Home Away From Home? Foreign Demand and London House Prices*

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Abstract

What explains house price variation in global cities like London and New York? A widely-held view is that foreign demand, especially during crises, is part of the explanation, but the rarity of crises renders pure time-series methods ineffective at evaluating this explanation. Our new approach is based on the insight that foreigners exhibit “home bias abroad,” concentrating demand in particular areas within global cities. Using the approach on large databases of housing transactions in London over two decades, we find that foreign risk strongly affects house prices and transactions volumes. The effects are long-lasting but temporary, and vary interestingly across countries.

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1 Introduction

Real estate prices around the world have come to be viewed as particularly volatile over the past two decades, exhibiting dramatic booms and occasional busts. A widely-held view is that foreign capital is at least partly responsible for these gyrations in prices, especially in global cities such as London, New York, or Singapore, which are relatively open to money and immigration from overseas. Residential real estate is the single largest asset held by most households, and so this question is deeply political. As a consequence, commentary on the purported effects of foreign capital on house prices has become increasingly strident.\(^1\)

Despite the widespread interest in this issue, there is little empirical evidence to link foreign demand with house prices. One important challenge is that precise data tracking cross-border transactions in residential real estate is generally not available. Another challenge is that demand deriving from capital preservation or “safe haven” motives generally appears during periods of country-specific or global crisis. Historical time-series data is short relative to the frequency of crises, and during crises, economic variables tend to move in lockstep. These issues make it problematic to use pure time-series methods to convincingly attribute price movements to price pressure from demand, rather than to movements in price-relevant information, the impact of regulation, variation in credit availability, or other crisis-induced market frictions.

Our paper brings a new approach to this question. We begin with the insight that foreign capital directed towards residential real estate may exhibit “home bias abroad.” In other words, capital from a source country may be primarily directed towards houses located in areas of the destination country with a high concentration

of source-country-origin residents. This insight lies at the heart of our method, which estimates the impacts of foreign demand on house prices using the cross-section of house prices in addition to the time-series. We view this methodology as more generally useful to identify demand effects on asset prices in any situation in which investors have different “preferred habitats” within a broad asset class.²

We employ this method on large databases of historical housing transactions in London, a city which is on the frontlines of this issue. London has received substantial attention for its unusual house price appreciation during the great recession, which has routinely been explained as the effect of safe-haven demand from foreign buyers.³ We provide evidence that foreign demand, especially from Southern Europe, the Middle East, and the Asia-Pacific region, is indeed an important part of the explanation.

To illustrate our identification strategy, in late 2009 and early 2010, there were large shocks to economic and political risk in Greece. To detect whether this generates Greek demand for properties in London, we conjecture that areas of London with relatively high pre-existing shares of Greek-born residents are preferred locations for Greek property purchases. If this conjecture is correct, following heightened risk in Greece, we would expect to see relatively higher prices in these specific areas of London, over and above the general level of London house prices.

We generalize this insight, sub-dividing London into smaller geographical areas, and enumerating the strength of the links of each London sub-area with specific foreign countries. When risk in a foreign country rises, our specifications forecast intra-London

²Preferred habitats have been used extensively to explain the term structure of interest rates (see, for example, Culbertson, 1957, Modigliani and Sutch, 1966, and Vayanos and Vila, 2009), and individual investor preferences for stocks with similar volatility as their pre-existing holdings (Dorn and Huberman, 2010).

rates of price appreciation in proportion to the strength of the links between London areas and the foreign country. We measure these links using the share of the population of each London sub-area (624 electoral wards in our empirical application) which originates from different countries; one can envision a wide range of other possible indicators of the attractiveness of particular areas of London to specific countries.

Why is it sensible to link particular foreign countries to areas of London in this way? First, at least some portion of housing demand is likely to be driven by calculations of subsequent physical movements, i.e., the potential for subsequent immigration to London. If so, the cultural affinity of prospective immigrants for their future neighbourhoods in London is very likely to affect their real estate demand. This line of reasoning does not require the immigration to materialize, only that it is a factor influencing property selection. That said, we find evidence that patterns of growth in electoral ward-level immigration shares in our data are consistent with this line of reasoning.\(^4\)

Second, even if an implicit or explicit future immigration motive is not a factor affecting purchasing decisions, there may be social network effects associated with foreign-origin settlement in particular London areas. These may help to lower informational asymmetries in property purchases. This could occur through direct communication between foreign-origin local London residents and overseas buyers. This could also occur because of specialty realtors, local legal firms, and other soft infrastructure set up

\(^4\)Interestingly, Saiz and Wachter (2011) find evidence from the US that growing immigration density leads to native flight and slower rates of house price appreciation, and Sa (2013) finds that immigration is associated with declines in house prices in local authorities in the UK between 2003 and 2010. This evidence makes it more likely that the “safe-haven for capital” channel is what we detect in our empirical work, rather than the direct immigration channel.
to match overseas purchasers of specific nationalities with property investment opportunities in specific London areas.\textsuperscript{5}

In our empirical work, we also consider the possibility that safe-haven demand is concentrated in desirable London regions, driven by high-net worth foreigners purchasing premium properties in response to risk in their home countries. Using this additional identification, we also find, over and above the foreign-origin channel, that London areas with high average income levels experience unusually high prices following increases in global political risk.\textsuperscript{6}

The foreign demand effects that we estimate are likely too low, as they do not contain effects which have homogenous impacts across London areas. Nevertheless, we find that they are large – in London areas with high (top quintile) shares of people originating from a particular country, we find that prices are approximately 1.70 percentage points higher in months following elevated (top quintile) risk in that country. Controlling for this effect, on average, house prices in London areas with higher average incomes are elevated by an additional 2.86 percentage points following rises in global risk. The foreign demand price-impacts that we estimate are long-lived but transitory, becoming statistically indistinguishable from zero after roughly two years. These effects survive a number of robustness checks and placebo tests.

We are also able to use our method to predict variation in London housing transactions volumes. When we instrument for housing transactions volume using foreign risk

\textsuperscript{5}See, for example, London-Tokyo property services (http://www.london-tokyo.co.uk/en/aboutus.php), and Celestial Globe (http://www.celestialglobe.co.uk/en/) which have been established to help prospective Japanese and Chinese buyers, respectively, in the London property market. The branch locations and property listings of each of these organizations appear concentrated in particular London wards.

\textsuperscript{6}All of our specifications control for the influence of a large set of hedonic property characteristics that are likely to influence London house prices, using the now standard approach of Rosen (1974). A recent example of the use of hedonic regressions to explain house prices is Campbell et al. (2011). Meese and Wallace (1997) discuss the benefits of the hedonic pricing method, relative to one based on repeat sales (for examples of the latter, see Case and Shiller (1987), and Bollerslev et al. (2013)).
we are well able to predict prices with instrumented volume. This suggests that foreign demand might also be helpful in explaining the well-documented association between house prices and transactions volumes (see, for example, Stein, 1995).

Finally, we uncover imprecisely estimated but intriguing variation in foreign demand effects emanating from different countries. Risk in Russia, parts of Africa, and the Middle East predicts price increases in premium areas of London, but not in areas in which there are pre-existing residents originating from these regions of the world. In contrast, risk in Southern European and some South Asian countries tends to generate price impacts in regions of London with higher shares of residents originating from their countries, but it appears that these countries are less likely to direct capital towards premium residential areas.

Our work is related to a number of different strands of the literature. Our identification strategy relies on “home bias abroad,” which connects our study to much work on home-bias and loyalty based portfolio choice (see, for example, Lewis 1999, Coval and Moskowitz, 2001, and Cohen, 2009). Our work also forms part of the burgeoning literature on the determinants of real estate prices (for a survey, see Ghysels et al., 2012). Our methodological approach is similar in spirit to Mian and Sufi (2009), who use zipcode level geographical variation across the United States to measure the impact of credit supply factors on household mortgage indebtedness and residential property markets. Our approach is also similar to Saiz and Wachter (2011), but our work is distinguished by our focus on the impacts of foreign demand on real estate prices. Aizenman and Jinjarak (2009), Jinjarak and Shefrin (2011), and Favilukis et al. (2013) also connect foreign capital flows with house prices—our work is distinguished by our new cross-sectional approach, and our implementation using transactions-level data.

Our work is also related to the literature on international capital flows and contagion. Using a wide range of methodologies and data sets, this literature finds that there is a strong relationship between risk and institutional quality in source countries, and
their extent of outbound direct and portfolio investment (see, for example, Alfaro et al. 2008, and Forbes and Warnock 2011). Our findings confirm that risk in a number of world regions predicts house prices in London over and above variation in hedonic characteristics, suggesting that these cross-border investments generate price pressure in destination country assets. Our results also suggest that cross-border capital flows and lending relationships can act as a channel for the international transmission of risk, joining a growing list of papers including Kaminsky et al. (2004), Broner et al. (2006), Jotikasthira et al. (2012) and Schnabl (2012).7 Our results are also related to the literature on estimating safe-haven effects in a variety of assets, which, with few exceptions, has tended to use time-series rather than cross-sectional approaches to this problem.8 Finally, our results provide some insights into the famous Lucas (1990) puzzle about the reason for the limited flow of capital from relatively rich to relatively poor countries.9 One reason for the “uphill” flow of capital from poor to rich countries may be the desire of a subset of poor-country residents to insure themselves against risk in their home countries.

The remainder of the paper is organized as follows. Section 2 describes our methodology, and Section 3 the data that we employ. Section 4 discusses the results from our empirical estimation, and Section 5 concludes. An online appendix contains supplementary results and robustness checks, and is available at the URL specified in the references section.

7Also see Claessens and Forbes (2001) and Karolyi (2003) for surveys of the literature on international financial contagion.
8See, for example, Longstaff (2004), Caballero and Krishnamurthy (2009), Beber et al. (2009), Campbell et al. (2010), Baur and McDermott (2010), and Ranaldo and Soderlind (2010).
9See, for example, Prasad, Rajan, and Subramanian (2007), and Carroll and Jeanne (2009) among others for more recent empirical evidence and theoretical rationales for this puzzle.
2 Methodology

This section describes our methodology for identifying the impacts of foreign purchases on London residential real estate prices. Our strategy uses variation in political risk in “source” countries to identify episodes of high foreign demand. Our key insight is that specific sub-regions of London populated by people originating from a particular country are potentially preferred destinations for capital arriving from the country, assuming that capital exhibits “home bias abroad.”

