instrument for immigration, first proposed in the context of immigration by Card (2001), which is now widely used in the literature on immigration.[25] It makes use of geographical variation in the historical inflow of foreigners, assuming that it is exogenous to recent developments in the housing market. Specifically, we predict the stock of foreign-born individuals in each region i at time t using the historical settlement patterns of immigrants as of 1980 by country of origin. According to Bartel (1989), migrant networks are an important driver of location choices of newly arriving immigrants. Immigrants tend to move to areas in which other immigrants of the same nationality reside already over-proportionally (e.g., Germans in Zurich, Portuguese in Fribourg, etc.), because the possibility to rely on a social network reduces migration costs. Thus, the instrument captures the supply-push component of recent immigrant inflows.[26]

The instrument for $\Delta I_{i,t}/Pop_{i,t-1}$, the annual change in the stock of foreigners in region i and time t relative to the region's initial total population, is constructed as

$$z_1 \equiv \frac{\Delta \hat{I}_{i,t}}{Pop_{i,t-1}} \text{ with } \hat{I}_{i,t} = \sum_c \left(\frac{I_{c,i,t_0}}{I_{c,t_0}}\right) I_{c,t}, \tag{2}$$

where $I_{c,t}$ is the total stock of foreigners from origin country c in t, t_0 is the base year, and I_{c,i,t_0} is the stock of foreigners with country of origin c in region i in t_0. Hence, the term in brackets is the share of people with country of origin c settling in region i at time t_0. It captures the size of the network of individuals from country c in region i. This share is multiplied by $I_{c,t}$, i.e., the stock of foreigners with country of origin c residing in Switzerland in year t.[27] Finally, the term is summed up over all countries of origin, in order to get a predicted stock of foreign-born individuals in each region i in year $t > t_0$. In our analysis, we use the year 1980 as the base year t_0. The instrument for the interaction term in eq. (1) is, accordingly,

$$z_2 = \frac{\Delta \hat{I}_{i,t}}{Pop_{i,t-1}} \times \mathbf{1}(t \geq 2002). \tag{3}$$

The validity of the instruments is based on the assumption that past migration patterns do not affect current housing prices through anything other than current immigration, i.e., are not correlated with the error term in the structural equation for housing price growth (exclusion restriction). The idea is to consider previous immigrant settlements far enough back in time for them to be independent of current housing demand factors. However, the exclusion restriction would be violated if the initial settlement pattern of migrants (in 1980) was correlated with current outcomes through other factors than present immigration, like region-specific time trends, e.g., stemming from regional adjustments to migration over time, such as out-migration of locals, and region-specific income trends (Saiz, 2007; Jaeger et al., 2018).

Combining the IV approach with the exogenous increase in immigration in response to the reform allows us to separately study the effects of immigrant inflows before and after the reform. As indicated in eq. (1), we therefore interact the immigrant inflow with a post-AFMP-reform dummy and instrument both the main variable and the interaction term. The differential effect of immigration in the post- relative to the pre-reform period can be interpreted as a DiD estimate of the effect of the AFMP reform on house prices. While the estimated coefficient β_1 could potentially absorb unaccounted region-specific trends that jointly affect the historical settlement pattern and housing price dynamics, the DiD effect captured by β_2 should be robust to a violation of the exogeneity condition of the instrument for this reason.

Moreover, we propose a second empirical approach that consists of an event study of the changes in house prices before and after the AFMP reform. With this approach we can verify that before the reform house price changes are indeed unrelated to the historical share of immigrants from EU-15 countries in 1980, which supports the validity of our shift-share instrument.

To further examine the validity and robustness of our results we implement series of sensitivity analyses inspired from the related literature. To mitigate the concern of omitted variable bias invalidating the shift-share instrument, Saiz (2007) includes metropolitan area

[25] The shift-share instrument is used, for example, by Ottaviano and Peri (2007), Saiz (2007), Fischer (2012), Gonzalez and Ortega (2013), Basten and Koch (2015), Sa (2015), and Degen and Fischer (2017) in the context of the housing market, by Dustmann et al. (2005), Card (2007), and Cortes (2008) in labour economics, by Hunt and Gauthier-Loiselle (2010) and Peri (2012) in the context of innovation and productivity, and by Bell et al. (2013) in the context of crime. The first application of the instrument was in Altonji and Card (1991) on the effects of immigration on labor market outcomes.

[26] The instrument has also been referred to as the supply-push or ethnic networks instrument.

[27] For instance, we multiply the share of Italians in region i in year t_0 by the total number of Italians in Switzerland in the year t, to get the predicted number of Italians in region i in the year t.

fixed-effects. We take a similar route by adding dummies for the major regions as robustness check in Sect. 7.[28]

Adãao et al. (2019), Borusyak et al. (2022), and Goldsmith-Pinkham et al. (2020) consider the properties of the shift-share instrument with a "leave-one-out" correction, originally employed in Autor and Duggan (2003) in the context of labor demand shifts, with the goal of strengthening the exogeneity assumption. In our context, this means that factors $I_{c,t}$ in eq. (2) should not contain region i when constructing $\hat{I}_{i,t}$. We show in the Appendix that our results are basically unaffected by the leave-one-out correction. However, the corrected instrumental variable can only be computed from 1990 onwards, i.e., we lose five years of observation at the MS-level. Moreover, for the cantonal level variation, which generally generates more imprecise estimates (as will become apparent), the first-stage results worsen. Thus, we focus in the main text on the uncorrected version of the instrument, in line with the previous literature on migration effects.

Another endogeneity issue is the potential measurement error in the explanatory variable. Immigration to Switzerland is measured by the annual change in the stock of foreigners in Switzerland. However, the number of foreign nationals is not only determined by immigration but also by births and deaths of foreigners as well as naturalizations. If the measurement error is correlated with the observed explanatory variable, namely the annual change in the stock of foreigners, the OLS regression gives a biased and inconsistent estimator (Wooldridge, 2012). To be precise, in this case the estimated coefficients on immigration variables are closer to zero than the true coefficient (attenuation bias). The inconsistency of OLS estimates will only be small if the variance in the unobserved explanatory variable, namely immigration, is large relative to the variance in the measurement error (Wooldridge, 2012).

4.2 Model 2: event study approach

In view of the potential invalidity of the IV approach, despite its common use, we specify a modified reduced form for the outcome variables at the MS-level that exploits the AFMP reform to implement an event study of housing price changes, where we group the regions according to their historical immigrant stocks from EU-15 countries. Specifically, we replace the first two terms on the right-hand side of eq. (1) as follows:

$$\begin{aligned}\Delta(\log P_{i,t}) = {} & \gamma_1 \cdot \mathbf{1}\left(\frac{I^{EU15}_{i,t_0}}{Pop_{i,t_0}} > b\right) \times \mathbf{1}(t \geq 2002) \\ & + \gamma_2 \cdot \mathbf{1}\left(a < \frac{I^{EU15}_{i,t_0}}{Pop_{i,t_0}} \leq b\right) \times \mathbf{1}(t \geq 2002) \\ & + \gamma_3 \cdot \mathbf{1}\left(\frac{I^{EU15}_{i,t_0}}{Pop_{i,t_0}} > b\right) \\ & + \gamma_4 \cdot \mathbf{1}\left(a < \frac{I^{EU15}_{i,t_0}}{Pop_{i,t_0}} \leq b\right) \\ & + \gamma_5 \cdot \Delta X_{i,t-1} + \alpha_t + \varepsilon_{i,t},\end{aligned} \tag{4}$$

where dummy variable $\mathbf{1}\left(I^{EU15}_{i,t_0}/Pop_{i,t_0} > b\right)$ indicates whether the stock of immigrants from EU-15 countries relative to the total population in region i at time $t_0 < t$ is above some threshold $b \in (0,1)$, and $\mathbf{1}\left(a < I^{EU15}_{i,t_0}/Pop_{i,t_0} \leq b\right)$ indicates an intermediate fraction of EU-15 immigrants in period t_0, $0 < a < b$. The other variables are the same as in eq. (1). Data availability dictates to focus on prices ($P_{i,t}$) for owner-occupied housing. As base year t_0 we again choose 1980, underlying the close connection of the historical measure of the immigration stock of a region from EU-15 countries, I^{EU15}_{i,t_0}, and the construction of the instrument z_1 in (2) for the IV approach that contains the historical immigration stocks I_{c,i,t_0} from a broader set of countries. Threshold levels a and b correspond to different percentiles of the regional distribution of the historical EU-15 immigration stock relative to the total population.

If we wrote eq. (4) in levels rather than first differences, the first four terms on the right-hand side would be interacted with a linear time trend. We thus allow for different regional time trends that may vary with the historical exposure to immigrants from the EU, in addition to national time trends (captured by α_t).

The coefficients of interest are γ_1 and γ_2 that estimate the effect of the AFMP reform by distinguishing regions with historically high and medium immigration from EU-15 countries, respectively, from those with low immigration. Again, this strategy is based on the idea of a supply-push effect of immigration to migrant networks with similar national background and that immigration from the EU triggered off by the AFMP has a different effect on prices of owner-occupied housing than pre-reform. A differential effect on housing price growth in regions with a historically higher exposure to immigration from EU-15 countries relative to those with low immigrant exposure

[28] Also recall that we control for the economic situation of a region to estimate eq. (1). As an alternative approach, Jaeger et al. (2018) add lagged immigration to the regression, which they instrument with a lagged version of the familiar instrument. However, the national stock of immigrants by country of origin used to construct the instrument may be highly correlated over time, making the two instruments highly correlated as well. If so, there may be multicollinearity, making causal inference of immigration effects very difficult or even impossible (Jaeger et al., 2018).

after the reform can be interpreted as DiD effect of the AFMP reform.

To show in more detail that there are no heterogeneous pre-trends according to the historical exposure of a region to immigrants from the EU-15 countries, we specify in addition the following event study model:

$$\Delta(\log P_{i,t}) = \alpha_t + \sum_{t=1986}^{2016} \delta_{1,t} \cdot \mathbf{1}\left(\frac{I_{i,t_0}^{EU15}}{Pop_{i,t_0}} > b\right) \times \mathbf{1}(year = t) + \sum_{t=1986}^{2016} \delta_{2,t} \cdot \mathbf{1}\left(a < \frac{I_{i,t_0}^{EU15}}{Pop_{i,t_0}} \leq b\right) \times \mathbf{1}(year = t) + \delta_3 \cdot \Delta X_{i,t-1} + \varepsilon_{i,t}. \quad (5)$$

The estimates of the coefficients $\delta_{1,t}$ for $t \geq 2002$ provide us with the effects of the AFMP reform on the historically highly exposed MS-regions and $\delta_{2,t}$ for $t \geq 2002$ the effects on moderately exposed MS-regions relative to regions with a low exposure to immigrants from EU-15 countries in 1980. We further hypothesize that before the reform the estimates of these coefficients are not significantly different from zero, i.e., $\delta_{1,t} = \delta_{2,t} = 0$ for $t < 2002$, supporting the validity of the shift-share instrument used in Model 1.

5 Descriptive statistics

Figure 2 lends support for the identification strategy of Model 2. It shows that the increase in the stock of foreigners between years 1985 and 2002 relative to the total population in 1985 in MS-regions is basically unrelated to the EU-15 immigration stock relative to the total population (blue dots) in the year 1980, while the change in the stock of foreigners between years 2002 and 2016 relative to the total population in 2002 is positively related to the historical share of EU-15 immigrants in the population.

Figure 3 plots the annual growth rates of the housing price indices (solid lines—left axis) in Switzerland, i.e., the dependent variables of the estimated equations, and the annual change in the stock of foreigners relative to the population size (dotted line—right axis) in Switzerland, $\Delta I_{i,t}/Pop_{i,t-1}$, as used on the right-hand side of eq. (1). We see the surge in house prices in the second half of the 1980 s, that was followed by a recession in Switzerland. The housing boom was related to high immigration, but immigration may have been triggered off by the boom rather than being causal. Immigration lifted off around the time of the AFMP reform.

Table 2 contains summary statistics for the variables used in the analysis with regional variation at the MS-level, distinguishing the pre-reform period 1985–2001 and the post-reform period 2002–2016. On average, prices for single-family homes (*HP*) in Switzerland increased by 1.4% per year before the AFMP reform and 3.0% after the reform. The difference is even more striking for prices of owner-occupied apartments (*AP*), that had an average annual growth rate of 0.9% before the reform and 3.9% after the reform.

