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

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This version: April 2015
First draft: November 2013

Abstract

Political economy explanations for the Lucas puzzle posit that private capital flies from poor countries to safe assets in rich countries to insure against domestic political risk. We find evidence that this mechanism helps to explain recent, widely-noted house price variation in global cities such as London and New York. We apply a new empirical approach, based on the insight that foreign capital may exhibit "home bias abroad," on large databases of housing transactions in London over the past two decades. The effects of foreign risk on London house prices are long-lasting but temporary, and vary interestingly across source countries. We also connect these findings to patterns of cross-border immigration.

JEL Codes: C53, D80, E47, F21, G12, G15.

*The previous version of this paper was circulated under the title "Preferred Habitats and Safe-Haven Effects: Evidence from the London Housing Market". We gratefully acknowledge Nationwide Building Society and the Land Registry for providing us with access to the data, and the Alfred P. Sloan Foundation for financial support. We thank John Campbell, Jonathan Cohn, Stijn Claessens, Yongheng Deng, Stefano Giglio, Cam Harvey, Erik Hurst, Ralph Koijen, Tim Landvoigt, Florencio Lopez-de-Silanes, John Muellbauer, Philipp Schnabl, Kevin Sheppard, Andrei Shleifer, seminar and conference participants at the Saïd Business School, the Oxford-Man Institute of Quantitative Finance, the Oxford Department of Economics, Goethe University, the MIT Sloan Finance Seminar, the Dutch National Bank, the Duisenberg Institute of Finance, Cass Business School, IMF, Imperial College, the University of Texas at Austin, the University of Illinois at Urbana-Champaign, the Darden International Finance Conference, the ABFER Meeting, the NBER Asset Pricing Program Meeting, the Paul Woolley Centre Conference at the LSE, the Federal Reserve Board and the National University of Singapore for useful comments, and Job Boerma and Shuang Zheng for dedicated research assistance.

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I. Introduction

Standard neoclassical models of trade and growth, as Lucas (1990) demonstrates, predict a "downhill" flow of capital from rich to poor countries, driven by the relatively higher marginal product of capital in poor countries. Paradoxically, the data tell the opposite story – capital seems to flow "uphill" from poor to rich countries – confronting economists with the now famous Lucas puzzle.[1]

Numerous theoretical explanations have been put forward for the Lucas puzzle. One class of explanations emphasizes differences in fundamentals, such as production technologies, between poor and rich countries. Another highlights that capital market frictions such as credit risk or asymmetric information impede flows of capital to poorer countries.[2] A particularly intriguing strand of the literature posits that private capital exits poor and risky countries in response to changes in the domestic political environment, to seek safety in rich-country assets.[3] These political economy explanations have important implications for the destination of the capital as well as the source, as such capital flight can create severe price distortions in a narrow range of "safe" rich-country assets. This effect will be magnified if the destination assets are relatively illiquid to begin with.

An important current issue is seemingly very consistent with the predictions of these political economy explanations for the Lucas puzzle. Global cities such as London, New York, and Singapore, which are relatively open to money and immigration from overseas, have witnessed sharp upward movements in residential real estate prices. These movements have been attributed – often stridently given their implications for housing affordability for local residents – to an influx of private capital seeking a "safe haven" from movements in risk in relatively poor and risky countries.[4]

Going from the intuitive plausibility of these stories to more rigorous empirical evidence requires confronting several obstacles. First, precise data tracking cross-border transactions in

[1] See, for example, Prasad, Rajan, and Subramanian (2007) for evidence of the strong uphill flow of capital.

[2] Lucas (1990) proposes human capital externalities as a fundamental solution to the puzzle. Gertler and Rogoff (1990), and Reinhart and Rogoff (2004), for example, emphasize the importance of capital market frictions.

[3] For example, Alesina and Tabellini (1988) show that uncertainty about the political environment (whether "workers" or "capitalists" are in power) can lead to capital flight from poor countries to rich countries, as poor country private capital seeks insurance against future taxation. Tornell and Velasco (1992) suggest that in the absence of adequate property rights in poor countries, an inferior production technology with good property rights enforcement in a rich country (or a "safe" bank account, in their terminology) will be the destination of uphill capital flows. In related work, Carroll and Jeanne (2009) solve a theory model in which poor, fast-growing country residents experience higher idiosyncratic risk and a greater precautionary motive to hold foreign assets.