Our approach may be more broadly applicable to detecting the effects of demand on asset prices. If it is possible to identify investor-group-specific demand for sub-categories of a broad asset class, cross-sectional variation in “preferred habitats” can be combined with time-series information on increases in demand to improve identification.

2.1 A Hedonic Pricing Model with Foreign Demand

Consider the following hedonic pricing model for residential properties in London:

\[
\ln P_{i,t} = \alpha + \Pi_{w,t} + \beta X_{i,t} + u_{i,t}.
\]  

Here, \( P_{i,t} \) is the price of property \( i \) (which is physically located in location \( w \)), measured in month \( t \). \( X_{i,t} \) is a vector of hedonic characteristics for property \( i \), and the second component on the right-hand side of equation (1), denoted \( \Pi_{w,t} \), denotes unrestricted location (\( w \)) cross time (\( t \)) fixed effects.

Our approach is to restrict these location-time fixed effects \( \Pi_{w,t} \) in an economically meaningful fashion. These restrictions allow us to identify the impact of foreign risk (which varies both across regions of the world, as well as through time), on the cross-location cross-time variation in London house prices.

More formally, let \( z_{k,t-1} \) be a measure of risk for a specific country \( k \), in period
We model $\Pi_{w,t}$ in the following fashion:

$$
\Pi_{w,t} = \rho_1 \Pi_{w,t-1} + \rho_2 \Pi_{w,t-2} + \delta_t + \phi_w + \sum_{k \in K} \gamma_t^{k} f^{k}_w z^{k}_t - 1 + \gamma_1 y_w z_{t-1}.
$$

(2)

In equation (2), $\delta_t$ are time fixed effects which eliminate common time-variation in London house prices as a source of identification; $\phi_w$ are location-specific fixed effects eliminating fixed price differentials across London locations as a source of identification, and $f^{k}_w$ links location $w$ with country $k$. In our empirical implementation, $f^{k}_w$ is the share of residents in London electoral ward $w$ (we occasionally refer to these simply as “wards”) that were born in country $k$, but it is easy to envision other useful “link” variables. The remaining coefficients are purely estimated off simultaneous cross-time- and cross-ward variation in house prices.

In the appendix, we show that estimating the model described in equations (1) and (2) is equivalent to estimating the following specification:

$$
\ln P_{i,t} = \mu + \delta_t + \phi_w + \sum_{k \in K} \gamma_t^{k} f^{k}_w z^{k}_t - 1 + \gamma_1 y_w z_{t-1} + \beta X_{i,t}
$$

$$
+ \varphi_1 \ln P_{w,t-1} + \varphi_2 \ln P_{w,t-2} + \nu_{i,t}.
$$

(3)

In equation (3), $\gamma_t^{k}$ measures the impact of foreign demand on London house prices. We expect this to be positive if foreign risk drives capital flows towards London locations which are linked to these countries, generating price impact. The simultaneous inclusion of multiple countries $k$ allows us to separately estimate the marginal impacts of risk from each country on specific locations within London.

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$^{10}$In our empirical estimation, we aggregate countries $k$ to world regions to reduce the number of separate parameters estimated. We represent $z$ as a lagged indicator, to allow for the possibility that economic and political conditions may build up over a period of time. In this sense, our approach is predictive, and not just purely explanatory.

$^{11}$This also means that while safe-haven flows may influence London house price appreciation across the board, our estimates do not take this into account, meaning that they are likely to be underestimates.
If foreign buyers choose to purchase London properties in locations not highly populated by residents originating from their countries, we would estimate $\gamma_0^k$ to be negative.\footnote{High political risk in particular countries could also be accompanied by the imposition of capital controls or other restrictions on the ability of country residents to move capital overseas at such times, and a consequent reduction in capital flows to their preferred London areas.} In this sense, the estimated coefficients allow us to derive interesting insights about specific countries in addition to identifying safe-haven effects. $\gamma_0^k$ could also be estimated negative if declines in risk are associated with increases in prosperity, leading to an interpretation as a wealth effect. Throughout, the obvious null hypothesis is that $\gamma_0^k = 0$.

$f_{kw}^k$ may be correlated with other characteristics which attract foreign capital. An obvious possibility here is that $f_{kw}^k$ is high in relatively more desirable parts of London (such as Chelsea or Mayfair), if high-net worth foreigners direct safe-haven capital to these locations. We therefore add a ward-level indicator of desirability (in empirical estimation, we use net average ward-level income) $y_w$ as an additional conditioning factor in the interaction term. To make our specification parsimonious, we do not try to separately identify demand effects from $k$ countries on areas with high net average income, and simply interact $y_w$ with $\tau_t$, a weighted average risk measure across countries, with weights given by country-$k$ population shares in London.
2.2 Implications for Housing Transaction Volume

If we are indeed picking up the impacts of foreign demand on London real estate prices, we should also see effects of our foreign risk measures on London residential housing transaction volumes.

We therefore estimate:

$$\ln V_{w,t} = \theta_w + \zeta_t + \sum_{k \in K} \chi_0^k f_{w}^k z_{t-1} + \chi_1 y_w \tau_{t-1} + v_{w,t}. \quad (4)$$

In equation (4), we test whether $\chi_0^k$ and $\chi_1$ are statistically different from zero.

Many authors such as Stein (1995), Genesove and Mayer (2001), and Ortalo-Magne and Rady (2006), show that aggregate housing market transactions volume is positively associated with prices.\(^{13}\) If demand fluctuations generated by safe-haven effects drive housing transactions volumes, and ultimately prices we should be able to instrument the price-volume relationship detected by these authors using our identified safe-haven effects.

We therefore instrument for volume using $\ln \hat{V}_{w,t} = \sum_{k \in K} \chi_0^k f_{w}^k z_{t-1} + \chi_1 y_w \tau_{t-1}$, and estimate:

$$\ln P_{i,t} = \delta_t + \phi_w + \beta X_{i,t} + \theta^S \ln \hat{V}_{w,t} + \ln \hat{P}_{w,t-1} + \ln \hat{P}_{w,t-2} + u_{i,t}. \quad (5)$$

We expect $\theta^S$ to be positive and statistically significant.

\(^{13}\) Stein (1993) presents a rational model in which price increases and declines affect available down-payment amounts for mortgage-holders, and hence, the ability of homeowners to move. Genesove and Mayer (2001) rationalize the observation using loss aversion of prospective sellers.
2.3 The Persistence of Price Impacts from Foreign Demand

As in much research in real estate, we do not observe a high number of repeat sales of individual properties. Therefore, we use the dynamics of house prices in our hedonic regression specification (3) to understand the duration of the impact of safe-haven effects on London house prices. To do so, we simply constrain the effects to be the same across all countries $k$ in the interests of focusing on time-series variation rather than the cross-sectional variation in these effects.

We therefore define $s^k_t$ as the period $t$ “price spread” between the logarithm of average ward-level prices in the top and bottom quintiles of London wards, ranked by the strength of their link to country $k$, i.e., sorted by $f^k_w$.

$$s^k_t = \prod^k_{t}^\text{high} - \prod^k_{t}^\text{low}$$

We then estimate:

$$s^k_t = \mu^k + \delta_t + \sum_{q=1}^{Q} \rho_q s^k_{t-q} + \sum_{q=1}^{Q} \zeta_q z^k_{t-q} + u^k_t,$$  \hspace{1cm} (6)

$$z^k_t = \theta^k + \sum_{q=1}^{Q} \pi_q z^k_{t-q} + \varepsilon^k_t.$$  \hspace{1cm} (7)

The second of these equations is a simple autoregressive model for the country-specific risk factors $z^k_t$. This specification relies on the (to our mind plausible) restriction that there is no feedback from London house prices to foreign risk. We use equations (6) and (7) to generate an impulse response function of $s^k_t$ to a shock $\varepsilon^k_t$ to risk, to check whether the effects are permanent or temporary – as we might expect if foreign capital

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14The appendix describes exactly how specification (6) derives from the benchmark model described in equations (1) and (2).
generates temporary liquidity-driven dislocations from fundamentals in the London housing market.

2.4 Placebo Tests

Crises are relatively infrequent, which may lead to limits to our ability to detect safe-haven effects. Another concern is that specific features of the distributions of \( f_w^k \) and \( z_t^k \) may be responsible for any effects that we may detect, rather than the specific safe-haven channel in which we are interested.

One simple test that we conduct to verify our results is to distinguish between high- and low-risk countries. We document in the online appendix that a number of countries, such as Japan, the US, Germany, and Scandinavian countries have considerably lower average levels of political risk than the other 41 countries in our sample. A priori, we might expect to observe no safe-haven effects for these countries, so we test the null hypothesis \( \zeta_q = 0 \) in equation (6) separately for these countries.

In a more formal placebo test, we construct a synthetic variable \( z_{t-1}^k \) by drawing with replacement from the set \( K \setminus k \) for each of the 41 high-risk countries \( k \). We then include \( z_{t-1}^k \) in our panel regression (6) and estimate:

\[
  s_t^k = \mu^k + \delta_t + \rho_1 s_{t-1}^k + \rho_2 s_{t-2}^k + \zeta z_{t-1}^k + \beta z_{t-1}^k + u_t^k. \tag{8}
\]

We run 2,000 such regressions with random draws of \( z_{t-1}^k \).

In one version of the test, we restrict \( \zeta = 0 \) and plot the distribution of the estimated \( \beta \) coefficient to assess the power of our identification approach, and in the other, we simply include both \( z_{t-1}^k \) and \( z_{t-1}^k \) to check the distributions of \( \zeta \) and \( \beta \) when both are included.
2.5 Sources of Foreign Demand

Public discussion of the effects of foreign demand on London house prices has focused on high-value purchases from high-net-worth individuals seeking a safe-haven for their capital in light of risk increases in their home countries. We therefore check using a version of equation (3) whether price impacts are highest in high-value properties, separately estimating coefficients $\gamma_0^k$ and $\gamma_1$ conditional on levels of house prices.