Table 3 shows the summary statistics at the cantonal level for the period 1998–2016, where *HP* increased by 2.4% and *AP* by 3.2% per year on average. The annual

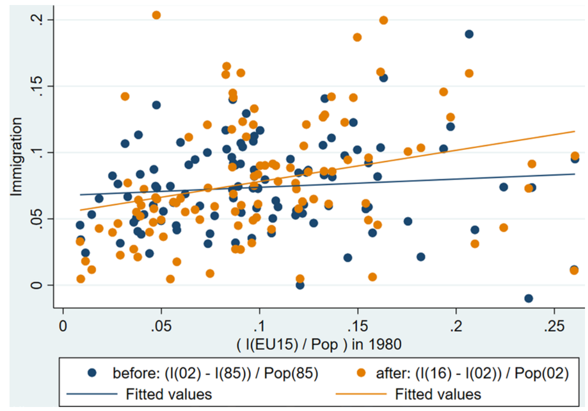

Fig. 2 Immigration before and after the AFMP reform as a function of the share of EU-15 immigrants in 1980. *Notes:* Each dot represents a MS-region. The y-axis displays the change in the stock of foreigners between years 2002 and 1985 relative to total population in 1985 (blue dots) and the change in the stock of foreigners between years 2016 and 2002 relative to total population in 2002 (orange dots). *Source:* Own calculations based on data from Federal Statistical Office

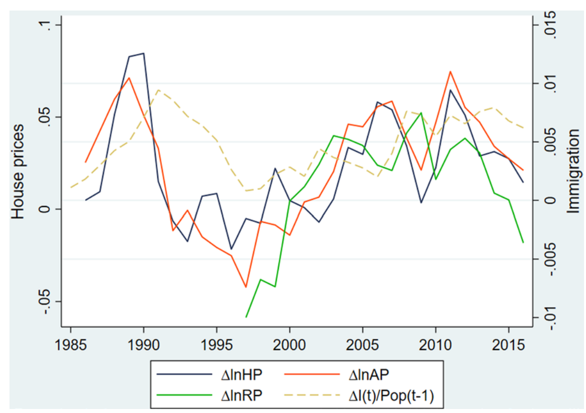

Fig. 3 Growth rate of housing price indices and immigration in Switzerland, 1985–2016. *Notes:* $\Delta \ln x$ denotes the annual change in the log of the price index *x*, where *HP* stands for single-family homes, *AP* for owner-occupied apartments and *RP* for rented apartments. $\Delta I_t/Pop_{t-1}$ is the annual change in the stock of foreigners relative to the initial population. Aggregate data for entire Switzerland. *Source:* Own calculations based on data from Wüest Partner and Federal Statistical Office

Table 2 Descriptive statistics at MS-level, 1985–2016. *Source:* Own calculations based on data from Wüest Partner and Federal Statistical Office

	1985–2001					2002–2016				
	Mean	SD	Min	Max	Obs	Mean	SD	Min	Max	Obs
HP_{it}	119.7	13.4	94.9	160.9	1802	160.0	34.6	103.4	282.0	1590
AP_{it}	119.4	12.4	92.2	151.2	1802	160.4	41.6	89.0	351.4	1590
$\Delta \ln HP_{it}$	0.014	0.038	−0.082	0.175	1696	0.030	0.034	−0.090	0.184	1590
$\Delta \ln AP_{it}$	0.009	0.039	−0.105	0.153	1696	0.039	0.031	−0.079	0.186	1590
Pop_{it}	64897.8	65074.3	4041.0	410715.0	1802	73300.9	71881.2	4042.0	487142.0	1590
I_{it}	11510.4	18068.9	117.0	154078.0	1802	16222.3	23328.7	495.0	197059.0	1590
$\Delta \ln Pop_{it}$	0.008	0.010	−0.048	0.087	1802	0.008	0.007	−0.017	0.033	1590
$\frac{\Delta I_{it}}{Pop_{it-1}}$	0.004	0.005	−0.013	0.024	1802	0.005	0.004	−0.009	0.024	1590
u_{it}	0.023	0.020	0.000	0.078	1797	0.029	0.011	0.007	0.074	1590
Δu_{it}	0.000	0.008	−0.028	0.030	1795	0.001	0.005	−0.018	0.025	1590

For the MS-region "Appenzell I.Rh." the unemployment rate is missing for the years 1985, 1987–1990

Table 3 Descriptive statistics at cantonal level, 1998–2016. *Source:* Own calculations based on data from Wüest Partner and Federal Statistical Office

	Mean	SD	Min	Max	Obs
HP_{it}	154.5	33.2	100.7	257.0	494
AP_{it}	158.0	45.1	101.2	351.4	494
RP_{it}	107.8	24.3	70.5	218.2	479
$\Delta \ln HP_{it}$	0.024	0.032	−0.079	0.125	494
$\Delta \ln AP_{it}$	0.032	0.035	−0.079	0.186	494
$\Delta \ln RP_{it}$	0.016	0.049	−0.175	0.259	479
Pop_{it}	293947.1	304468.4	14880.0	1477197.0	494
I_{it}	63583.6	72209.5	1442.0	389195.0	494
$\Delta \ln Pop_{it}$	0.007	0.006	−0.016	0.023	494
$\frac{\Delta I_{it}}{Pop_{it-1}}$	0.004	0.003	−0.005	0.014	494
u_{it}	0.027	0.013	0.003	0.074	494
$wage_{it}$	5576.8	464.0	4446.0	6671.0	494
CP_{it}	116.0	9.1	100.0	135.0	494
VAC_{it}	1660.7	1704.4	41.0	9309.0	494
Δu_{it}	−0.001	0.007	−0.028	0.025	494
$\Delta \ln wage_{it}$	0.012	0.014	−0.012	0.047	468
$\Delta \ln CP_{it}$	0.011	0.020	−0.030	0.045	468
$\Delta \ln VAC_{it}$	−0.007	0.188	−0.789	1.499	494

For the canton "Appenzell I.Rh." the rental price index is missing for the years 2002–2016

growth rate of prices for rental apartments (*RP*) was somewhat smaller, with by 1.6% per year.

There exists substantial variation across cantons, as displayed in Fig. 4 for prices for owner-occupied apartments. Urban cantons and cantons near bigger cities experienced the strongest growth in housing prices. For example, prices for owner-occupied apartments in Geneva increased by 218% between 1985 and 2016, while in the canton of Jura they increased by only 63%. A similar pattern can be observed for prices for single-family homes and for rental apartments.

Turning to immigration, according to Table 2, the average annual change in the stock of foreigners relative to initial population was 0.4% in the period before and 0.5% after the AFMP reform. Table 3 shows a similar figure (0.4%) at the cantonal level for the period 1998–2016. Again there exists regional variation, as displayed in Fig. 5. A high increase in the immigrant population relative to the total population was again observed in urban cantons and in cantons near bigger cities such as in the cantons of Zug, Vaud, and Fribourg, where a growth of 28% was recorded between 1985 and 2016. Rural cantons experienced much smaller increases. For example, the stock of foreigners relative to population size increased in the cantons of Appenzell by only 6.0% between 1985 and 2016. A similar pattern can be observed for the period from 1998 to 2016, where immigration is on average responsible for more than half of population growth in Switzerland.

Finally, regarding the other control variables, on average across cantons, the monthly gross salary (*wage*) increased on average by 1.2% per year and the construction price index (*CP*) by 1.1%, while vacancies of housing units (*VAC*) decreased by 0.7% (Table 3). The unemployment rate (*u*) changed very little on average, according to both Tables 2 and 3.

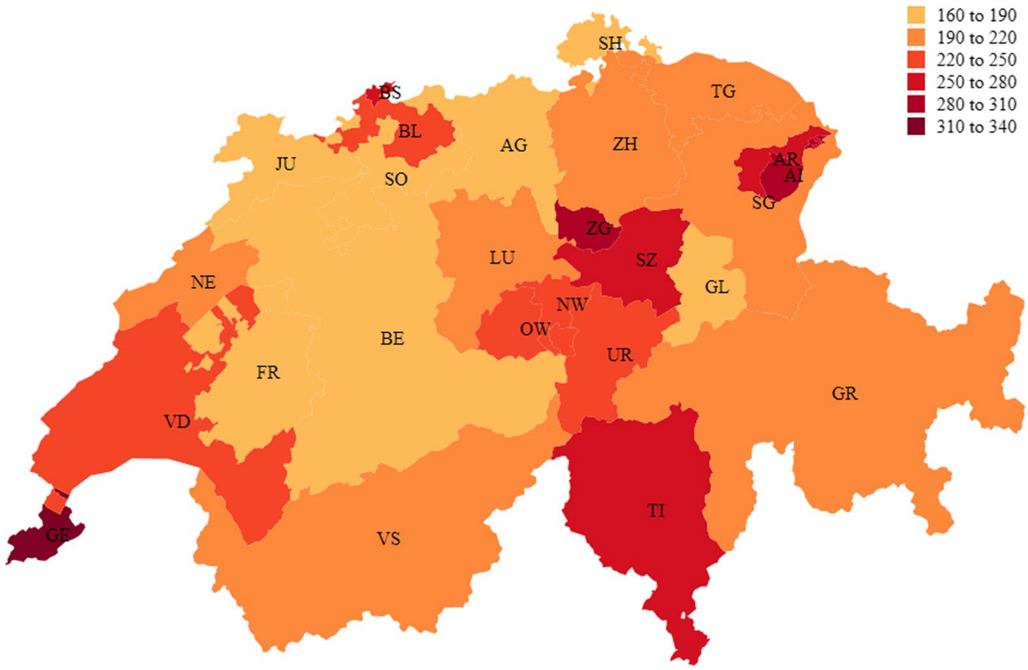

Fig. 4 Price index for owner-occupied apartments in 2016 by canton. *Notes:* Base year = 1985. The range (x to y) includes the lower number (x) and excludes the upper number (y). *Source:* Own calculations based on data from Wüest Partner

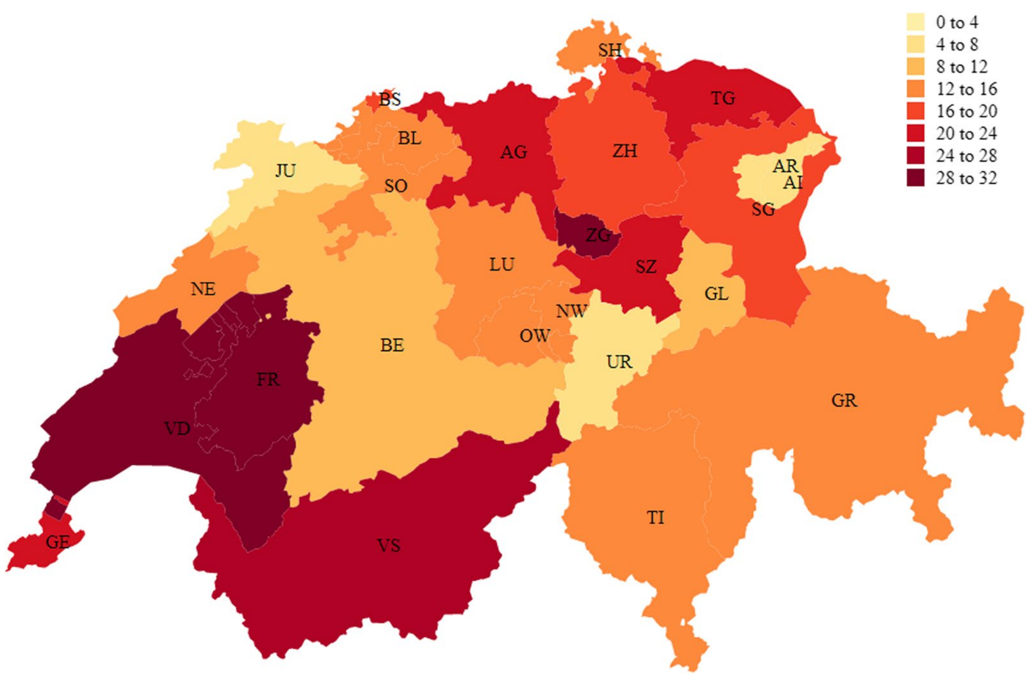

Fig. 5 Growth in the stock of foreigners relative to initial population from 1985 to 2016 in % by canton. *Notes:* Mapped variable: $(I_{i,2016} - I_{i,1985}) \, / \, Pop_{i,1985} \cdot 100$. The range (x to y) includes the lower number (x) and excludes the upper number (y). *Source:* Own calculations based on data from Federal Statistical Office