It is also possible that relatively less wealthy individuals move capital towards London or other safe-haven investments in response to risk in their home countries. Given the high unit-value of residential real estate investments, it is more likely that such investments incorporate an implicit or explicit future consideration of future London-bound immigration.

To explore this channel, we estimate:

$$
\Delta f_{w,2011}^k = \alpha + \rho^k f_{w,2001}^k + \beta \Delta f_{w,2011}^{UK} + e_{w,2011}.
$$

(9)

We condition the change between 2011 and 2001 $\Delta f_{w,2011}^k = f_{w,2011}^k - f_{w,2001}^k$, in the share of people in ward $w$ originating from country $k$, on the starting level of this share in 2001 $f_{w,2001}^k$. The sign and significance of the coefficient $\rho^k$ indicates the degree to which immigrants from country $k$ move into wards with a pre-existing high share of people originating from their home country. $\Delta f_{w,2011}^{UK}$ is the change in the share of UK-origin residents in London wards, and eliminates variation in immigration shares that are mechanically generated by aggregate variation in the relative populations of foreign-vs-domestic origin residents.

(9) also provides further evidence on our identification strategy, if $\rho^k$ is estimated to be statistically different from zero – although it is worth noting again that our identification strategy is not conditional on immigration flows actually materializing. This possibility of the immigration channel also relates to the relative pace of financial
market transactions and those in real goods and services markets. When political or economic risks actually materialize, relatively fast moving capital flows towards London properties may be followed by relatively slow-moving subsequent increases in immigration. The online appendix investigates, to the extent permitted by the data, whether price increases in wards with higher shares of foreign-born people can forecast increases in future immigration into those wards.

2.6 Notes on Estimation

In our empirical applications, we construct $z^k_{t-1}$ as dummy variables which take the value of one if the underlying risk indicator lies in the top quintile of its country-specific time series distribution. We verify this approach in the online appendix using equation (6), which we estimate using decile dummies for $z^k_{t-1}$. As expected, the effects of $z^k_{t-1}$ are large and statistically significant only when $z^k_{t-1}$ is in its top few deciles. We also transform the foreign-born people shares $f^k_w$, and the net average income levels $y_w$ into dummy variables which take the value of one when these variables lie in the top 20% of their cross-ward distribution, and zero otherwise. The two exceptions to this rule are when we generate impulse response functions to foreign risk shocks using (6) and (7), and in the small final note on country-specific effects. In these two cases, we simply estimate effects which are linear in $z^k_{t-1}$ and $f^k_w$.\footnote{In earlier versions of the paper available online, we estimated linear specifications in which we did not transform these variables into quantiles, with very similar results. The current specification captures the impact of high movements in country-risk on locations in London with a high foreign share, permitting relatively sharper identification.}

This transformation implies that in a ward in which the share of people born in country $k$ is in the top quintile of the cross-ward distribution, when risk moves into the top quintile of its time series distribution, we predict an additional house price premium of $\gamma^k_{10}$ for transactions occurring in the subsequent month. In all tables and figures we
multiply the estimated coefficients by 100, so that they can be interpreted as percentage points.

We deal with persistence in ward-level prices by estimating and reporting robust standard errors in all tables and figures, clustered along two dimensions (see Cameron et al. (2011) and Thompson (2011)). The two dimensions are borough (a broader geographical unit of aggregation than wards – there are 32 boroughs in London encapsulating 624 wards) to capture potential cross-ward-time correlation, and time. In specification (6), we employ Driscoll-Kraay (1998) robust standard errors, which allow for cross-sectional dependence and serial correlation in the error terms, with a lag length of 12 months.\footnote{Driscoll-Kraay standard errors correct for cross-sectional dependence in the residuals at each point in time, as well as serial correlation in the average residuals over a pre-specified lag (in our case, 12 months).}

We conduct extensive robustness checks on our results, which we discuss briefly in the text, leaving details to the online appendix to the paper. One important check is that we interact the share of UK-origin residents in each ward with the level of the UK mortgage interest rate to pick up any potential source of cross-time cross-ward variation in house prices that may be correlated with mortgage credit availability or other domestic policy-related sources. This does not materially affect our inferences about foreign demand.

3 Data

We employ four datasets in our study, the first, from the UK Land Registry, the second, from the Nationwide Building Society, the third, census data from the Office for National Statistics in the UK, and the fourth, time-series indexes of economic and political risk measures.
3.1 Registry data

We obtain data on the complete set of house purchases in London from HM Land Registry. All purchasers of houses are required to report transactions to the Land Registry, and the data cover 2,445,057 transactions over the period from 1995 to 2013. This constitutes 13% of roughly 19 million residential property sales the Land Registry has lodged for England and Wales. Property characteristics reported with these data include the type of house (whether it is an apartment, semi-detached or terraced house, for example), the tenure status (whether the property is a leasehold or a freehold property) and an indicator of whether the property is newly built.

The Land Registry provides a postcode for each traded property. In the UK, postcodes allow for very granular geographical identification of properties, often covering just a segment of a street. This allows us to link each property to London electoral-ward-level information, as we describe below, allowing us to control for price-relevant characteristics of the location in which each property is located, and most importantly for our purposes, allowing us to connect property prices with electoral-ward-level foreign-origin share and income.

3.2 Loan-level data

In addition to the Land Registry data, we use proprietary loan-level mortgage data covering the period 1996 to 2012, obtained from the Nationwide Building Society. Nationwide is the second largest mortgage lender in the UK, with a market share of 14.8% of gross lending in 2012, and their house price index is considered one of the benchmark indexes characterizing the evolution of the UK housing market.

These data are collected following the completion of valuation reports on properties serving as mortgage collateral, and cover 154,137 observations of house purchases widely spread across London electoral wards over the sample period. This amounts to 11% of
the roughly 1.4 million transactions which are reported by Nationwide for the entire UK since 1996.

Despite its substantially smaller size than the Land Registry data, we use Nationwide data for two reasons. First, the ward fixed effects do not completely eliminate variation across properties arising from property-specific hedonic characteristics. The Nationwide data allow us to better control for hedonic characteristics of properties which may be important for house-price determination. For each individual property in the Nationwide London sample, in addition to the loan approval date and the purchase price, we know the geographical location at postcode level, the tenure status, the house type, the year of construction, the floor area, the number of bathrooms, bedrooms, and garages, and whether the borrower is a first-time home buyer. Second, the Nationwide data are most likely associated with domestic residents rather than foreign demand. If we find effects of foreign demand on transactions prices in these data as well, they are informative about the extent to which price impacts of foreign demand spillover into the domestic housing market.

Panel A of Table 1 shows the total number of transactions in each year, that are reported by the Land Registry as having occurred in London, as well as the number of mortgages issued by Nationwide which are associated with housing transactions in London. Interestingly, mortgage volumes and housing transactions both dip in 2008 and remain low relative to historical averages during the credit crunch, while London house prices are increasing over the same period, relative to prices in the remainder of the UK.

3.3 UK Office for National Statistics

We implement the majority of our analysis at the level of electoral wards. The 624 wards in London function as political sub-divisions, but also as administrative entities
within the city. The average number of people residing in each ward is roughly 13,000. We identify properties belonging to wards using property postcodes, using the UK Office for National Statistics’ (ONS) postcode directory. There are 1.8 million postcodes active in the UK, corresponding to 29 million postal addresses, an average of roughly 16 buildings per postcode. For each individual housing transaction from both Nationwide and Land Registry samples, we match postcodes to wards, and acquire demographic and geographical characteristics associated with the location of the house, also available from the ONS.

Data from the ONS on the composition of electoral wards is available in years 2001 and 2011, and we use data from 2001 throughout our study to identify ward-level characteristics. A key variable in our analysis is the share of each ward’s population that was born in foreign countries $k$, $f^k_w$. To select the $k \in K$ countries which we employ in our analysis, we do not impose any a priori criteria. Instead, we consider the entire set of countries represented in London, which we aggregate into nine world regions: Northern Europe and North America, Southern Europe, Eastern Europe, Russia, the Middle East, Africa, South Asia, the Asia-Pacific and South and Central America. The complete set of 55 countries which we employ and their regional groups are shown in Panel B of Table 1. As described earlier, we transform the foreign-born people shares $f^k_w$ and the net average income levels $y_w$ into dummy variables which take the value of one when these variables lie in the top 20% of their cross-ward distribution, and zero otherwise.

The online appendix shows histograms of a number of other ward-level variables. It does seem that wards more densely inhabited by people of foreign origin do have important differences with the remainder, but these static characteristics of wards are not allowed to affect our estimation as we include ward-fixed effects in our estimation. We do see that the fraction of owners with mortgages appears to be lower in wards with high shares of foreign-born people. While it is of course possible that foreigners’ access
to the UK mortgage system is less straightforward than access for UK-born London residents, we see this fact as consistent with the argument that foreign-born people draw upon sources of funds which lie outside the UK mortgage system, and at least partially, may come from overseas.\footnote{As mentioned earlier, we included a control for this variation in mortgage usage by interacting the total UK-origin share with the level of the UK mortgage interest rate, and found that our results were unaffected by the use of this control. Also, in order to account for the tendency of foreigners to cluster around specific areas and for the possible correlation with factors affecting UK demand, we consider a specification in which we interact time dummies with the total foreign-born origin share. In the online appendix, we show that our results are robust to the inclusion of these additional controls.}

We aggregate country-level variables into regions by weighting individual country data by the extent to which these countries are represented in London’s population. The ONS does not report country-specific population shares for some of the countries in our sample, so we simply use the overall share of people born in the Middle East for countries Qatar, UAE, Saudi Arabia, Syria, Lebanon, and Israel; the shares of people born in North Africa for Tunisia, Egypt, Algeria, and Libya, and the shares of people born in South America for Argentina, Brazil, Chile, and Mexico.