Table 4 House prices and immigration—MS-regions from 1985–2016: OLS and IV regressions. *Source:* Own calculations based on data from Wüest Partner and Federal Statistical Office

Dep. var.:	**OLS regressions**				**IV regressions**			
	$\Delta \ln HP_{it}$		$\Delta \ln AP_{it}$		$\Delta \ln HP_{it}$		$\Delta \ln AP_{it}$	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
$\frac{\Delta I_{it}}{Pop_{it-1}}$	0.918***	0.449*	0.948***	0.269	1.536***	−2.589***	3.696***	0.478
	[0.000]	[0.076]	[0.000]	[0.266]	[0.006]	[0.008]	[0.000]	[0.473]
$\frac{\Delta I_{it}}{Pop_{it-1}} \times \mathbf{1}(t \geq 2002)$		0.865**		1.251***		6.907***		5.388***
		[0.030]		[0.001]		[0.000]		[0.004]
$\Delta\, u_{it-1}$	−0.206	−0.223	−0.047	−0.072	−0.185	−0.335*	0.047	−0.070
	[0.172]	[0.137]	[0.754]	[0.634]	[0.207]	[0.062]	[0.775]	[0.656]
Year FE	yes	yes	yes	yes	yes	yes	yes	yes
Obs	3279	3279	3279	3279	3279	3279	3279	3279
R^2	0.538	0.540	0.634	0.637	0.535	0.461	0.573	0.554
K-P F-stat					11.52	5.60	11.52	5.60
S-W F-stat, immigration						11.41		11.41
S-W F-stat, interaction						13.25		13.25

p-values in brackets. Standard errors are clustered by MS-regions. The regressions are run at the MS-regional level for the period from 1985–2016. Δ indicates first differences. The instrument is the change in the predicted stock of foreigners divided by initial population. The stock of foreigners is predicted by the settlement pattern of immigrants in 1980. The *K-P F-stat* is the Kleibergen-Paap rk Wald *F*-statistic. The *S-W F-stat* is the Sanderson and Windmeijer (2016) multivariate *F*-test of excluded instruments for weak identification of each endogenous regressor separately; immigration and interaction refer to endogenous regressors $\Delta I_{it}/Pop_{it-1}$ and $\Delta I_{it}/Pop_{it-1} \times \mathbf{1}(t \geq 2002)$, respectively

Significance levels: $^{*}\, p < 0.10$, $^{**}\, p < 0.05$, $^{***}\, p < 0.01$

6 Main results

6.1 Estimates for Model 1: IV approach

6.1.1 MS-region variation and pre- versus post-reform effects

Table 4 reports the OLS and IV estimates of model (1) with MS-regions as observational units. The dependent variable is the annual change in the log of the price index for single-family homes and owner-occupied apartments in a MS-region. The main independent variable is the annual change in the stock of foreigners relative to initial population, without and with interacting it with the post-AFPM reform dummy. All regressions include year fixed effects to capture national trends and are estimated in first differences to control for time-invariant region-specific factors.

Columns 1 and 3 restrict $\beta_2 = 0$ (no interaction effect with post-reform dummy). They show that the OLS estimates of β_1 are highly significant. They suggest that an annual increase in the stock of foreigners equal to 1% (i.e., about two standard deviations) of the initial population leads to an increase in single-family home prices (*HP*) of 0.92% and in owner-occupied apartment prices (*AP*) of 0.95%. The analogous IV estimates presented in columns 5 and 7 reveal higher price increases of 1.5% and 3.7%, respectively.[29] Again, *p*-values are below 1%. When including the immigration variable interacted with the post-AFMP reform dummy, we see that the estimate of coefficient β_2 is highly significant in all specifications, while the estimate of coefficient β_1 becomes small and sometimes insignificant or even negative. For instance, the IV estimates in column 6 suggest that, after the AFMP reform, an increase in the predicted migration variable of 1% raises the prices of single-family homes by $(-2.59 + 6.91 =)$ 4.32%, while the estimated effect is even negative before the reform. According to the IV estimates in column 8, the price effect for owner-occupied apartments is not significantly different from zero before the reform and amounts to $(0.48 + 5.39 =)$ 5.87% after the reform. If anything, a change in the unemployment rate has the expected negative effect, but the coefficients are mostly not significantly different from zero and small in magnitude.

Comparing OLS and IV estimates suggests that the OLS-estimate of β_1 is upward biased whereas the OLS-estimate of β_2 is downward biased. An upward bias of the estimated β_1 may be explained by the booming economy in the second half of the 1980 s − driven by the real estate market − that has caused a large immigration inflow (see Fig. 3) with immigrants locating in regions with high income growth. A downward bias of the estimated β_2 is consistent with the interpretation that immigrants prefer

[29] That the IV estimates of immigration effects on housing costs are considerably larger than the OLS estimates is in line with Saiz (2007), Gonzalez and Ortega (2013), and Degen and Fischer (2017).

Table 5 First-stage regressions. *Source:* Own calculations based on data from Wüest Partner and Federal Statistical Office

Dep. var.:	$\frac{\Delta I_{it}}{Pop_{it-1}}$		$\frac{\Delta I_{it}}{Pop_{it-1}} \times \mathbf{1}(t \geq 2002)$	$\frac{\Delta I_{it}}{Pop_{it-1}}$	
	(1)	**(2)**	**(3)**	**(4)**	**(5)**
$\frac{\Delta \hat{I}_{it}}{Pop_{it-1}}$	0.222***	0.198***	0.000	0.172*	0.172*
	[0.001]	[0.001]	[0.983]	[0.095]	[0.094]
$\frac{\Delta \hat{I}_{it}}{Pop_{it-1}} \times \mathbf{1}(t \geq 2002)$		0.043	0.241***		
		[0.612]	[0.008]		
$\Delta\, u_{it-1}$	−0.020	−0.022	0.000	−0.057	
	[0.595]	[0.561]	[0.983]	[0.525]	
$\Delta \ln wage_{it-1}$					0.033*
					[0.087]
$\Delta \ln CP_{it-1}$					0.012
					[0.431]
$\Delta \ln VAC_{it-1}$					−0.001**
					[0.026]
Year FE	Yes	Yes	Yes	Yes	Yes
Obs	3279	3279	3279	442	442
R^2	0.400	0.400	0.576	0.438	0.444
Years	85-16	85-16	85-16	98-16	98-16
Region	MS	MS	MS	Canton	Canton

p-values in brackets. In columns 1–3, standard errors are clustered by MS-regions. In columns 4–5, standard errors are clustered by cantons and estimated by the wild bootstrap. Δ indicates first differences. The settlement pattern of immigrants in 1980 is used to predict the stock of foreigners in each region and year

Significance levels: $^{*}\ p < 0.10$, $^{**}\ p < 0.05$, $^{***}\ p < 0.01$

to locate in areas where housing costs are growing more slowly.

6.1.2 First-stage

Columns 1–3 of Table 5 report the first-stage estimates corresponding to the IV estimates in Table 4. Column 1 applies for the second stage estimates that restrict $\beta_2 = 0$ (columns 5 and 7 of Table 4). It shows that the coefficient of the predicted stock of foreigners relative to initial population, $z_1 = \Delta\hat{I}_{i,t}/Pop_{i,t-1}$, is highly significant. Columns 2 and 3 show the results without restriction of β_2 to zero. Here, the same is true for the coefficient of $z_1 = \Delta\hat{I}_{i,t}/Pop_{i,t-1}$ when $\Delta I_{i,t}/Pop_{i,t-1}$ is the dependent variable (column 2) and for the coefficient of $z_2 = \Delta\hat{I}_{i,t}/Pop_{i,t-1} \times \mathbf{1}(t \geq 2002)$ when $\Delta I_{i,t}/Pop_{i,t-1} \times \mathbf{1}(t \geq 2002)$ is the dependent variable (column 3), corresponding to second stage estimates in columns 6 and 8 of Table 4.

6.1.3 Cantonal variation and rents

As rental prices are not available for the entire period and only at the cantonal level, Tables 6 and 7 present the OLS and IV estimates for the period 1998–2016 with cantons as observational units, respectively, restricting $\beta_2 = 0$. We control for the economic situation of a region with the annual change in the unemployment rate (columns 1, 3, 5) or the annual changes in the log of the monthly gross wage, the construction price index, and the number of vacant apartments (columns 2, 4, 6), all lagged by one year. The effects for single-family homes (columns 1 and 2) and owner-occupied apartments (columns 3 and 4) show point estimates that are, in particular for the IV estimations, comparable in size to the post-reform effects of immigration in Table 4, where MS-regions rather than cantons were the observational unit. However, because the estimations are quite imprecise given the lower number of observations, the *p*-values for the estimates of β_1 using wild bootstrap standard errors are mostly above 0.1. Also the first-stage estimates of the immigration variable in columns 4 and 5 of Table 5 are significantly different from zero at the 10% level only, albeit similar in size to column 1.

The estimated coefficients on the immigration variable are significantly different from zero though for rental prices, according to columns 5 and 6 of Tables 6 and 7 (at 1% level for OLS and 5% level for IV estimations). According to the OLS estimates (Table 6), an annual increase in the stock of foreigners equal to 1% of the initial population leads to an increase in rental prices of more than 2%. The IV estimates (Table 7) suggest that it leads to rent increase of 7.4%, which is higher than the effect on the other prices.

Table 6 Housing prices and immigration—cantons from 1998–2016: OLS regressions. *Source:* Own calculations based on data from Wüest Partner and Federal Statistical Office

Dep. var.:	**$\Delta \ln HP_{it}$**		**$\Delta \ln AP_{it}$**		**$\Delta \ln RP_{it}$**	
	(1)	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**
$\frac{\Delta I_{it}}{Pop_{it-1}}$	1.881**	1.720**	1.097*	1.089*	2.387***	2.175***
	[0.011]	[0.011]	[0.073]	[0.059]	[0.006]	[0.005]
Δu_{it-1}	-0.406		-0.342		1.592	
	[0.417]		[0.488]		[0.206]	
$\Delta \ln wage_{it-1}$		0.660**		0.535		0.245
		[0.026]		[0.114]		[0.282]
$\Delta \ln CP_{it-1}$		0.314*		-0.183		0.303
		[0.053]		[0.431]		[0.407]
$\Delta \ln VAC_{it-1}$		-0.009		0.000		0.000
		[0.440]		[0.977]		[0.999]
Year FE	Yes	Yes	Yes	Yes	Yes	Yes
Obs	442	442	442	442	427	427
R^2	0.432	0.453	0.407	0.415	0.149	0.145

p-values in brackets. Standard errors clustered by cantons are estimated by the wild bootstrap. The regressions are run at the cantonal level for the period from 1998–2016. Δ indicates first differences

Significance levels: $^{*}\ p < 0.10, ^{**}\ p < 0.05, ^{***}\ p < 0.01$

Table 7 Housing prices and immigration—cantons from 1998–2016: IV regressions. *Source:* Own calculations based on data from Wüest Partner and Federal Statistical Office

Dep. var.:	**$\Delta \ln HP_{it}$**		**$\Delta \ln AP_{it}$**		**$\Delta \ln RP_{it}$**	
	(1)	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**
$\frac{\Delta I_{it}}{Pop_{it-1}}$	3.865	3.925*	6.285	6.743	7.407**	7.382**
	[0.115]	[0.097]	[0.319]	[0.274]	[0.037]	[0.024]
Δu_{it-1}	−0.287		−0.030		1.944	
	[0.557]		[0.958]		[0.109]	
$\Delta \ln wage_{it-1}$		0.607**		0.399		0.124
		[0.041]		[0.167]		[0.600]
$\Delta \ln CP_{it-1}$		0.272*		−0.289		0.197
		[0.061]		[0.218]		[0.595]
$\Delta \ln VAC_{it-1}$		−0.005		0.008		0.009
		[0.503]		[0.331]		[0.588]
Year FE	Yes	Yes	Yes	Yes	Yes	Yes
Obs	442	442	442	442	427	427
R^2	0.406	0.422	0.246	0.226	0.071	0.061
K-P F-stat	21.98	21.90	21.98	21.90	16.91	16.18

p-values in brackets. Standard errors clustered by cantons are estimated by the wild bootstrap. The regressions are run at the cantonal level for the period from 1998–2016. Δ indicates first differences. The instrument is the change in the predicted stock of foreigners divided by initial population. The stock of foreigners is predicted by the settlement pattern of immigrants in 1980. The *K-P F-stat* is the Kleibergen-Paap rk Wald *F*-statistic

Significance levels: $^{*}\ p < 0.10, ^{**}\ p < 0.05, ^{***}\ p < 0.01$

The wage rate has the expected, positive effects on prices, but the estimated coefficient is only significantly different from zero for single-family homes. The other control variables do not play a role.