Panel A of Figure 1 shows the correlation between the shares of people born in different regions of the world (darker shaded areas of the matrix represent higher correlations), and confirms that London residents that were born abroad tend to cluster around specific areas, which are different from the ones preferred by residents born in the UK. At the same time, people that come from culturally and geographically proximate countries tend to live near one another. For example, London wards with high shares of residents originating in Southern Europe also have high shares of residents from Northern Europe, and relatively low shares of people originating in South Asia.
3.4 Measures of Foreign Risk

We measure economic and political risk in foreign countries using three separate variables. First, we obtain a list of indicators from the International Country Risk Guide (ICRG), which are monthly data on the political situation around the world. These indicators rate each country along 12 dimensions, each of which contributes to the total country rating according to the number of points indicated in parentheses: government stability (12 points), socioeconomic conditions (12 points), investment opportunities (12 points), the potential for internal and external conflicts (24 points), the level of corruption (6 points), the influence of military in politics (6 points), the prevalence of religious tensions (6 points), general conditions in terms of law and order (6 points), ethnic tensions (6 points), democratic accountability (6 points), and the quality of bureaucracy (4 points). For each country, we build a composite index by simply summing across these 12 risk categories (by ICRG construction, this adds up to 100), and we again use weights based on the shares of respective populations in London to build our time series for world regions.18

Glaeser et al. (2004) note that the ICRG indicators do not perfectly describe the permanent state of country-level political institutions, but they reflect actual changes that happen through time. For the purposes of our analysis, it is precisely this time series dimension in which we are most interested – capturing changes in the political situation in a given country, rather than solely differences across institutions. They also emphasize that the ICRG indicators are subjective assessments of risk, which is also relevant, as we are interested in drivers of foreign demand, which we believe will be correlated with prospective buyers’ sentiment about their local political environment. The ICRG data are also used by Erb, Harvey, and Viskanta (1996 a,b), who show

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18 As constructed, the index ranges from 0 to 100, with 0 indicating the highest possible risk. In our empirical implementation, we simply replace this with 100 minus the original values so that high levels of the index indicate high levels of risk and vice versa.
that ICRG ratings are correlated with expected stock and bond returns in a variety of countries.

Panel B of Figure 1 plots the evolution of the ICRG measures in the nine regions of the world that we study. The plot shows that these measures capture major political events and periods of widespread turmoil. Our sample period includes the Russian political crisis of 1999, the South-American crises around the turn of the millennium, the outbreak of the Arab Spring at the end of 2010, as well as the gradual increase in political uncertainty in Southern Europe in the wake of the 2009 sovereign debt crisis.

As described earlier, in our empirical implementation, we transform the ICRG and other risk variables into dummy variables which take the value of one if the respective country-level risk indicator lies in the top quintile (top 20%) of its time-series distribution. As described in our placebo test, in some parts of our analysis we distinguish between high- and low-risk countries, classified based on their average levels of risk. In the online appendix, we show the cross-sectional patterns of average political risk. We choose a threshold value equal to 20, to separate low from high-risk countries, meaning that the Czech Republic lies just above the threshold while Japan lies just below.

We also check robustness using two alternative measures of country risk. First, we use country-specific bond yield data, taking the spread of the country-level 10-year bond yield over the equivalent 10-year UK government bond yield.\footnote{Longstaff et al. (2011) find that there is less independent variation in country-specific bond yields than might otherwise be expected – showing that a global factor explains a large fraction of country-level bond yield variation. We therefore use both ICRG and bond measures to ensure that we are picking up country-specific variation in our empirical analysis.} These data are retrieved from Reuters via Datastream. Bond yield data are available for different periods for different countries in the sample; the date range across countries spans the period from 1995 to 2013. For some countries long-term bond data are unavailable; we describe the specific data restrictions for these variables in the online appendix.
Second, we use the economic policy uncertainty indexes of Baker, Bloom, and Davis (2013), who show that this measure is an important factor in determining the allocation of capital, and that it negatively affects investment, output, and employment dynamics in a variety of countries. We use the series at monthly frequency for Spain, Italy, China, and India, for the period between 2004 and 2013. We find that higher economic policy uncertainty is associated with increases in our political risk measure – in the online appendix, we report short-run co-movement between 12-month changes in the ICRG indexes of political risk and the economic policy uncertainty measure, country-by-country. Since the Baker, Bloom and Davis (2013) measure is based on news coverage and forecaster disagreement, it spikes during periods in which the initial turmoil occurs and tends to revert quickly. In contrast, the ICRG indexes tend to be more persistent.

4 Results

4.1 Preliminary Results: Time-Series Patterns and Total Foreign-Origin Share

The top panel of Figure 2, shows the widely documented emergence of a house price gap between London and the rest of the UK. The figure uses indexes reported by Nationwide, Halifax (now owned by Lloyds), the Land Registry, and the UK ONS, and plots the hedonic-adjusted percentage spread between the price of the average house in London relative to a population in 2001- weighted average house price in the remainder of the UK. All four series clearly show that the spread between London house prices and those in the remainder of the UK is very large on average. Moreover, this spread fluctuates substantially over time. There is a pronounced increase in this spread beginning in 1998, a period of heightened international political and economic uncertainty owing to the Asian and Russian financial crises. Following this period,
London prices appear to grow at roughly the same rate (or even slightly lower) as the remainder of the UK during the early part of the decade beginning in 2000. Finally, there is strong growth in the spread following the onset of the financial crisis beginning in 2008.

How much of this increase is attributable to external political and economic uncertainty? The bottom panel of Figure 2 shows the relationship between the London house price spread and a 12-month trailing moving average of two economic and political indicators (bond yield spreads over the UK, and the ICRG political risk index). These indicators are weighted by the extent to which people born in each foreign country are represented in London’s population, across all of the non-UK countries in our sample available in each time period. In the time-series, these foreign risk indicators appear closely related with the level of London house prices relative to those in the remainder of the UK.

Interpreting this time-series correlation as evidence of the impact of foreign demand on London house prices is problematic for a number of reasons. First, pure time-series relationships such as the one plotted in the figure are difficult to attribute to any single cause. Over the period under study, there were many dislocations in capital markets and a number of key determinants of house prices (most notably the availability of credit) were very likely highly correlated with one another. Sorting out their independent effects solely in the time series is difficult because of the limited degrees of freedom available in the time-dimension. Second, characterizing the entire external environment by aggregating all non-UK countries is unsatisfying. By doing so, we eliminate the possibility of separately identifying the effects by the country of origin of the safe-haven demand. A multivariate time-series analysis suffers from the very same small-time-series sample-size issue, i.e., simply putting all country-level measures on the right-hand-side in a time-series regression would make the problem no better. Third, better identification of foreign demand effects on London house prices might allow us
to explore the underlying economic motivations for this demand in a richer fashion.

A preliminary illustration of our results, which aggregates country-specific information into a single foreign-born share, is provided in Figure 3. The figures show maps of London, with electoral wards color-coded in a “heatmap” – red shades indicate greater levels of particular variables, and blue shades indicate lower levels. Panel A of the figure shows the average house price appreciation in London electoral wards between 2001 and 2006, a period of relatively low global uncertainty. Panel B of the figure shows the average house price appreciation in London electoral wards between 2007 and 2012, a period of elevated global risk. When juxtaposed, the figures clearly reveal that the London wards with the greatest house price appreciation shifted towards the centre and northwest of London. Panel C of the figure shows that this appears to be correlated with the high net-income areas of London as measured in the 2001 census, although the high-income areas in the South and South-West of London appear not to have experienced particularly high house price appreciation over the 2007 to 2012 period. Finally, Panel D of the figure shows the total share of foreign-origin people in London measured in the 2001 census. The pattern is visually striking – the house price appreciation between 2007 and 2012 appears to line up well with the share of foreign-origin people in London wards. Indeed, the simple cross-ward correlation between 2001-2006 price appreciation and 2001 foreign-origin shares is $-12\%$, while the cross-ward correlation between 2007-2012 price appreciation and 2001 foreign-origin shares increases to $+38\%$.

The online appendix shows that the use of the single foreign-born origin share and a measure of global uncertainty yields strong and statistically significant results in our specifications. However, this masks the important and economically interesting cross-country variation in safe-haven effects which is possible to estimate using our method. We first turn to describing the effect of hedonics, and then turn to these cross-country results.
4.2 Hedonics

Table 2 shows the coefficients $\beta$ from a simple hedonic regression specification in which we do not include our interaction terms to identify safe-haven effects, but do include time and ward fixed effects. Panel A of the table shows the estimated coefficients on the hedonic characteristics that are common to both Land Registry and Nationwide datasets, with the left (right) part of the panel showing the estimated effects in the former (latter) dataset. While the Land Registry data is somewhat sparse on hedonic characteristics, the estimated signs for the common characteristics are strongly consistent across the datasets. Panel B of the table shows the set of hedonic characteristics that are present only in the Nationwide data.

The estimated contributions to marginal hedonic utility of bedrooms, bathrooms, parking spaces, and floor area (the omitted category is the smallest possible unit in all cases) appear very reasonable. Older properties, all else equal, are valued higher than newer ones (with the exception of properties built in the 2000s which are not new builds), and detached houses are worth more than any other category of houses, once floor-space area is controlled for.\footnote{Our empirical specifications focus on interactions between ward-level characteristics and foreign risk; we do not consider how safe-haven demand may be associated with particular property-level hedonic characteristics. It is of course straightforward to extend our approach to account for these possibilities.}

Finally, Panel C of Table 2 shows the estimated time fixed effects, averaged across all months in each year. The time fixed effects show steady appreciation in London house prices, with the only recorded declines in 2008 and 2009, but strong recovery in 2010 and 2011. We next turn to presenting the main point of our analysis, which is the empirical identification of foreign demand effects from our specifications.
4.3 The Effects of Foreign Demand

Figure 4 shows the coefficients $\gamma_k^0$ estimated from equation (3) for each of the nine world regions, as well as the coefficient $\gamma_1$, which captures safe-haven effects in wards with high net average income.

The figure shows that risk in Southern EU, the Asia-Pacific region, the Middle East, and South- and Central America are associated with subsequent price premia in London wards with high shares of people born in these regions. Controlling for this tendency, there are also elevated prices in high-income wards in London following periods of high overall foreign risk.