6.1.4 Interpretation

Our findings support the hypothesis that an inflow of immigrants raises housing demand, especially demand for rental apartments. The results thus meet the

Table 8 House prices and immigration—MS-regions from 1985–2016: Event study. *Source:* Own calculations based on data from Wüest Partner and Federal Statistical Office

Cut-offs (percentiles):	a: 33th, b: 67th				a: 50th, b: 75th			
Dep. var.:	$\Delta \ln HP_{it}$		$\Delta \ln AP_{it}$		$\Delta \ln HP_{it}$		$\Delta \ln AP_{it}$	
	(1)	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**	**(7)**	**(8)**
$\mathbf{1}(\frac{I^{EU15}_{i,1980}}{Pop_{i,1980}} > b) \times \mathbf{1}(t \geq 2002)$	0.010***	0.010***	0.012***	0.011***	0.010***	0.010***	0.011***	0.011***
	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]
$\mathbf{1}(a < \frac{I^{EU15}_{i,1980}}{Pop_{i,1980}} \leq b) \times \mathbf{1}(t \geq 2002)$	− 0.001		0.002		0.001		0.000	
	[0.689]		[0.428]		[0.661]		[0.927]	
$\mathbf{1}(\frac{I^{EU15}_{i,1980}}{Pop_{i,1980}} > b)$	− 0.001	− 0.002	− 0.000	− 0.001	− 0.002	− 0.003*	− 0.001	− 0.001
	[0.380]	[0.113]	[0.773]	[0.335]	[0.102]	[0.061]	[0.728]	[0.452]
$\mathbf{1}(a < \frac{I^{EU15}_{i,1980}}{Pop_{i,1980}} \leq b)$	0.001		0.002		0.001		0.002	
	[0.240]		[0.247]		[0.559]		[0.197]	
$\Delta\, u_{it-1}$	− 0.254	− 0.253	− 0.106	− 0.102	− 0.246	− 0.244	− 0.097	− 0.095
	[0.111]	[0.112]	[0.512]	[0.528]	[0.124]	[0.128]	[0.549]	[0.557]
Year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Obs	3279	3279	3279	3279	3279	3279	3279	3279
R^2	0.537	0.537	0.635	0.634	0.535	0.535	0.633	0.633

p values in brackets. Standard errors are clustered by MS-regions. The regressions are run at the MS-regional level for the period from 1985–2016. Δ indicates first differences. Cut-off levels refer to the respective percentile in the distribution of the stock of EU-15 foreigners relative to total population in year 1980 across MS-regions

Significance levels: $^{*}\, p < 0.10$, $^{**}\, p < 0.05$, $^{***}\, p < 0.01$

expectation that there are no offsetting reactions caused by immigration within the considered, relatively large regional units (MS-regions and cantons). However, the evidence in Table 4 also suggests that immigration significantly raised prices for owner-occupied housing only when the AFMP came into force. There are two related explanations for this finding. First, there has been a changing composition of immigrants toward EU citizens, particularly Germans, who presumably earn more than the average foreigner in Switzerland and thus demand more housing services. Second, in regions with particularly high immigration from the EU caused by the reform, the housing market could not absorb the higher demand anymore, i.e., there is a non-linear effect of an immigration inflow that depends on the stock of immigrants (or the population size in general) in a region. We investigate this further by turning to the event study approach presented in Sect. 4.2.

6.2 Estimates for Model 2: event study approach

Table 8 presents the results of estimating eq. (4). Columns 1 and 3 present the effects of historical immigration from the EU on prices of single-family homes (*HP*) and owner-occupied apartments (*AP*) when we split the MS-regions in three groups of the same size, i.e., threshold levels a and b correspond to the 33th and 67th percentile of the distribution of the EU-15 immigrant population share, respectively. We see that the estimate of γ_1 is highly significant in both columns while estimates of the other coefficients (including the one on the change in the unemployment rate) are insignificant. That the estimates of γ_3 and γ_4 are basically zero suggests there have not been pre-existing trends, lending support to our IV strategy when estimating eq. (1). That the estimate of γ_2 is basically zero makes it appropriate to separate the regions in two categories, with a high level of immigration from the EU-15 (again, according to the 67th percentile of the distribution) and the rest. The results are shown in columns 2 and 4. They suggest that, after the AFMP reform, the annual growth rate of house prices is by one percentage point higher in the regions with a historically high EU-15 immigration stock than in the other regions.

According to columns 5–8 of Table 8, the results are pretty similar when choosing the 50th and 75th percentile of the distribution of the EU-15 immigrant population for a and b, respectively. Switching from a region with a fraction of EU-15 immigrants below the 75th percentile of the distribution to a region belonging to the quarter of regions with the highest fraction of EU-15 immigrants raises the growth rate of house prices by slightly more than one percentage point after the year 2002.

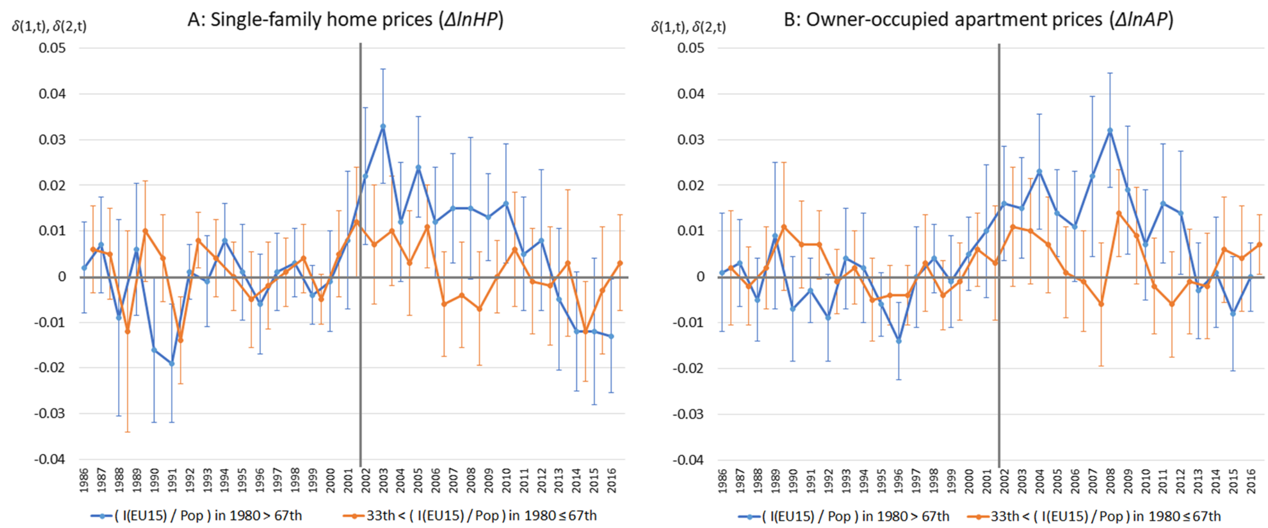

Fig. 6 House prices and immigration—MS-regions from 1985–2016: Event study by year. *Notes:* The figure shows the estimated coefficients $\delta_{1,t}$ and $\delta_{2,t}$ and associated 95% confidence intervals of equation (5). The effect for the highly treated MS-regions ($\delta_{1,t}$) is shown in blue and the effect for the moderately treated MS-regions ($\delta_{2,t}$) is shown in orange. In Panel A, the dependent variable is the annual change in the log of the price index for single-family homes. In Panel B, the dependent variable is the annual change in the log of the price index for owner-occupied apartments. All regressions include year fixed effects and the annual change in the unemployment rate lagged by one year as a control variable. Standard errors are clustered by MS-regions. The 67th and 33rd percentiles of the distribution of the stock of EU-15 foreigners relative to total population in year 1980 across MS-regions are used as cut-off levels. *Source:* Own calculations based on data from Wüest Partner and Federal Statistical Office

Figure 6 provides a graphical representation of the estimated coefficients $\delta_{1,t}$ and $\delta_{2,t}$ from regression eq. (5) in the case of single-family homes (Panel A) and owner-occupied apartments (Panel B). It confirms the conjecture that before the reform immigration did not have a significant effect on prices of owner-occupied housing. In line with the estimates shown in Table 8, this evidence again supports the validity of the exclusion restriction of the shift-share instrument used in the IV estimations of Model 1. Importantly, the evidence also suggests significant impacts of immigration on housing prices in the highly treated regions up until about 10 years after the AFMP reform. For the moderately treated regions, the AFMP reform effects were mostly insignificant. These insights are robust to using the alternative cut-offs as in Table 8 to classify the regions (available upon request). Overall, short-run housing shortages triggered off by the immigration shock appear to have vanished after an extended adjustment period.

7 Sensitivity analyses and reduced forms for Model 1

To further probe the robustness of our estimates for Model 1 (IV approach) we conduct several sensitivity analyses that we present in Tables 9, 10, 11, 12 and in the next subsections. Finally, we implement the specification of Degen and Fischer (2017) with our data and present the results of the reduced form regression for housing prices corresponding to the main results for Model 1 (IV approach).

7.1 Adding fixed effects for major regions

Table 9 focusses on the estimates with post-reform interaction effects of immigration at the MS-regional level. Column 1 repeats the baseline estimates of the coefficients of interest β_1 and β_2 in the price regressions (1) of single-family homes (*HP*) and owner-occupied apartments (*AP*) from columns 6 and 8 of Table 4, respectively.

Column 2 of Table 9 presents the results when we add fixed effects for the seven major regions (Fig. 10 in the Appendix). We thereby aim at controlling for region-specific dynamic effects of the initial settlement pattern of immigrants on the housing market, thus raising confidence in the exclusion restriction (Saiz, 2007). We find that OLS results are similar to the baseline estimates. The IV estimates suggest that, after the AFMP reform, an increase in the predicted migration variable of 1% raises the prices of single-family homes by $(-2.74 + 7.61 =)$ 4.87% and those of owner-occupied apartments by $(0.93 + 6.72 =)$ 7.65%, compared to the 4.32% and 5.87% (baseline estimates) when not controlling for major region fixed effects, respectively. The estimated coefficient β_2 of the interaction term is, again, highly significant. Moreover, there is no evidence of an effect of immigration on *HP* and *AP* before the reform.

Table 9 House prices and immigration—MS-regions from 1985–2016: Sensitivity analysis. *Source:* Own calculations based on data from Wüest Partner and Federal Statistical Office

	Baseline (1)	**Major reg. FE (2)[1]**	**One year lag (3)**	**3-year diff. (4)[2]**	**Net mig. (5)[3]**	**$\leqslant$ 20% sec. homes (6)[4]**
OLS regressions						
Dep. var.: Δ ln HP_{it}						
β_1	0.449*	0.422*	0.550*	0.265	0.102	0.147
	[0.076]	[0.096]	[0.064]	[0.332]	[0.637]	[0.652]
β_2	0.865**	0.851**	0.723*	1.477***	1.162***	1.461**
	[0.030]	[0.034]	[0.084]	[0.001]	[0.000]	[0.014]
Dep. var.: Δ ln AP_{it}						
β_1	0.269	0.301	0.315	0.193	0.187	0.038
	[0.266]	[0.216]	[0.212]	[0.471]	[0.306]	[0.887]
β_2	1.251***	1.254***	1.352***	1.666***	1.310***	1.659***
	[0.001]	[0.001]	[0.000]	[0.000]	[0.000]	[0.001]
IV regressions						
Dep. var.: Δ ln HP_{it}						
β_1	− 2.589***	− 2.736***	− 2.421**	− 2.748***	− 2.093***	− 1.987**
	[0.008]	[0.008]	[0.014]	[0.008]	[0.005]	[0.047]
β_2	6.907***	7.614***	5.921***	7.149***	3.872***	7.254***
	[0.000]	[0.001]	[0.001]	[0.000]	[0.000]	[0.007]
Dep. var.: Δ ln AP_{it}						
β_1	0.478	0.934	− 0.229	0.646	0.402	0.065
	[0.473]	[0.346]	[0.738]	[0.396]	[0.445]	[0.926]
β_2	5.388***	6.721***	5.350***	4.945**	2.014***	6.657**
	[0.004]	[0.007]	[0.001]	[0.013]	[0.002]	[0.012]
Obs	3279	3279	3279	3067	3279	2318
K-P F-stat	5.60	4.93	5.87	5.91	14.87	2.22
S-W F-stat, immigration	11.41	14.63	12.06	14.83	31.97	20.02
S-W F-stat, interaction	13.25	13.25	14.44	14.32	47.67	20.47

p values in brackets. Standard errors are clustered by MS-regions. The regressions are run at the MS-regional level for the period from 1985–2016. Δ indicates first differences. β_1 is the estimated coefficient on the annual immigration (ΔI_{it}) relative to the initial population size (Pop_{it-1}). β_2 is the coefficient on the interaction of $\Delta I_{it}/Pop_{it-1}$ with $\mathbf{1}(t \geq 2002)$. All regressions include year fixed effects and the annual change in the unemployment rate lagged by one year as a control variable. The instrument is the change in the predicted stock of foreigners divided by initial population. The stock of foreigners is predicted by the settlement pattern of immigrants in 1980. In columns 3, the instrument is likewise lagged by one year. The *K-P F-stat* is the Kleibergen-Paap rk Wald *F*-statistic. The *S-W F-stat* is the Sanderson and Windmeijer (2016) multivariate *F*-test of excluded instruments for weak identification of each endogenous regressor separately; immigration and interaction refer to endogenous regressors $\Delta I_{it}/Pop_{it-1}$ and $\Delta I_{it}/Pop_{it-1} \times \mathbf{1}(t \geq 2002)$, respectively

[1] Column 2 additionally includes fixed effects for major regions

[2] In column 4, a 3-year difference is applied to all variables, e.g.: $\ln HP_{it} - \ln HP_{it-3}$ or $(I_{it} - I_{it-3})/Pop_{it-3}$

[3] In column 5, international net migration of foreigners is used instead of the annual change in the stock of foreigners to measure ΔI_{it}

[4] In column 6, only MS-regions with a share of second homes of 20% or less are considered for the estimation

Significance levels: $^{*}\,p < 0.10$, $^{**}\,p < 0.05$, $^{***}\,p < 0.01$

We also experimented with adding major region fixed effects to the estimates of Model 2 (event study), that we have presented in Table 8. Results were very similar (not shown).

7.2 Alternative timing

As housing prices may not react immediately to immigration, we next report estimates with a lagged migration variable and a time difference of three years (between years $t-3$ and t) for both the explanatory variables and the dependent variable instead of annual changes.

In column 3 of Table 9, the annual change in the stock of foreigners relative to population size is lagged by one year, i.e., we replace $\Delta I_{i,t}/Pop_{i,t-1}$ by $\Delta I_{i,t-1}/Pop_{i,t-2}$ in eq. (1). The OLS estimates change very little. Moreover, the estimated sum of coefficients $\beta_1 + \beta_2$ in the IV regressions (post-reform effect of immigration) is only slightly lower than the baseline estimates. The estimated β_1 (pre-reform effect) is non-positive.

Table 10 Housing prices and immigration—cantons from 1998–2016: Sensitivity analysis. *Source:* Own calculations based on data from Wüest Partner and Federal Statistical Office

	Baseline (1)	**One year lag (2)**	**3-year diff. (3)[1]**	**Net mig. (4)[2]**	**$\leqslant$ 20% sec. homes (5)[3]**	**2002–2016 (6)**
OLS regressions						
Dep. var.: $\Delta \ln HP_{it}$						
β_1	1.881**	1.775***	2.242***	1.516**	1.856**	1.639**
	[0.011]	[0.009]	[0.009]	[0.012]	[0.028]	[0.019]
Dep. var.: $\Delta \ln AP_{it}$						
β_1	1.097*	1.374**	1.398	1.697***	1.317*	1.313**
	[0.073]	[0.044]	[0.101]	[0.005]	[0.072]	[0.026]
Dep. var.: $\Delta \ln RP_{it}$						
β_1	2.387***	1.793*	2.650***	2.085***	2.680***	2.865***
	[0.006]	[0.071]	[0.004]	[0.001]	[0.009]	[0.006]
IV regressions						
Dep. var.: $\Delta \ln HP_{it}$						
β_1	3.865	3.146	4.602*	1.290*	4.199	4.100
	[0.115]	[0.200]	[0.060]	[0.094]	[0.177]	[0.164]
Dep. var.: $\Delta \ln AP_{it}$						
β_1	6.285	5.373	7.278	2.097	8.057	6.365
	[0.319]	[0.148]	[0.264]	[0.306]	[0.176]	[0.248]
Dep. var.: $\Delta \ln RP_{it}$						
β_1	7.407**	5.403	7.665*	2.218**	10.321**	8.045**
	[0.037]	[0.113]	[0.052]	[0.043]	[0.033]	[0.029]
Obs, HP & AP	442	442	442	442	357	390
Obs, RP	427	427	427	427	342	375
K-P F-stat, HP & AP	21.98	35.68	46.30	182.99	15.33	19.57
K-P F-stat, RP	16.91	28.52	38.81	178.09	9.89	14.58

p values in brackets. Standard errors clustered by cantons are estimated by the wild bootstrap. The regressions are run at the cantonal level for the period from 1998–2016, except in column 6. Δ indicates first differences. β_1 is the estimated coefficient on the annual immigration (ΔI_{it}) relative to the initial population size (Pop_{it-1}). All regressions include year fixed effects and the annual change in the unemployment rate lagged by one year as a control variable. The instrument is the change in the predicted stock of foreigners divided by initial population. The stock of foreigners is predicted by the settlement pattern of immigrants in 1980. In columns 2, the instrument is likewise lagged by one year. The *K-P F-stat* is the Kleibergen-Paap rk Wald *F*-statistic

[1] In column 3, a 3-year difference is applied to all variables, e.g.: $\ln HP_{it} - \ln HP_{it-3}$ or $(I_{it} - I_{it-3})/Pop_{it-3}$

[2] In column 4, international net migration of foreigners is used instead of the annual change in the stock of foreigners to measure ΔI_{it}

[3] In column 5, only cantons with a share of second homes of 20% or less are considered for the estimation

Significance levels: $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$

When we consider a time difference of three years (between years $t-3$ and t) instead of annual changes (column 4 of Table 9), the estimated β_1 and β_2 are again similar to those in the baseline specification. This suggests that there is no large measurement error in the year-to-year estimates.

Table 10 presents sensitivity analysis with cantonal variation for the period 1998–2016 (where we restrict $\beta_2 = 0$), with the change in the unemployment rate as control variable X (analogously to Table 9). Column 1 restates the baseline estimates from Table 6 (OLS regressions) and Table 7 (IV regressions). We again see little difference to the baseline estimates with a time difference of three years (column 3 of Table 10). With the lagged migration variable (column 2), the point estimates of β_1 somewhat shrink and the estimated coefficient of interest in the IV regression becomes insignificant also for rental prices (*RP*).

7.3 Employing net immigration of foreigners

A potential issue is the employed approximation of the immigration flow by a change in the stock of foreigners rather than a change in the foreign-born population.[30]

[30] Children of foreigners who are born in Switzerland are typically non-citizens at the time of birth. Thus, on the one hand, we also capture the impact of births (and deaths) on housing demand by looking at changes in the stock of foreign nationals rather than that of the foreign-born population. On the other hand, we do not account for naturalizations.

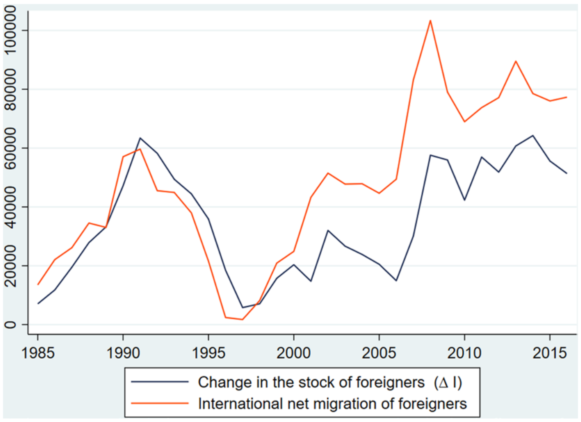

Fig. 7 Change in the stock of foreigners versus international net migration of foreigners in Switzerland, 1985–2016. *Notes:* Aggregate data for entire Switzerland. *Source:* Own calculations based on data from Federal Statistical Office

It is commonly used in the literature on immigration and the housing market for data availability reasons and makes our results on house price effects comparable to the previous literature. Alternatively, we may measure the immigration flow as the number of foreigners settling in Switzerland minus the number of foreigners leaving the country in a certain period (net migration). Because measuring immigrants and emigrants in this way again contains foreigners that are born in Switzerland, it is still different to the change in the foreign-born population though. Figure 7 displays the time series of annual net migration and the change in the stock of foreigners for the period 1985–2016. It shows that the former is increasing faster than the latter from the end of the 1990 s onwards, while both series are similar before. The increasing difference in both series is consistent with a change in the law on the acquisition of Swiss citizenship in 1992 that facilitated naturalizations of foreigners married to a Swiss citizen and keeping the foreign citizenship after naturalization.

Column 5 of Table 9 and column 4 of Table 10 report the results when we replace the annual change in the stock of foreigners employed in the baseline specification with net international migration of foreigners. The IV analysis still employs the same instrument (2) (based on the change in the predicted stock of foreigners) as in the baseline estimates. The first-stage estimates suggest that it is equally relevant (not shown).

At the MS-level, the estimated OLS coefficients β_1 and β_2 in column 5 of Table 9 are quite similar to the baseline estimates. However, the estimated IV coefficients β_2 are considerably lower than in column 1, while the suggested immigration effects before the reform remain similar. The estimated $\beta_1 + \beta_2$ suggest that after the reform an annual net migration of foreigners of 1% relative to the initial population causes *HP* to increase by $(-2.09 + 3.87 =)$ 1.78% and *AP* to increase by $(0.40 + 2.01 =)$ 2.41%.

At the cantonal level, column 4 of Table 10 also suggests that the IV coefficients of the net migration variable are lower than instrumenting the change in the stock of foreigners. (Again, the OLS coefficient of interest is similar to the one in column 1.) For instance, an annual net migration of foreigners of 1% relative to the initial population raises rental prices (*RP*) by 2.2% (rather than 7.4% in column 1). The estimated coefficient β_1 (restricting $\beta_2 = 0$) is still significantly different from zero at the 5% level. The estimated coefficients for the other prices (*HP* and *AP*) are similar in magnitude as the corresponding estimates in Table 9.

The lower post-reform immigration effects when employing the change in net migration rather than the change in the stock of foreigners is not surprising, as Fig. 7 suggested a higher immigration boom after the AFMP reform using the former measure. Thus, a change in the number of foreigners relative to the total population of 1% corresponds to a higher change than 1% in the net migration variable, thus mechanically leading to higher price effects. As naturalizations increased considerably over the considered time period and particularly since the end of the 1990 s where the two time series displayed in Fig. 7 diverge, the net migration variable may approximate the change in the foreign-born population better.[31] Instrumenting the change in the number of foreigners as immigration measure rather than the net migration variable with the shift-share instrument (2) based on the stock of foreigners seems more natural though.