In terms of magnitudes, the effect appears large, even though it is likely an underestimate of the total effect, since these coefficients do not contain the common impacts of foreign risk on all London wards. Averaging across world regions, a shift to a high-risk regime predicts that house prices in the subsequent month in London wards with the highest shares of foreign-origin people will be elevated by 1.70 percentage points. Controlling for this foreign-born share effect, the effect of foreign political risk amounts to 2.86 percentage points on wards with high average income. These numbers come from the Land Registry estimates; comparable magnitudes are evident from the Nationwide data.

Our identification of safe haven effects uses both the time series dynamics of foreign risk and the cross-sectional variation of house prices across the 624 London wards. In the online appendix, we check whether these effects are localized at the ward level, or reflect broader effects on larger areas of the city, estimating $\gamma_k^0$ and $\gamma_1$ in a specification which includes borough-cross-year fixed effects (there are 32 boroughs within London). The specification confirms that broader geographical effects do matter for our estimation of $\gamma_k^0$, but even controlling for these broader effects, within-borough, cross-ward variation in the foreign-origin share is useful for identifying safe-haven effects from Southern
Europe and the Asia-Pacific region. Interestingly, the $\gamma_1$ coefficients remain statistically significant and even increase in magnitude - this suggests that within-borough cross-ward income variation is important in identifying safe-haven effects.

### 4.4 Evidence on Transaction Volumes

Table 3 shows results from estimating equations (4)-(5). Panel A of the table shows that our foreign risk measures are associated with an increase in transaction volumes. For example, high risk in Southern Europe is associated with 8.47 percent higher transaction volumes in wards with the highest share of people born in Southern Europe. The bottom of panel A shows that the hypothesis that $\chi_0^k$ and $\chi_1$ are jointly zero is strongly rejected, with a close-to-zero p-value.

Panel B of the table first shows that there is a strong positive correlation between volume and prices in the London housing market, and that this correlation survives the introduction of time and ward fixed-effects. In the last column of Panel B, we report results from (5), and find that volume instrumented with our measures of foreign-demand has a significantly positive association with prices. We view this as encouraging evidence that our method is successful at identifying foreign demand effects that impact volume, and ultimately prices in the London residential real estate market.

### 4.5 The Persistence of Foreign Demand Effects

As described in the methodology section, we complement our hedonic regressions by computing price “spreads” $s_t^k$ in each period $t$, and for each country $k$, between the top and bottom quintiles of wards sorted by the share of residents originating in country $k$.

To begin with, we simply estimate (6) using two lags of $s_t^k$ and a single lag of $z_{t-1}^k$, on the sample of 41 high-risk countries indicated by their higher average levels of risk. Table 4 summarizes the findings from this regression. The first column presents
estimated coefficients from a simple specification in which we condition the panel of spreads on lagged values of top-quintile-dummies for ICRG indexes in each country. The results using the Registry dataset show that high levels of risk are associated with prices that are 0.17 percentage points higher in London wards with high foreign-origin shares. When we eliminate common shocks using monthly time dummies, the second column shows that the estimated effects of foreign demand reduce to 0.11 percentage points, but continue to be statistically significant.

The third column shows results when using 10-year bond yield spreads versus the UK as measures of foreign risk in place of the ICRG indexes. Following periods when foreign yield spreads over the UK are elevated, prices are 0.25 percentage points higher in areas which are associated with the respective country of origin. All of the above patterns are also visible in the fourth column, in which we replace the ICRG risk measures with the Baker, Bloom and Davis (2013) economic policy uncertainty indexes. In both specifications, we do not introduce time fixed effects because of the far more limited cross-section of countries (13 for bond yield spreads, and 5 for EPU) available on account of the use of these alternative measures.

In the online appendix, we show that the estimated effects of foreign demand are particularly important during periods in which foreign risk is unusually high. We use the country-specific time-series distributions of the ICRG indexes to generate decile dummy variables. The results show that the top three deciles are the main driver of the effects. This suggests that the relationships that we identify are primarily driven by precautionary savings and “flight-to-safety” motives during periods of extreme political and economic risk.

In Figure 5, we estimate the evolution of the spread in response to a one standard deviation shock to foreign political risk, using equations (6) and (7).\textsuperscript{21} The estimation

\textsuperscript{21}We choose the lag length of 20 months, to balance precision in estimation against the need to capture the substantial delays in housing transactions (information acquisition, legal process, international
results using the Registry dataset show that the predicted spread in house prices begins at 0.2 percent, and reaches a peak at 0.8 percent. 19 months after the shock, the effect becomes statistically indistinguishable from zero. In the Loans dataset, the effects are also estimated to be predominantly positive, although estimated far precisely because of the substantially smaller sample size.

The estimated impulse response functions suggest that the effects of foreign demand on the London housing market, while long-lived, are temporary. Properties in wards with higher foreign-origin shares trade at higher prices following increases in foreign political risk, but after 19 months, the predicted price gap arising from this source is statistically indistinguishable from zero. The relatively long lag before the effects dissipate are consistent with the high persistence of political risk, and substantial house price inertia, a likely consequence of the difficulty of arbitraging price discrepancies in this market.
4.6 Placebo Tests

As discussed earlier, we identify 14 low-risk countries (Japan, Belgium, the USA, Portugal, Germany, Singapore, Australia, Canada, Austria, Denmark, New Zealand, the Netherlands, Sweden and Finland) with far lower average levels of ICRG-measured risk. Our first test is to estimate (6) for this low-risk sample. We report the results in the last column of Table 4, which show that variation in risk for low-risk countries has effects on London house prices which are statistically indistinguishable from zero in the Registry sample, and negative and statistically significant in the Loans sample. While our results are consistent with a “flight-to-safety” for the high-risk countries, these results suggest that there are either no demand effects conditional on variation in risk in low-risk countries, or that there may be a milder wealth effect in these countries which affects London house prices.

Figure 6 reports the results of the placebo tests described in equation (8). In Panel A, we constrain the coefficient $\zeta$ to equal zero, and plot the estimated distribution of the coefficient $\beta$ in equation (8) across a set of N=2,000 draws with replacement from the set of countries $K \setminus k$. The point estimate on $\zeta$ in Table 4 is 0.11 with a t-statistic of 2.75. Panel A of the figure shows that this value lies in the extreme right tail of the distribution of the coefficient under random draws with replacement, providing additional reinforcement that our result is not merely a statistical artefact.

Panel B of the figure shows what happens when we estimate both $\zeta$ and $\beta$ simultaneously. The estimated coefficient $\zeta$, plotted in the left panel varies between 0.077 and 0.148 percentage points, with a 5% lower percentile equal to 0.099 percentage points. In contrast, the coefficient $\beta$ on randomly assigned risk, plotted in the right panel varies between -0.207 and 0.152 with far greater dispersion.
4.7 Sources of Demand Effects

In Figure 7, we check whether the estimated impacts of foreign demand differ with the price of the property. This is to investigate whether the safe-haven effects are driven by the wealthiest purchasers from foreign nations seeking to preserve capital in response to elevated risk in their countries of residence. The figure shows that the impacts of foreign risk on London house prices are generally increasing with the price of the house across all countries. Of course, the reduction in sample size in the very highest price category means that some of these effects are not very precisely estimated, but the economic magnitudes of the effects are substantial. In some cases, the effects cause the interactions to change sign in a direction which is more consistent with our initial hypothesis – for example, controlling for the impact of global risk on the prices of houses in high net average income wards, we now find that for the highest price category of houses, there is also price appreciation of houses in London wards with a high share of Russian-origin residents. Overall, it does appear that “safe-haven” motivations drive some portion of foreign demand for London housing.

Figure 8 shows estimates of equation (9) for the countries for which we are able to track changes in ward-level shares of foreign-born people between the two (2001 and 2011) waves of the census. The figure shows that for a large set of countries there is a strong, statistically significant correlation between these changes and the initial levels of ward-level foreign-born shares in 2001. This lends credibility to our identification approach despite the reduced set of countries for which we are able to estimate this. It is the case, however, that for a smaller set of countries which includes Jamaica, Kenya, Bangladesh, and Zimbabwe, that we see the opposite pattern, namely that immigration...

22There are also frequent statements in the popular press that foreign-origin safe-haven demand is most prevalent at the very top of the housing market. See, for example, “Foreign buyers behind half of £2m+ home sales in London,” The Guardian, 6 May 2013, and “Half of central London’s £1m-plus homes go to non-UK buyers,” The Telegraph, 8 October 2013.
to wards with high shares of residents originating from these countries is significantly lower than in other areas of London.

These patterns in immigration shares suggest that countries may differ along this and other dimensions, which we may be able to derive interesting insights from. We therefore analyze the cross-country patterns of foreign demand effects from individual countries, now allowing the income interaction on $y_w^k$ to also be estimated independently for each country. Rather than considering a few episodes of crisis country-by-country, we now allow for a linear dependence between foreign risk and ward-level house prices.

Figure 9 plots these income and foreign-origin share coefficients ($\gamma_0^k$ and $\gamma_1^k$) against one another. To aid understanding of this plot, consider a hypothetical country $k$ for which $\gamma_0^k$ and $\gamma_1^k$ are both estimated to be high. From such a country, controlling for the overall desirability of London wards, capital is more likely to be directed towards wards with higher shares of residents originating from the country. Moreover, controlling for the foreign-origin share, purchasers from country $k$ are more likely to direct investment towards more desirable regions of London. As discussed earlier, there are several possible interpretations of negative coefficients. Our preferred interpretation is that foreign demand from such countries is directed towards London wards with relatively low average income levels, or low foreign-born shares.

Figure 9 shows a negative relationship between the two estimated coefficients, which we believe captures economically interesting cross-country variation in the “home bias abroad” of demand emanating from particular countries. The world seems to be broadly divided into two groups. Countries in the Middle East, parts of Africa, and Russia appear to prefer premium areas of London, and holding this tendency constant, appear to disfavor areas of London in which there are pre-existing denizens originating from their countries.

On the other end of the spectrum countries in Southern and Eastern Europe and parts of South Asia appear to direct capital towards areas of London populated by
their compatriots, and are less likely to direct capital towards premium residential real estate.

To illustrate this point with a specific example, following heightened risk in Pakistan, prices in both Wimbledon Park (a relatively high-income ward around the famous tennis club) and Southall (a relatively low-income ward which is part of “Little South Asia” in London) would be expected to experience increases as they both have relatively high shares of Pakistan-born individuals. However, our empirical estimates suggest that there would be relatively higher increases in Southall than in Wimbledon Park through the income channel.

Despite the significant imprecision of these estimates, and the lack of controls for global variation in risk simultaneously driving country-risk in these country-specific plots, it is tempting to conjecture that these patterns might line up well with the skewness of the wealth distribution across countries, as well as the relative frictions associated with moving capital or labour across borders. Southern European country-capital flows to London real estate may be driven by a wider cross-section of the domestic population in countries in this world region, given both the relatively less skewed wealth distribution in these countries, and low frictions associated with relocating within the European Union. Russian, Central African, and Middle Eastern capital flows to London, on the other hand, may be more closely associated with ultra-wealthy purchasers and super-premium areas of London.

5 Conclusions

In this paper, we propose a novel method to identify the impact of demand on asset prices, which relies on the fact that investors may have different “preferred habitats” for their capital within a broad asset class. Using this method and large micro-datasets of housing transactions in London, we find economically large, statistically significant, and
robust effects of foreign risk on house prices in locations in which the share of foreign-origin London residents are high, and controlling for this effect, on London areas which are wealthier and more desirable.

Our empirical results provide a careful analysis of whether foreign capital flows have affected global real estate prices and price volatility, especially in global cities such as London and New York. This is a phenomenon that has been widely commented upon in the press, and been the subject of numerous policy debates, but has not been supported by rigorous empirical work thus far. We view our results as a more general contribution to the analysis of macro-variation in prices, by using an identification strategy that is grounded in agents’ microeconomic motivations.
APPENDIX

Hedonic House Price Specification

Equation (1) implies that the average level of transaction prices across all properties in a ward in a time period can be written as:

$$\ln P_{w,t} = \alpha + \Pi_{w,t} + \beta X_{w,t} + u_{w,t}. \quad (10)$$

where $X_{w,t}$ are the average hedonic characteristics of properties in ward $w$ in period $t$.

Our identification of foreign demand effects relies on properties in London areas with higher shares of foreign-origin residents trading at a price premium to those with weaker international ties, and a similar premium between areas which are relatively more or less affluent, following elevated risk in foreign countries. To account for persistence in this price gap across London wards, we propose that ward-level prices follow an $AR(2)$, as in equation (2). This implies:

$$\Pi_{w,t} = \rho_1 (\ln P_{w,t-1} - \alpha - \beta X_{w,t-1} - u_{w,t-1}) + \rho_2 (\ln P_{w,t-2} - \alpha - \beta X_{w,t-2} (11)$$
$$-u_{w,t-2}) + \delta_t + \phi_w + \sum_{k \in K} \gamma^k_{0 \delta k} x_k + \gamma^k_{y w} y_{w t-1} + \gamma_{1 y w} z_{t-1}.$$

Combining equations (1) and (11), our benchmark hedonic house price model can be expressed as:

$$\ln P_{i,t} = \mu + \delta_t + \phi_w + \sum_{k \in K} \gamma^k_{0 \delta k} x_k + \gamma^k_{y w} y_{w t-1} + \beta X_{i,t} + \varrho_1 \ln P_{w,t-1} (12)$$
$$+ \varrho_2 \ln P_{w,t-2} + b_1 X_{w,t-1} + b_2 X_{w,t-2} + \nu_{i,t}.$$

This specification is equivalent to one in which we express price changes as a linear function of lagged price changes and the lagged price level.

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Since we observe little time variation in the observed hedonic characteristics of properties, we restrict \( b_1 = b_2 = 0 \) in our benchmark estimated specification. In this case, any cross-ward variation in hedonics is absorbed in the ward-level fixed effects. In the online appendix, we also report results from an unconstrained version of equation (12), which produces very similar results.

The autoregressive nature of ward-level prices implies that \( \nu_{i,t} = \rho_1 \Pi_{u,t-1} + \rho_2 \Pi_{u,t-2} + u_{i,t} \), leading to both serial- and ward-level correlation of error terms. We deal with this issue by reporting robust standard errors in all tables and figures, which are clustered along two dimensions, namely, borough and time.

**Spreads Specification**

Let \( \Pi_t^{\text{high},k} \) be the hedonic-adjusted average price level in period \( t \) in the top 20% of wards with the highest shares of people born in country \( k \). From equation (2), we obtain:

\[
\Pi_t^{\text{high},k} = \rho_1 \Pi_{t-1}^{\text{high},k} + \rho_2 \Pi_{t-2}^{\text{high},k} + \delta_t + \bar{\phi}^{\text{high},k} + \gamma_0 \bar{f}^{\text{high},k} z_{t-1} + \sum_{k' \in K, k' \neq k} \gamma_{0} \bar{f}^{k'|\text{high},k,k'} z_{t-1} + \gamma_1 \bar{y}^{\text{high},k} z_{t-1},
\]

where \( \bar{\phi}^{\text{high},k} \) is the ward fixed effect, \( \bar{f}^{\text{high},k} \) is the average share of people born in country \( k \), \( \bar{f}^{k'|\text{high},k} \) is the average share of people born in country \( k' \) and \( \bar{y}^{\text{high},k} \) is the average income level across these wards. Analogously, for the hedonic-adjusted price level in the wards with the lowest shares of people born in country \( k \), we obtain:

\[
\Pi_t^{\text{low},k} = \rho_1 \Pi_{t-1}^{\text{low},k} + \rho_2 \Pi_{t-2}^{\text{low},k} + \delta_t + \bar{\phi}^{\text{low},k} + \gamma_0 \bar{f}^{\text{high},k} z_{t-1} + \sum_{k' \in K, k' \neq k} \gamma_{0} \bar{f}^{k'|\text{low},k,k'} z_{t-1} + \gamma_1 \bar{y}^{\text{low},k} z_{t-1}.
\]
The spread $s^k_t$ is the difference between $\Pi^\text{high},k_t$ and $\Pi^\text{low},k_t$:

$$s^k_t = \rho_1 s^k_{t-1} + \rho_2 s^k_{t-2} + (\bar{\phi}^\text{high},k - \bar{\phi}^\text{low},k) + \gamma_0 (\bar{f}^\text{high},k - \bar{f}^\text{low},k) z^k_{t-1} + \sum_{k' \in K, k' \neq k} \gamma_0 \left( \bar{f}^{k'|\text{high},k} - \bar{f}^{k'|\text{low},k} \right) z^k_{t-1} + \gamma_1 \left( \bar{y}^\text{high},k - \bar{y}^\text{low},k \right) z^k_{t-1}.$$

Thus, changes in foreign risk affect house prices in wards with strong links to country $k$ in two ways. First, foreign demand from country $k$ has a direct effect on prices in these areas. Second, demand from other countries can also lead to price pressure to the extent that $k$-origin people live in areas which are preferred by residents from those other countries, or areas with high average income levels.

In the online appendix, we report estimation results when explicitly considering the role of these indirect safe-haven effects. We find that estimated $\gamma_0$ is very similar in this specification to the spreads specification with time fixed effects. The key to this is that the last two terms in the above equation do not vary greatly across countries – the first principal component of the set of interaction terms $\left( \bar{f}^{k'|\text{high},k} - \bar{f}^{k'|\text{low},k} \right) z^k_{t-1}$ explains 85% of the total variation across our full sample of countries.

We also find that it is difficult to separately identify direct and indirect effects of foreign demand whenever populations from similar countries are too geographically proximate. In our benchmark hedonic framework, we therefore group similar countries into nine distinct regions of the world.

As far as the net average income channel is concerned, we find that $\gamma_1$ is large, but imprecisely estimated. This is because of the low variation in income levels across wards populated by different foreign-origin residents. This fact motivates our choice to constrain $\gamma_1$ to be the same across countries in our benchmark specification.
6 References


Table 1
Summary statistics

In Panel A, we report the number of observations per year, for both the transaction-level data from the Land Registry and the loan-level data from the Nationwide Building Society. A hyphen indicates that our data samples do not cover the respective years. In Panel B, we list the countries which are part of our analysis concerning the effects of external factors on the London housing market. We group the countries in nine regions and weight them according to their average population shares in London.

Panel A
Number of transactions

<table>
<thead>
<tr>
<th>Year</th>
<th>Registry dataset</th>
<th>Loans dataset</th>
<th>Year</th>
<th>Registry dataset</th>
<th>Loans dataset</th>
<th>Year</th>
<th>Registry dataset</th>
<th>Loans dataset</th>
</tr>
</thead>
<tbody>
<tr>
<td>1995</td>
<td>84,525</td>
<td>-</td>
<td>2001</td>
<td>162,705</td>
<td>7,522</td>
<td>2007</td>
<td>166,417</td>
<td>9,143</td>
</tr>
<tr>
<td>1998</td>
<td>133,784</td>
<td>12,463</td>
<td>2004</td>
<td>165,471</td>
<td>7,565</td>
<td>2010</td>
<td>91,877</td>
<td>5,068</td>
</tr>
<tr>
<td>2000</td>
<td>147,865</td>
<td>8,958</td>
<td>2006</td>
<td>172,360</td>
<td>10,712</td>
<td>2012</td>
<td>92,151</td>
<td>9,892</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
<td>2013</td>
<td>108,383</td>
<td>-</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td><strong>Total</strong></td>
<td><strong>154,137</strong></td>
<td></td>
<td><strong>2,445,057</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

Panel B
List of countries

**Northern Europe and North America**
- Austria
- Belgium
- Denmark
- Finland
- Netherlands
- Germany
- Sweden
- USA
- Canada

**Southern Europe**
- Italy
- Spain
- Portugal
- France
- Greece

**Eastern Europe**
- Poland
- Romania
- Czech Rep.

**Russia**
- China
- Hong Kong
- New Zealand
- Malaysia
- Singapore
- Australia
- Japan

**Asia-Pacific**
- India
- Bangladesh
- Pakistan
- Sri Lanka

**Africa**
- Nigeria
- Sierra Leone
- Congo
- Kenya
- South Africa
- Somalia
- Zimbabwe

**Middle East**
- Cyprus
- Turkey
- Iran
- Iraq
- Tunisia
- Libya
- Algeria
- Egypt
- Qatar
- S. Arabia
- UAE
- Lebanon
- Syria
- Israel

**South- and Central America**
- Jamaica
- Brazil
- Argentina
- Chile
- Mexico
Table 2

Estimated coefficients in the hedonic regression framework

The table reports estimated coefficients from the following hedonic regression:

\[ \ln P_{i,t} = \phi_w + \delta_t + \beta X_{i,t} + u_{i,t}, \]

where \( X_{i,t} \) are property-level characteristics. The reported time fixed effects are yearly averages of the estimated monthly coefficients. A hyphen indicates that our data samples do not cover the respective years. The standard errors are double-clustered at the borough and time level. *, **, *** denote statistical significance at the 10%, 5%, and 1% level, respectively.