7.4 Addressing the second home initiative

The acceptance of the second home initiative implied a construction ban of new second homes in municipalities with more than 20% of second homes. Column 6 of Table 9 and column 5 of Table 10 provide the estimates when excluding those MS-regions and cantons with more than 20% of second homes in 2012, respectively. Immigration is again measured by the annual change in the number of foreigners. At the MS-level, we see that the estimated β_1 and β_2 change relatively little. According to the IV estimates, an increase in the predicted migration variable of 1% after the AFMP reform raises the prices of single-family homes by $(-1.99 + 7.25 =)$ 5.26% and of owner-occupied apartments by $(0.07 + 6.66 =)$

[31] The average number of naturalizations of persons living in Switzerland was 6'261 per year in the period 1987–1991, 15'921 per year in the period 1992–1999, and 37'091 per year in the period 2000–2016. See State Secretariat for Migration SEM (2022).

Table 11 Housing prices and immigration—MS-regions from 2001–2006: OLS and IV regressions. *Source:* Own calculations based on data from Wüest Partner and Federal Statistical Office

	OLS regressions				**IV regressions**			
Dep. var.:	**Δ ln HP_{it}**		**Δ ln AP_{it}**		**Δ ln HP_{it}**		**Δ ln AP_{it}**	
	(1)	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**	**(7)**	**(8)**
$\frac{\Delta I_{it}}{Pop_{it-1}}$	1.821***	2.091***	2.002***	2.086***	7.493***	7.804***	6.288**	5.982***
	[0.000]	[0.000]	[0.000]	[0.000]	[0.006]	[0.004]	[0.014]	[0.004]
$\Delta\ u_{it-1}$	0.228	0.357	0.649	0.738	− 0.355	− 0.163	0.208	0.384
	[0.598]	[0.376]	[0.257]	[0.157]	[0.395]	[0.724]	[0.710]	[0.465]
Year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Maj. reg. FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Obs	510	636	510	636	510	636	510	636
R^2	0.566	0.532	0.496	0.489	0.340	0.277	0.358	0.353
K− P F-stat					12.89	14.35	12.89	14.35
Regions	85 MS	106 MS	85 MS	106 MS	85 MS	106 MS	85 MS	106 MS

p values in brackets. Standard errors are clustered by MS-regions. The regressions are run at the MS-regional level for the period from 2001–2006. Δ indicates first differences. Maj. reg. FE are fixed effects for major regions. The instrument is the change in the predicted stock of foreigners divided by initial population. The stock of foreigners is predicted by the settlement pattern of immigrants in 1997. The *K-P F-stat* is the Kleibergen-Paap rk Wald *F*-statistic

Significance levels: $^{**}p < 0.05$, $^{***}p < 0.01$

6.73%, which is slightly (less than one percentage point) higher than the baseline estimates (column 1). At the cantonal level, the estimated β_1 suggests that rental prices are affected considerably more by immigration when we exclude cantons with a high share of second homes. Moreover, the suggested downward bias of the OLS coefficient becomes higher.

7.5 Cantonal level estimates for time period 2002–2016

We have so far looked for the cantonal level estimates on the period 1998–2016 for data availability reasons. One may ask whether results change if we confine the analysis on the time period after the AFMP reform, i.e., on 2002–2016. Column 6 of Table 10 shows that the OLS estimates of β_1 remain pretty similar to the baseline estimates in column 1. IV estimates are slightly higher, consistent with a higher effect of immigration after the AFMP reform, albeit significance levels do not change.

7.6 Comparison to Degen and Fischer (2017)

Albeit our focus is on the role of the AFMP reform on housing costs, it is interesting to compare our results to Degen and Fischer (2017), who consider the effect of immigration on prices for single-family homes and owner-occupied apartments in the period 2001–2006. Like we do, they measure immigration by the change in the number of foreigners. Moreover, they use a similar IV strategy, with the only difference that the base year t_0 for constructing the instrument (2) is 1997 rather than 1980. However, they use house price index data from the *Informations- und Ausbildungszentrum für Immobilien* (IAZI) rather than the data from *Wüest Partner* and restrict the data set to the 85 MS-regions that had a residential population of at least 25’000 inhabitants in 2001 (excluding mostly mountain regions). Noteworthy, aggregate house and rental price indices for Switzerland as a whole evolved quite similar for both time series (comparing our Fig. 1 with Figure 2 in Degen and Fischer (2017).

We now take up their suggestion that a “direct comparison between the two indexes would be desirable and is left for future study” (Degen and Fischer, 2017, p.22). That is, we now replicate their empirical strategy of estimating eq. (1) with restriction $\beta_2 = 0$ for the change in the two house price index variables *HP* (for single-family homes) and *AP* (for owner-occupied apartments) as dependent variables, the change in the unemployment rate, year fixed effects, and major region fixed effects as control variables, choosing base year 1997 when constructing the instrument, and sometimes restricting the data set to the 85 (rather than 106) MS-regions they consider.

Results are displayed in Table 11. We see that the estimated β_1 on the immigration variable is highly significant and positive also for the OLS regressions and, again, much higher for the IV regressions. Our instrument seems relevant, with a similar Kleibergen–Paap $F-$statistic (first-stage weak identification test) than Degen and Fischer (2017).[32] Whether or not we include

[32] Slight deviations from Degen and Fischer (2017) at the first-stage are possible because we use updated FSO data on the population by region and year for the time period under consideration.

Table 12 Housing prices and immigration: reduced-form regressions. *Source:* Own calculations based on data from Wüest Partner and Federal Statistical Office

Dep. var.:	**Δ ln HP_{it}**	**Δ ln AP_{it}**	**Δ ln HP_{it}**		**Δ ln AP_{it}**		**Δ ln RP_{it}**	
	(1)	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**	**(7)**	**(8)**
$\frac{\Delta\hat{I}_{it}}{Pop_{it-1}}$	− 0.513***	0.095	0.666	0.676	1.084	1.161	1.105**	1.089***
	[0.002]	[0.457]	[0.119]	[0.116]	[0.227]	[0.221]	[0.032]	[0.003]
$\frac{\Delta\hat{I}_{it}}{Pop_{it-1}} \times \mathbf{1}(t \geq 2002)$	1.552***	1.317***						
	[0.000]	[0.000]						
$\Delta\, u_{it-1}$	− 0.277*	− 0.079	− 0.508		− 0.389		1.441	
	[0.081]	[0.623]	[0.366]		[0.454]		[0.286]	
$\Delta \ln wage_{it-1}$				0.737**		0.623		0.357
				[0.017]		[0.106]		[0.161]
$\Delta \ln CP_{it-1}$				0.319**		− 0.209		0.304
				[0.037]		[0.329]		[0.425]
$\Delta \ln VAC_{it-1}$				− 0.010		− 0.000		− 0.002
				[0.319]		[0.970]		[0.920]
Year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Obs	3279	3279	442	442	442	442	427	427
R^2	0.538	0.636	0.417	0.442	0.418	0.429	0.142	0.140
Years	85–16	85–16	98–16	98–16	98–16	98–16	98–16	98–16
Region	MS	MS	Canton	Canton	Canton	Canton	Canton	Canton

p values in brackets. In columns 1 and 2, standard errors are clustered by MS-regions. In columns 3–8, wild bootstrap standard errors are clustered by cantons. Δ indicates first differences. The stock of foreigners is predicted by the settlement pattern of immigrants in 1980

Significance levels: $^{*}\, p < 0.10, ^{**}\, p < 0.05, ^{***}\, p < 0.01$

the mountain regions makes little difference. Most strikingly, our IV estimates are almost three times as high for single-family homes and twice as high for apartments. According to column 5 and 7 of Table 11, an increase in the foreign population relative to initial population raises *HP* by 7.5% and *AP* by 6.3%, compared to 2.6% and 2.8% in Degen and Fischer (2017), respectively. Moreover, *p*-values in Table 11 are below or close to 1%, whereas the estimated β_1 in Degen and Fischer (2017) is barely significantly different from zero.

We can thus draw two lessons from the comparison. First, despite the fact that the house price indices from *Wüest Partner* and IAZI evolve quite similar in the aggregate, the results are quite sensitive to the data set chosen. Second, comparing Table 11 with our results from Table 4, the estimated immigration effects for the period 2001–2006 are somewhat larger than the post-reform estimates with the longer data set until 2016. This suggests that the effects were particularly high shortly after the AFMP reform.

7.7 Reduced-form estimates

Finally, Table 12 shows the reduced-form estimates for all main specifications of eq. (1), where we regress the annual change in the log of the housing price indices on the excluded instruments, i.e., on the annual change in the predicted stock of foreigners relative to the initial population, $z_1 = \Delta\hat{I}_{i,t}/Pop_{i,t-1}$, and, for the MS-level estimations (period 1985–2016), also on $z_2 = \Delta\hat{I}_{i,t}/Pop_{i,t-1} \times \mathbf{1}(t \geq 2002)$. Since the reduced-form coefficient is the product of the first-stage coefficient and the coefficient in the second-stage (structural) equation, it can only be different from zero if both coefficients are different from zero. The reduced-form estimates are thus important to gain confidence in the causal interpretation of the IV estimates (Angrist and Pischke, 2009).

The signs as well as the significance levels of the estimated coefficients of interest (β_1 and β_2) of the reduced-form regressions are similar to those of the two-stage least squares estimations. Overall, the results support the relevance of the instruments and the existence of a causal, positive effect of immigration on housing costs after the AFMP reform. However, the magnitudes of the estimated β_1 and β_2 are considerably smaller. Columns 1 and 2 suggest that an increase in z_1 by 1% raises the price for owner-occupied housing by only somewhat more than 1% after the AFMP reform. For the cantonal variation, similarly, an increase in z_1 by 1% raises rental prices, again, by slightly more than 1%. The effects on owner-occupied housing are similar, but the estimated β_1 is not significantly different from zero when standard errors clustered by cantons are estimated by wild bootstrap.

8 Conclusion

We have studied the short-run effects of immigration into Switzerland on housing prices, distinguishing single-family homes, owner-occupied apartments, and rented apartments. We proposed two empirical strategies, an IV approach and an event study. Both exploit the Agreement on the Free Movement of Persons (AFMP) with the European Union (EU), enacted in 2002, as a source of exogenous variation in immigration. The results suggest that the growth of the foreign population has had a sizable positive impact on the prices of single-family homes and owner-occupied apartments after the AFMP reform came into effect, but not pre-reform. A possible reason for the higher effects of immigration on house prices after the reform is the change in the composition of immigrants toward EU citizens, who often are better skilled and earn more than other immigrants. Moreover, the AFMP reform induced short-run housing shortages because the housing demand increases triggered off by immigration were not fully anticipated.

The immigration effect on apartment prices are somewhat higher than on prices of single-family homes, according to the IV estimates at the MS-level, using the widely-used shift-share instrument based on historical immigration patterns. At the cantonal level, the effect of immigration is higher on rental prices than on prices for owner-occupied housing. This may be explained by the fact that immigrants usually first move into rented apartments when arriving in Switzerland (Graf et al., 2010). At the same time, however, rents have risen less strongly on average than prices for owner-occupied housing for the considered time period. The estimated housing demand effects of immigration were also somewhat higher in regions with less than 20% of second homes. The estimated effects were lower when measuring immigration by the difference between immigration of foreigners and emigration of foreigners rather than the change in the number of foreigners over time. Ideally, future research should aim to measure immigration by net immigration of the foreign-born population.

The event study analyzed the differential price change in owner-occupied housing between regions with a historically high, medium and low stock of EU-15 immigrants in interaction with the AFMP reform. It suggests that in regions with a historically high stock of EU immigrants, the annual growth rate of prices for owner-occupied housing was one percentage point higher than in the other regions after 2002. Interestingly, however, this difference in price growth eventually vanished about 10 years after the AFMP reform. In both the IV approach and the event study, adding fixed effects for major regions changes the results only marginally.

Our results have potentially important policy implications. Despite the undisputed positive effects of (particularly) high-skilled immigration on labor market outcomes and the economic development of an advanced economy such as Switzerland (e.g., Beerli et al., 2021, Grossmann, 2021), associated increases in housing prices particularly harm low-income individuals who do not own housing property. Ignoring these effects can generate resistance to liberal migration policies, as observed in Switzerland and elsewhere. For instance, the exit of Great Britain from the EU, that has generated harmful labor shortages, serves as a prime example that policy makers need to address distributional effects of immigration. Compensating measures via the tax-transfer system and deregulation of zoning restrictions to stimulate housing construction could tackle distributional consequences of immigration in the medium run and help avoiding political backlashes to the free movement of labor in Europe.

Appendix

Short-run housing supply effects of immigration

To back our demand-side explanation of the effects of immigration on housing costs, we now demonstrate that there are no positive short-run effects of immigration on housing supply.

Empirical model

Using regional variation at the MS-level and municipal level we estimate the following model in first differences (annual changes):

$$\Delta(\log Q_{i,t}) = \eta_1 \cdot \frac{\Delta I_{i,t}}{Pop_{i,t-1}} + \eta_2 \cdot \Delta(\log wage_{i,t-1}) + \alpha_t + \varepsilon_{i,t}, \tag{6}$$

where $Q_{i,t}$ is the number of housing units in region i and year t. We distinguish the total number of apartments ($Q^T_{i,t}$), apartments in single-family-homes ($Q^H_{i,t}$), and flats in apartment buildings ($Q^A_{i,t}$). Again, $\Delta I_{i,t}/Pop_{i,t-1}$ is the annual change in the stock of foreigners relative to the initial population size. As for the housing costs, we will alternatively replace $\Delta I_{i,t}/Pop_{i,t-1}$ with the net international migration of foreigners (*NetMig*) relative to population, $NetMig_{i,t}/Pop_{i,t-1}$. For the IV regressions, both

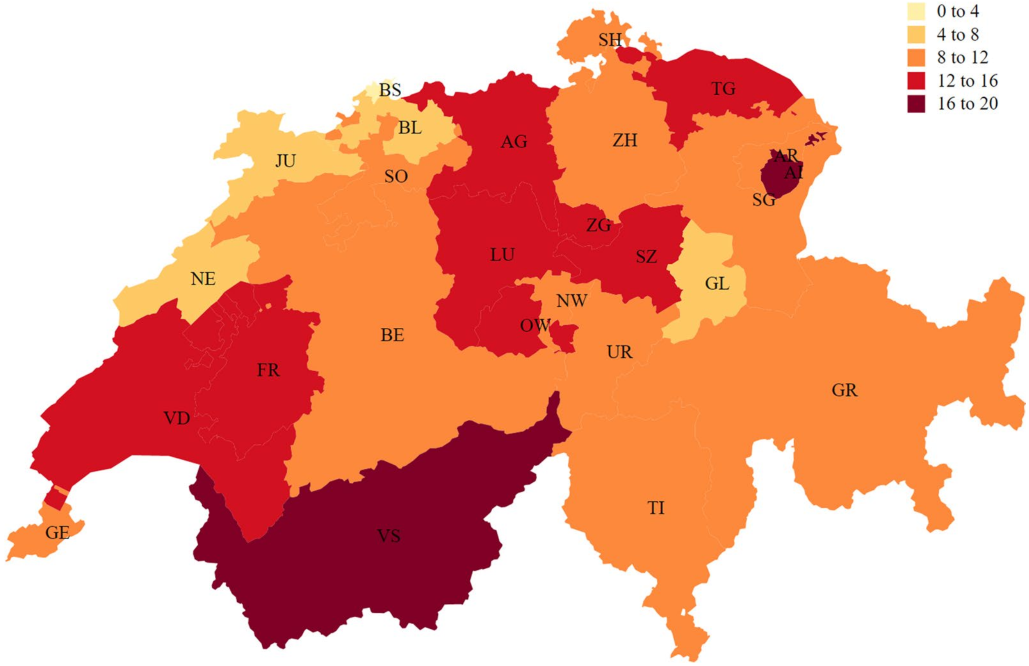

Fig. 8 Growth in housing supply from 2009 to 2016 in % by canton. *Notes:* Mapped variable: $(Q^T_{i,2016} - Q^T_{i,2009}) / Q^T_{i,2009} \cdot 100$. The range (x to y) includes the lower number (x) and excludes the upper number (y). *Source:* Own calculations based on data from Federal Statistical Office

migration variables are instrumented as before, using (2) with base year 1980. $wage_{i,t-1}$ is the (lagged) average monthly gross wage, α_t are year fixed effects to control for national trends, and $\varepsilon_{i,t}$ is the error term.[33] The first differences again account for time-invariant region-specific factors. Standard errors are heteroscedasticity-robust and clustered at the analyzed regional level in all regressions.

Data

To measure changes in housing supply, we employ data by region and year from the buildings and dwellings statistics of the Federal Statistical Office (FSO). The other data sources are the same as described in Sect. 3 and in Table 1. The data is available annually from 2009 to 2016 at the municipal level. The number of apartments in a region is counted on December 31 each year. The data distinguishes between apartments in single-family homes and in apartment buildings (Federal Statistical Office, 2017b).[34]

[33] Unlike in the price estimations presented in Tables 6 and 7, the change in the number of vacant apartments is not included in the estimation on housing supply because of endogeneity issues. We also left out the construction price index because its coefficients were never statistically significant and point estimates close to zero.

[34] An apartment is by definition the totality of rooms that form a structural unit. An apartment has its own access and cooking facilities.

Table 13 Descriptive statistics at MS-level, 2009–2016. *Source:* Own calculations based on data from Federal Statistical Office

	Mean	SD	Min	Max	Obs
Q^T_{it}	39562.6	37519.6	4138.0	229551.0	848
Q^H_{it}	9055.7	6143.6	840.0	39394.0	848
Q^A_{it}	22176.1	23160.3	1826.0	139113.0	848
$\Delta \ln Q^T_{it}$	0.015	0.008	− 0.024	0.063	742
$\Delta \ln Q^H_{it}$	0.008	0.011	− 0.017	0.139	742
$\Delta \ln Q^A_{it}$	0.020	0.014	− 0.111	0.104	742
Pop_{it}	75961.3	74578.0	4042.0	487142.0	848
I_{it}	17728.5	24921.2	533.0	197059.0	848
$\Delta \ln Pop_{it}$	0.010	0.007	− 0.017	0.033	848
$\frac{\Delta I_{it}}{Pop_{it-1}}$	0.006	0.004	− 0.005	0.024	848
$\frac{NetMig_{it}}{Pop_{it-1}}$	0.009	0.004	− 0.005	0.027	848
$wage_{it}$	5981.4	334.6	4983.0	6671.0	848
$\Delta \ln wage_{it}$	0.009	0.011	− 0.012	0.044	848

Descriptive statistics

Table 13 contains descriptive statistics at the level of MS-regions. On average, the number of apartments in total grew by 1.5% per year in the period from 2009 to 2016, which is mainly driven by the increase of flats in apartment buildings. On average, the number of apartments in apartment buildings rose by 2.0% per year, while the number of single-family houses grew by 0.8% per year.

Figure 8 displays the cantonal variation for the growth rate of the total number of apartments. We see that housing supply increased particularly little in urban areas like Basel-Stadt, pointing to land scarcity as limiting factor of housing construction.[35]

Table 13 also shows that the average annual change in the stock of foreigners relative to initial population was 0.6%. The annual average growth of the total population was 1.0%. Monthly gross wages grew by 0.9%. As background information, it may be noted that during this period, prices for single-family homes grew on average at the MS-level by 3.0% per year and prices for owner-occupied apartments by 4.0% per year. At the cantonal level, rental prices grew by 2.0% per year.

Regarding the migration variables, we see that the average annual change in the stock of foreigners relative to initial population, $\Delta I_{it}/Pop_{it-1}$, is 0.6%, while the alternative migration variable, $NetMig_{it}/Pop_{it-1}$, increased by 0.9%. That net migration increased more for the time period of consideration is also visible in Fig. 7 and consistent with naturalizations.[36]

Estimation results

Column 1 of Table 14 presents both the OLS and IV estimates of eq. (6), where the dependent variable is the annual change in the log of the stock of apartments in a MS-region.[37] Although OLS estimates of the coefficient of interest, η_1, are positive and highly significant except for apartments in single-family-homes (Q^H_{it}), the IV estimates are all negative.[38] An upward bias in the OLS estimates is consistent with immigrants settling in regions where housing supply is growing rapidly.

According to column 2, the same picture emerges when we control for major region fixed effects (like in Table 9). While house price data is only available for MS-regions and cantons, column 3 presents the estimates of the impact of immigration on housing supply at the municipal level.[39] The OLS estimates of η_1 are now all positive and highly significant. However, the

Table 14 Housing supply and immigration—MS-regions from 2009–2016. *Source:* Own calculations based on data from Federal Statistical Office

	Baseline	**Major region FE**	**Municipal level**	**Net migration**
	(1)	**(2)[1]**	**(3)**	**(4)[2]**
OLS regressions				
Dep. var.: $\Delta \ln Q^T_{it}$				
η_1	0.440***	0.405***	0.457***	− 0.091
	[0.000]	[0.001]	[0.000]	[0.400]
Dep. var.: $\Delta \ln Q^H_{it}$				
η_1	− 0.108	− 0.144	0.115***	− 0.176
	[0.525]	[0.386]	[0.004]	[0.219]
Dep. var.: $\Delta \ln Q^A_{it}$				
η_1	0.785***	0.720***	0.744***	− 0.502**
	[0.001]	[0.001]	[0.000]	[0.025]
IV regressions				
Dep. var.: $\Delta \ln Q^T_{it}$				
η_1	− 2.068*	− 4.604	0.062	− 0.628***
	[0.064]	[0.227]	[0.917]	[0.000]
Dep. var.: $\Delta \ln Q^H_{it}$				
η_1	− 1.468	− 4.145	− 1.672***	− 0.446*
	[0.144]	[0.194]	[0.000]	[0.062]
Dep. var.: $\Delta \ln Q^A_{it}$				
η_1	− 4.332**	− 9.311	− 0.715	− 1.315***
	[0.044]	[0.219]	[0.411]	[0.000]
Obs	742	742	15526	742
K-P F-stat	4.48	1.57	16.49	35.95

p-values in brackets. Standard errors are clustered by MS-regions, resp. municipalities in column 3. The regressions are run at the MS-regional level (in column 3 at the municipal level) for the period from 2009-2016. Δ indicates first differences. η_1 is the estimated coefficient on the annual immigration (ΔI_{it}) relative to the initial population size (Pop_{it-1}). All regressions include year fixed effects and the annual change in the log of the monthly gross wage lagged by one year as a control variable. The instrument is the change in the predicted stock of foreigners divided by initial population. The stock of foreigners is predicted by the settlement pattern of immigrants in 1980. The *K-P F-stat* is the Kleibergen-Paap rk Wald *F*-statistic

[1] Column 2 additionally includes fixed effects for major regions

[2] In column 4, international net migration of foreigners is used instead of the annual change in the stock of foreigners to measure ΔI_{it}

Significance levels: $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$

[35] We also have data for 2000, which confirms the picture. The largest increases in housing supply between 2000 and 2016 were registered in the cantons of Valais (45%), Schwyz (38%), and Fribourg (38%), while the urban cantons of Basel-Stadt (2.0%) and Geneva (12%) recorded the lowest increases.

[36] In the period 2009–2016, the average number of naturalizations of persons living in Switzerland was 37'502 per year (State Secretariat for Migration SEM, 2022).

[37] We checked whether the results were sensitive to using the annual change in the stock of apartments divided by initial population ($\Delta Q_{it}/Pop_{it-1}$) as measure of a change in housing supply, following Sa (2015), instead of using the annual change in the log of the stock of apartments. Overall, the sign of the coefficients and *p*-values are similar.

[38] With respect to the first-stage, the $F-$statistics suggest relevance of the instrument for all specifications displayed in Table 14.

[39] Compared to MS-regions, municipalities in Switzerland are small. On average, at the municipal level, the total number of apartments grew by 1.6% per year, the number of apartments in multifamily buildings by 2.5% per year, and the number of single-family houses by 1.0% per year in the period from 2009 to 2016. The average annual change in the stock of foreigners relative to initial population was 0.5%. These descriptive statistics are thus similar to the ones at the MS-regional level, displayed in Table 13.