Panel A

<table>
<thead>
<tr>
<th>Registry dataset</th>
<th>Loans dataset</th>
</tr>
</thead>
<tbody>
<tr>
<td>Semi detached house</td>
<td>-0.426***</td>
</tr>
<tr>
<td>Terraced house</td>
<td>-0.580***</td>
</tr>
<tr>
<td>Flat</td>
<td>-0.744***</td>
</tr>
<tr>
<td>Semi detached bungalow</td>
<td>-0.065***</td>
</tr>
<tr>
<td>Purpose built flat</td>
<td>-0.274***</td>
</tr>
<tr>
<td>Flat conversion</td>
<td>-0.208***</td>
</tr>
<tr>
<td>New property</td>
<td>0.232***</td>
</tr>
<tr>
<td>Leasehold indicator</td>
<td>-0.257***</td>
</tr>
</tbody>
</table>

Panel B

<table>
<thead>
<tr>
<th>Loans dataset</th>
</tr>
</thead>
<tbody>
<tr>
<td>First purchase</td>
</tr>
<tr>
<td>Two bedrooms</td>
</tr>
<tr>
<td>Three bedrooms</td>
</tr>
<tr>
<td>Four or five bedrooms</td>
</tr>
<tr>
<td>More than five bedrooms</td>
</tr>
<tr>
<td>Two bathrooms</td>
</tr>
<tr>
<td>Three bathrooms</td>
</tr>
<tr>
<td>More bathrooms</td>
</tr>
<tr>
<td>Parking space</td>
</tr>
<tr>
<td>Single garage</td>
</tr>
<tr>
<td>Double garage</td>
</tr>
</tbody>
</table>

<table>
<thead>
<tr>
<th>Construction date</th>
<th>Floor area</th>
</tr>
</thead>
<tbody>
<tr>
<td>1900 to 1920</td>
<td>50 to 70 m²</td>
</tr>
<tr>
<td>1920 to 1940</td>
<td>70 to 90 m²</td>
</tr>
<tr>
<td>1940 to 1960</td>
<td>90 to 110 m²</td>
</tr>
<tr>
<td>1960 to 1980</td>
<td>110 to 130 m²</td>
</tr>
<tr>
<td>1980 to 2000</td>
<td>130 to 150 m²</td>
</tr>
<tr>
<td>after 2000</td>
<td>150 to 170 m²</td>
</tr>
<tr>
<td></td>
<td>above 170 m²</td>
</tr>
</tbody>
</table>
Table 2
Estimated coefficients in the hedonic regression framework (continued)

Panel C

<table>
<thead>
<tr>
<th>Estimated time fixed effects</th>
<th>Registry dataset</th>
<th>Loans dataset</th>
</tr>
</thead>
<tbody>
<tr>
<td>1996</td>
<td>0.052***</td>
<td>-</td>
</tr>
<tr>
<td>1997</td>
<td>0.184***</td>
<td>0.217***</td>
</tr>
<tr>
<td>1998</td>
<td>0.311***</td>
<td>0.354***</td>
</tr>
<tr>
<td>1999</td>
<td>0.436***</td>
<td>0.511***</td>
</tr>
<tr>
<td>2000</td>
<td>0.635***</td>
<td>0.699***</td>
</tr>
<tr>
<td>2001</td>
<td>0.742***</td>
<td>0.810***</td>
</tr>
<tr>
<td>2002</td>
<td>0.902***</td>
<td>0.982***</td>
</tr>
<tr>
<td>2003</td>
<td>1.027***</td>
<td>1.080***</td>
</tr>
<tr>
<td>2004</td>
<td>1.108***</td>
<td>1.149***</td>
</tr>
<tr>
<td>2005</td>
<td>1.139***</td>
<td>1.171***</td>
</tr>
<tr>
<td>2006</td>
<td>1.203***</td>
<td>1.255***</td>
</tr>
<tr>
<td>2007</td>
<td>1.342***</td>
<td>1.391***</td>
</tr>
<tr>
<td>2008</td>
<td>1.327***</td>
<td>1.337***</td>
</tr>
<tr>
<td>2009</td>
<td>1.233***</td>
<td>1.255***</td>
</tr>
<tr>
<td>2010</td>
<td>1.340***</td>
<td>1.350***</td>
</tr>
<tr>
<td>2011</td>
<td>1.358***</td>
<td>1.386***</td>
</tr>
<tr>
<td>2012</td>
<td>1.409***</td>
<td>1.418***</td>
</tr>
<tr>
<td>2013</td>
<td>1.490***</td>
<td>-</td>
</tr>
</tbody>
</table>
Table 3
Foreign risk and transaction volumes

In Panel A we report the coefficients $\chi_k^0$ and $\chi_1$ from the following regression:

$$\ln V_{w,t} = \theta_w + \xi_t + \sum_{k \in K} \chi_k^0 f_k^w z_t^{k-1} + \chi_1 y_w z_t - 1 + v_{w,t},$$

where $V_{w,t}$ is the number of transactions in ward $w$ in period $t$. In Panel B we report the estimated coefficients $\theta$ and $\theta^S$ from the following specifications:

$$\ln P_{i,t} = \delta_t + \phi_w + \beta X_i,t + \theta \ln V_{w,t} + \ln P_{w,t-1} + \ln P_{w,t-2} + u_{i,t},$$

$$\ln P_{i,t} = \delta_t + \phi_w + \beta X_i,t + \theta^S \ln V_{w,t} + \ln P_{w,t-1} + \ln P_{w,t-2} + u_{i,t},$$

where $z_t^k$ is a top-quintile indicator of the ICRG index of political risk, corresponding to the nine world regions listed in Table 1. The estimated coefficients are multiplied by 100, for easier interpretation as percentage points. The estimation is carried out in the Registry sample of housing transactions. In parentheses, we report standard errors, double-clustered at the borough and time level. *, **, *** denote statistical significance at the 10%, 5%, and 1% level respectively.

Panel A
Explaining the evolution of transaction volumes

<table>
<thead>
<tr>
<th>Region</th>
<th>Coefficient</th>
<th>Standard Error</th>
</tr>
</thead>
<tbody>
<tr>
<td>Southern Europe</td>
<td>8.47***</td>
<td>(2.78)</td>
</tr>
<tr>
<td>Russia</td>
<td>6.97***</td>
<td>(2.04)</td>
</tr>
<tr>
<td>Asia-Pacific</td>
<td>9.39***</td>
<td>(2.26)</td>
</tr>
<tr>
<td>Eastern Europe</td>
<td>-8.80***</td>
<td>(2.39)</td>
</tr>
<tr>
<td>Middle East</td>
<td>-1.74</td>
<td>(3.94)</td>
</tr>
<tr>
<td>South Asia</td>
<td>-2.79</td>
<td>(4.23)</td>
</tr>
<tr>
<td>Africa</td>
<td>8.98**</td>
<td>(3.58)</td>
</tr>
<tr>
<td>Sth. and C. America</td>
<td>6.16**</td>
<td>(2.82)</td>
</tr>
<tr>
<td>North EU and America</td>
<td>8.44**</td>
<td>(3.64)</td>
</tr>
<tr>
<td>Net average income</td>
<td>5.38*</td>
<td>(3.23)</td>
</tr>
</tbody>
</table>

$F$-test for joint significance of safe-haven effects: $F(10, 7295) = 8.31$.

$P$-Value: 0.00

Panel B
Hedonic house price regression

<table>
<thead>
<tr>
<th></th>
<th>Benchmark</th>
<th>IV spec.</th>
</tr>
</thead>
<tbody>
<tr>
<td>Transaction volume</td>
<td>2.98***</td>
<td>26.23***</td>
</tr>
<tr>
<td></td>
<td>(0.42)</td>
<td>(4.00)</td>
</tr>
</tbody>
</table>
Table 4
Cross-country panel analysis

The table reports the estimated coefficients $\zeta$ from panel regressions where the dependent variable is the price spread $s^k_t$ in period $t$ between the top and bottom 20% of wards with respect to the share of people born in country $k$. In each of the three sets of columns, we report estimates for the ICRG indexes of political risk, the 10-year bond yield spreads versus the UK and the Economic Policy Uncertainty indexes of Baker, Bloom and Davis (2013), respectively. The specification is given by:

$$s^k_t = \mu^k + \delta_t + \rho_1 s^k_{t-1} + \rho_2 s^k_{t-2} + \zeta z^k_{t-1} + u^k_t,$$

where $z^k_t$ is an indicator variable which takes a value of one if the respective risk measure from country $k$ is in the top 20% of its time series distribution. In parentheses, we report Driscoll-Kraay (1998) standard errors with a lag length of 12 months.