Table 15 Houseing prices and immigration—MS-regions from 1990-2016: LOO instrument. *Source:* Own calculations based on data from Wüest Partner and Federal Statistical Office

Dep. var.:	IV regressions				First-stage regressions		
	$\Delta \ln HP_{it}$		$\Delta \ln AP_{it}$		$\frac{\Delta I_{it}}{Pop_{it-1}}$		$\frac{\Delta I_{it}}{Pop_{it-1}} \times \mathbf{1}(t \geq 2002)$
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
$\frac{\Delta I_{it}}{Pop_{it-1}}$	2.498***	-2.285	4.487***	0.846			
	[0.003]	[0.230]	[0.000]	[0.563]			
$\frac{\Delta I_{it}}{Pop_{it-1}} \times \mathbf{1}(t \geq 2002)$		6.664**		5.074**			
		[0.019]		[0.024]			
Δu_{it-1}	− 0.057	− 0.283	0.106	− 0.065	− 0.031	-0.035	− 0.000
	[0.714]	[0.107]	[0.571]	[0.705]	[0.402]	[0.334]	[0.997]
$\frac{\Delta \hat{I}^{LOO}_{it}}{Pop_{it-1}}$					0.210***	0.158***	− 0.000
					[0.003]	[0.002]	[0.997]
$\frac{\Delta \hat{I}^{LOO}_{it}}{Pop_{it-1}} \times \mathbf{1}(t \geq 2002)$						0.084	0.242**
						[0.322]	[0.011]
Year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Obs	2754	2754	2754	2754	2754	2754	2754
R^2	0.449	0.371	0.537	0.531	0.411	0.412	0.536
K-P F-stat	9.42	4.94	9.42	4.94			
S-W F-stat, immigration		9.91		9.91			
S-W F-stat, interaction		11.69		11.69			

p-values in brackets. Standard errors are clustered by MS-regions. The regressions are run at the MS-regional level for the period from 1990-2016. Δ indicates first differences. The instrument is the change in the predicted stock of foreigners divided by initial population. The stock of foreigners is predicted by the settlement pattern of immigrants in 1980. In contrast to before, here the respective MS-region i is not taken into account in $I_{c,t}$ when estimating the predicted stock of foreigners *(LOO: leave-one-out estimator)*. The *K-P F-stat* is the Kleibergen-Paap rk Wald F statistic. The *S-W F-stat* is the Sanderson and Windmeijer (2016) multivariate *F*-test of excluded instruments for weak identification of each endogenous regressor separately; immigration and interaction refer to endogenous regressors $\Delta I_{it}/Pop_{it-1}$ and $\Delta I_{it}/Pop_{it-1} \times \mathbf{1}(t \geq 2002)$, respectively

Significance levels: $^{**}p < 0.05$, $^{***}p < 0.01$

IV estimates of η_1 suggest that the short-run impact of immigration on the total number of apartments is basically zero and, if anything, negative when looking at the two subgroups of apartments. Finally, in column 4 we report the results when we replace the annual change in the stock of foreigners employed in the baseline specification (shown in column 1) with net international migration of foreigners. Compared to the baseline estimates in column 1, the IV estimates are still negative and significantly different from zero, but smaller in magnitude. Moreover, now not only the IV estimates but also the OLS estimates are non-positive (albeit not significantly different from zero except for apartments in apartment buildings).

Overall, we do neither see a robust positive or negative effect of immigration on the supply of housing units in the short-run, as expected given the amount of time needed for construction. This justifies the focus of our housing cost analysis on demand effects. The housing supply estimates also suggest that construction is correlated with omitted factors attracting immigrants. For instance, housing supply could be driven by expectations of higher future housing demand that cannot be satisfied in big cities because of limited land supply (reflected in high land prices). Indeed, the largest increases in housing supply were in the cantons of Valais, Schwyz and Fribourg, i.e., in metropolitan areas near big cities.

IV results at MS-level with leave-one-out correction

Table 15 re-estimates the IV analysis in Table 4 with the leave-one-out correction when constructing the instrumental variables, as described in Sect. 4.1.2, at the MS-level. It also presents first-stage results, analogously to Table 5. As discussed, for data availability reasons we can only compute the corrected instrumental variable from

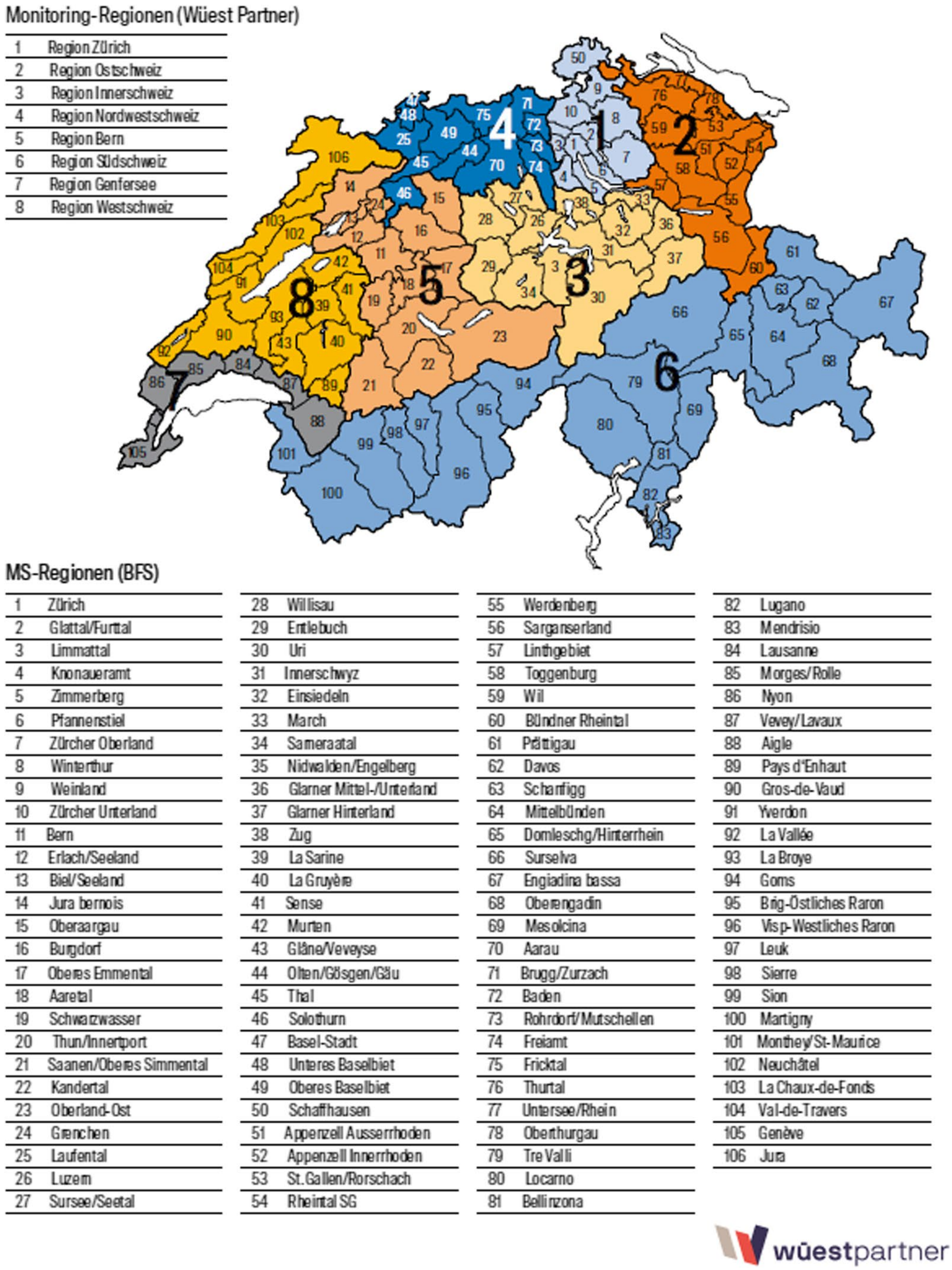

Fig. 9 MS-regions in Switzerland. *Source:* Wüest Partner (2019)

1990 onwards, i.e., we confine the analysis to the period 1990-2016.

Comparing columns 1-4 of Table 15 with columns 5-8 of Table 4, we see that the IV estimate of β_1 when restricting $\beta_2 = 0$ is somewhat higher, whereas the combined effect of the immigration variable after the reform $(\beta_1 + \beta_2)$ in columns 2 and 4 is very similar to the respective counterparts without the leave-one-out correction (columns 6 and 8 of Table 4). Also first-stage results in columns 5-7 are quite similar to the respective counterparts in columns 1-3 of Table 5.

Regions in Switzerland

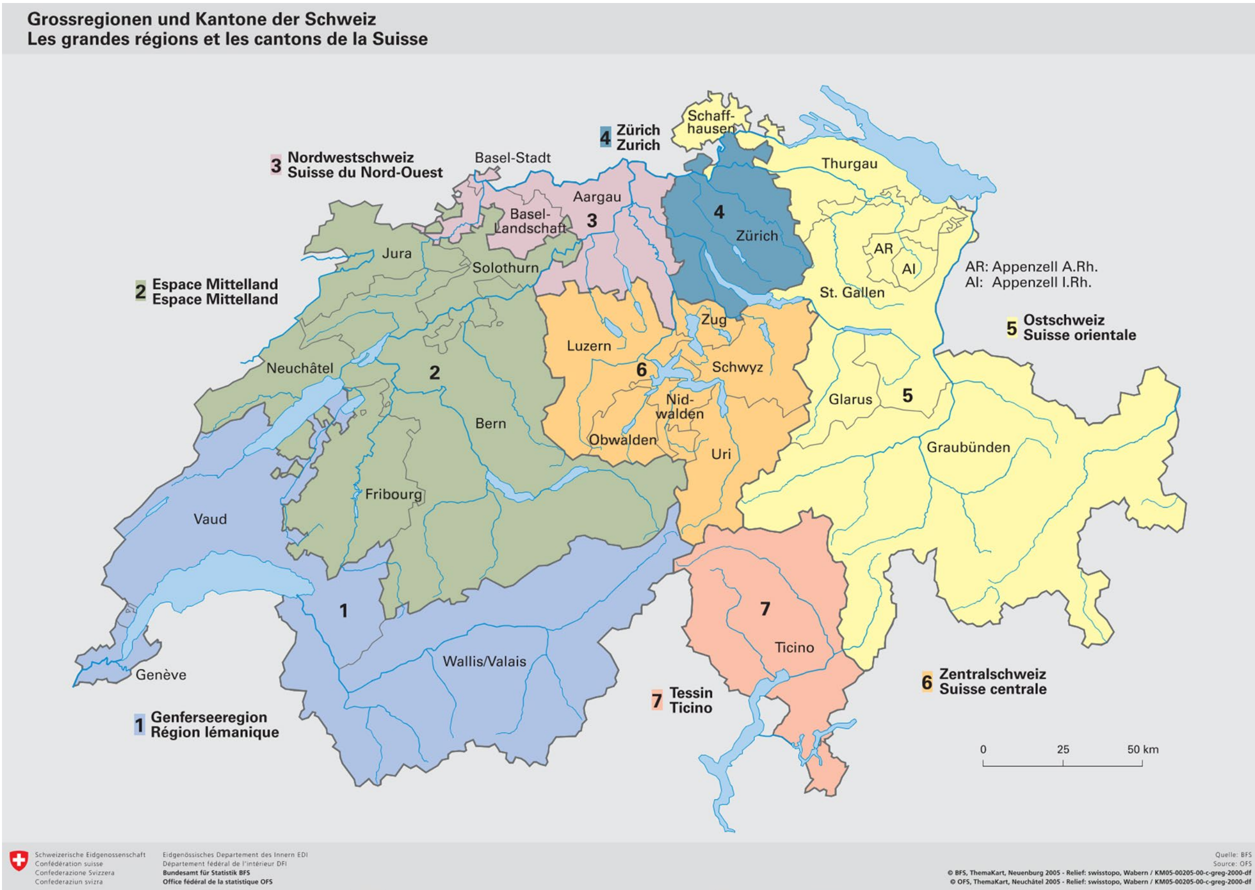

Fig. 10 Cantons and major regions in Switzerland. *Source:* Federal Statistical Office (2019).

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Author contributions
All authors contributed to concept, research questions, methodology, and writing of the paper. FH collected and analyzed the data. All authors read and approved the final manuscript.

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Availability of data and materials
The data on house prices that support the findings of this study are available from *Wüest Partner* but restrictions apply to the availability of these data, which were used under license for the current study, and so are not publicly available. All other data are available from Fabienne Helfer upon reasonable request.

Declarations

Competing interests
The authors declare that they have no competing interests.

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