<table>
<thead>
<tr>
<th>ICRG index vs. the UK uncertainty index</th>
<th>(1)</th>
<th>(2)</th>
<th>(3)</th>
<th>(4)</th>
<th>(5)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Registry dataset</td>
<td>0.17***</td>
<td>0.11**</td>
<td>0.25**</td>
<td>0.42**</td>
<td>-0.05</td>
</tr>
<tr>
<td></td>
<td>(0.06)</td>
<td>(0.04)</td>
<td>(0.11)</td>
<td>(0.20)</td>
<td>(0.09)</td>
</tr>
<tr>
<td>Loans dataset</td>
<td>0.27</td>
<td>0.02</td>
<td>0.70**</td>
<td>1.35***</td>
<td>-0.57***</td>
</tr>
<tr>
<td></td>
<td>(0.23)</td>
<td>(0.13)</td>
<td>(0.33)</td>
<td>(0.50)</td>
<td>(0.20)</td>
</tr>
<tr>
<td>Country fixed effects</td>
<td>Yes</td>
<td>Yes</td>
<td>Yes</td>
<td>Yes</td>
<td>Yes</td>
</tr>
<tr>
<td>Time fixed effects</td>
<td>No</td>
<td>Yes</td>
<td>No</td>
<td>No</td>
<td>Yes</td>
</tr>
</tbody>
</table>
Figure 1

Shares of foreign-born people and foreign political risk

In Panel A, we report the pairwise correlation coefficients between the shares of people born in foreign countries, across the 624 London wards. Blocks in darker shading indicate a higher tendency of the population from the respective countries to cluster around similar areas of the city. We calculate the fractions of foreign-born people relative to the total ward population by using 2001 census data, indicating the country of birth. In Panel B, we report the time series evolution of political risk in the nine world regions. The regional aggregates are calculated by using weights given by respective population shares in London. The list summarizing our country coverage is given in Table 1 and the data sources are indicated in the online appendix.
Panel A shows the evolution of London house prices relative to the UK. We collect the alternative series reported by four different UK institutions and aggregate the regional indexes by using 2001 census population weights. The indexes produced by Nationwide, the Lloyds Group and the ONS are based on data on mortgage loans. The one from the Land Registry is based on repeat sales. Panel B shows lagged monthly 1-year moving averages of the ICRG measure of political risk and the 10-year bond yield spread versus the comparable UK bond. We generate aggregate values by weighting the observations according to the respective population shares in London. The ICRG indicators and bond yield spreads are normalized by subtracting the in-sample mean and dividing by the standard deviation.
Figure 3
House prices, demographic structure and economic conditions at ward level

The figure reports the cross-sectional distribution of house prices and characteristics, across the 624 London wards. In Panels A and B, we plot the increase in average ward-level prices for two different sub-periods, calculated in the Land Registry sample of housing transactions. Panel C illustrates the ward-level average net income, as reported by the Office for National Statistics in 2001. Panel D shows the ward-level fraction of people born outside the UK, as indicated in the 2001 census.

A: House price appreciation between 2001 and 2006
B: House price appreciation between 2007 and 2012

C: Net average income
ONS (2001)

D: Foreign-born people share
2001 census
The table reports the coefficients $\gamma_{k0}^k$ and $\gamma_1^k$ from the following hedonic regression:

$$\ln P_{i,t} = \delta_i + \phi_{w} + \beta X_i + \sum_{k \in K} \gamma_{0f_{w}^{k}}^{k} + \gamma_{1w}^{k} + \ln P_{w,t-1} + \ln P_{w,t-2} + u_{i,t},$$

where $f_{w}^{k}$ are the shares of people in ward $w$ born in world region $k$, $y_{w}$ is average net income and $P_{w,t}$ are average transaction prices in ward $w$ in period $t$. In this specification, $z_{1}^{k}$ is a top-quintile indicator of the ICRG index of political risk, corresponding to the nine world regions listed in Table 1. We report absolute values of all estimated coefficients and indicate negative values by using light shading. The coefficients are multiplied by a factor of 100, for easier interpretation as percentage points relative price appreciation. Statistical significance is reported through error bars, indicating 95% confidence intervals. The standard errors are double-clustered at the borough and time level.

![Figure 4](image-url)

**Figure 4**
Safe haven effects across world regions

Residents linked to foreign countries
- estimation at the level of world regions -

<table>
<thead>
<tr>
<th>Region</th>
<th>Percent price change following risk shock</th>
</tr>
</thead>
<tbody>
<tr>
<td>Southern Europe</td>
<td>10</td>
</tr>
<tr>
<td>Middle East</td>
<td>9</td>
</tr>
<tr>
<td>South and Central America</td>
<td>8</td>
</tr>
<tr>
<td>Asia-Pacific</td>
<td>7</td>
</tr>
<tr>
<td>Northern EU and North America</td>
<td>6</td>
</tr>
<tr>
<td>Russia</td>
<td>5</td>
</tr>
<tr>
<td>Africa</td>
<td>4</td>
</tr>
<tr>
<td>Eastern Europe</td>
<td>3</td>
</tr>
<tr>
<td>South Asia</td>
<td>2</td>
</tr>
<tr>
<td>Net average income</td>
<td>1</td>
</tr>
</tbody>
</table>
The price impact of safe haven flows: time series dynamics

The figure reports the estimated average response of house prices in wards with high shares of foreign born people, following a one standard deviation shock to foreign political risk. The empirical specification corresponds to the following system of equations:

\[ s^k_t = \mu^k + \sum_{q=1}^{Q} \rho_q s^k_{t-q} + \sum_{q=1}^{Q} \zeta_q z^k_{t-q} + u^k_t, \quad \text{and} \quad z^k_t = \varphi^k + \sum_{q=1}^{Q} \pi_q z^k_{t-q} + \varepsilon^k_t, \]

where \( s^k_t \) is the price spread in period \( t \) between the top and bottom 20% of wards with respect to the share of people born in country \( k \) and \( z^k_t \) is the level of the ICRG index of political risk. Here, we consider the case \( Q = 20 \) months. The estimation is carried out in the sample of high-risk countries. The shaded areas indicate 90% confidence intervals, based on Driscoll-Kraay (1998) standard errors with a maximum lag length of 12 months.
Placebo tests

The figure reports estimated distributions of the coefficients $\zeta$ and $\beta$ from the following dynamic monthly panel regression specification:

$$s^k_t = \mu^k + \delta_t + \rho_1 s^k_{t-1} + \rho_2 s^\tilde{k}_{t-2} + \zeta z^k_{t-1} + \beta z^\tilde{k}_{t-1} + u^k_t,$$

with $\tilde{k} \in K$ and $\tilde{k} \neq k,$ across a set of $N=2,000$ draws during which we pair each country $k$ with a randomly drawn country $\tilde{k}$ (with replacement). In each draw, we consider the political risk indicator in country $\tilde{k}$ as an additional explanatory variable for the price spread between the top and bottom 20% of wards with respect to the share of people born in country $k$. We re-estimate the panel coefficients $\zeta$ and $\beta$ for each complete cross-country set of draws. In this specification, $z^k_t$ is a top-quintile indicator of the ICRG index of political risk. In Panel A, we report the distribution of estimated coefficients $\zeta$ when restricting $\zeta = 0$. The estimation is carried out in the sample of high-risk countries. The dashed line indicates the estimated coefficient $\zeta$, as reported in Table 4, in a specification which includes time fixed effects.

Panel A

Effect of randomly assigned political risk index levels on cross-ward house price spreads

Panel B

Effect of political risk on the same country’s price spread

Effect of alternative (placebo) randomly assigned levels of political risk
The figure reports the estimated coefficients $\gamma^{k,\eta}_0$ and $\gamma^{\eta}_1$ from the following hedonic regression:

$$\ln P_{i,t} = \phi_w + \delta_t + e^t + \beta X_i + \sum_{\eta=1}^3 \left( \sum_{k \in K} f^{k,\eta}_w x_k + \gamma^{k,\eta}_0 f^{k,\eta}_w z_{k,t} + \gamma^{\eta}_1 y_w t - 1 \right) + \ln P_{w,t-1} + \ln P_{w,t-2} + u_{i,t},$$

where $f^{k,\eta}_w$ are the shares of people in ward $w$ born in world region $k$ and $y_w$ is average net income. The parameter $\eta$ indicates the price category of property $i$. The thresholds which determine the price category are given by the cross-sectional 70th and 90th percentiles of the distribution of prices in each year and borough. The coefficients $\gamma$ correspond thus to a triple interaction term between the ward-level characteristics $f^{k,\eta}_w$ or $y_w$, the external factor $z_{k,t}^\eta$ and price category dummies. In this specification, $z_{k,t}^\eta$ is a top-quintile indicator of the ICRG index of political risk, corresponding to the nine world regions listed in Table 1. We report absolute values of all estimated coefficients and indicate negative values by using light shading. The coefficients are multiplied by a factor of 100, for easier interpretation as percentage points relative price appreciation. Statistical significance is reported through error bars, indicating 95% confidence intervals. The standard errors are double-clustered at the borough and time level.
The evolution of foreign-born people shares through time

The figure reports the coefficients $\rho^k$ from the regression:

$$\Delta f_{w,2011}^k = \alpha + \rho^k f_{w,2001}^k + \beta \Delta f_{w,2011}^{UK} + \epsilon_{w,2011}$$

where we condition the change between 2011 and 2001 in the share of people in ward $w$ originating from country $k$ on the starting level of this share in 2001. The estimation sample consists of the 624 London wards. We normalize the variables by subtracting the mean and dividing by the standard deviation. We report absolute values of all estimated coefficients and indicate negative values by using light shading. Statistical significance is reported through error bars, indicating 95% confidence intervals. The estimated standard errors are White heteroskedasticity-robust.
Figure 9
Cross-country overview of estimated coefficients

The figure documents the cross-sectional patterns of the estimated coefficients identifying safe haven effects in the London housing market. On the horizontal axes, we report the point estimates of the coefficients $\gamma_0^k$ from country-by-country versions of equation (3):

$$\ln P_{i,t} = \delta_t + \phi_w + \beta X_{i,t} + (\gamma_0^k f_k + \gamma_1^k y_w) z_{t-1}^k + \ln P_{w,t-1} + \ln P_{w,t-2} + u_{i,t},$$

where $f_k^w$ are the shares of people in ward $w$ born in country $k$ and $y_w$ is average net income. $z_t$ is the ICRG index of political risk. On the vertical axes, we report the corresponding point estimates of the coefficients $\gamma_1^k$, for each country $k$. The lines indicate univariate cross-country fitted values.

Registry dataset

Loans dataset