[4] See, for example "Real estate goes global," *The New Yorker*, 26 May 2014, "Stash pad," *New York Magazine*, 29 June 2014, "Vancouver's real estate boom: The rising price of heaven," *The Globe and Mail*, 10 October 2014, and "Immigration is driving up home prices," *Daily Mail*, 26 October 2014.

residential real estate is generally not available. Second, capital flight driven by "safe haven" motives occurs 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 confound the use of pure time-series methods to convincingly attribute these house price movements to price pressure from foreign capital flows, rather than to movements in price-relevant information, the impact of regulation, variation in credit availability, or other crisis-induced market frictions.

In this paper, we follow a new identification strategy to resolve these obstacles, and apply it on large databases of historical housing transactions in London, a city on the frontlines of this issue.[5] We find that political risk in the Middle East, China, Russia, and the Asia-Pacific region, and importantly, during the recent sharp downturn in European fortunes, from countries in Southern Europe, help to explain London's recent house price fluctuations. We view this as evidence supporting the predictions of the political economy explanations of the Lucas puzzle. In addition, our paper provides some of the only rigorous empirical evidence to underpin important contemporary debates on the effects of foreign capital on domestic house prices.

Our identification strategy relies on the idea 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 city with a high concentration of source-country-origin residents.[6] Our method therefore estimates the impacts of foreign demand on house prices using the within-city cross-sectional variation in house prices in combination with time-series variation in foreign risk.

Consider the following example of 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 units, 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 rates of price

[5] See, for example "Rule of law is central to London's safe haven status," *The Financial Times*, 11 March 2013, "London remains top safe haven for property investors", *The Telegraph*, 13 March 2013, and "Live and let buy: Why an influx of foreign money is good for London's property market," *The Economist*, 9 November 2013.

[6] We provide a number of justifications for this identifying assumption in the Methodology section below.

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, but this can easily be generalized to other indicators of the attractiveness of particular areas of London to specific countries. The design of the method allows us to derive causal inferences with some confidence, as concerns of reverse causality, where London sub-area price increases drive aggregate foreign country risk, are limited in this setup.

The foreign demand effects that we estimate are likely a lower bound, 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 1.62 percentage points higher in months following elevated (top quintile) risk in that country. Adding a new dimension to the current political debate on this issue, we find that the foreign demand price-impacts that we estimate are long-lived but transitory, becoming statistically indistinguishable from zero after roughly two years.

Using our identification strategy, we are also able to predict variation in London housing transactions volumes. Perhaps more importantly, we are well able to predict prices when we instrument for volume using foreign risk. 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).

These effects survive a number of robustness checks and placebo tests. It is worth mentioning here that the "home bias abroad" channel is distinct from a related effect, which is the concentration of safe-haven demand from high net-worth foreigners in upscale London regions. Following increases in foreign risk, we find that London areas with high average income levels experience house price appreciation of an average of 2.21 percentage points over and above the effects on high foreign-origin-share regions of London.[7]

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

[7] 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)).

higher shares of residents originating from their countries, but it appears that these countries are less likely to direct capital towards premium residential areas.

In addition to our main focus on the Lucas puzzle, our work is related to a number of different literatures. Our identification strategy relies on "home bias abroad," which connects our study to work on home-bias and loyalty based portfolio choice (see, for example, Lewis 1999, Coval and Moskowitz, 2001, and Cohen, 2009), as well as to work on the effect of "preferred habitats" on portfolio choice.[8] 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 on mortgage indebtedness and residential property markets.[9]

Our identification strategy using the foreign-origin share links our paper with the work of Card and DiNardo (2000), Card (2001), and Saiz and Wachter (2011); our work is distinguished by its focus on the impacts of foreign demand on real estate prices. Aizenman and Jinjarak (2009), Jinjarak and Sheffrin (2011), and Favilukis et al. (2013) also analyze how foreign capital flows affect house prices – our work is distinguished from these papers by our novel cross-sectional approach, and our implementation using transaction data.

Others have identified that capital flows are not just "pulled" by returns in destination countries, but also "pushed" by movements in risk in source countries (see, for example Griffin et al., 2004, and Forbes and Warnock, 2011). Our results also suggest that cross-border capital flows 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).[10] Finally, we view our cross-sectional identification strategy as a useful complement to the time-series approaches that have generally been used to identify safe-haven effects on asset prices.[11]

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

[9]Our work is also related to Guerrieri et al. (2013), who analyze the relationship between demographics and house price dynamics across city neighbourhoods. Their focus is different from ours, in that they are concerned with spatial equilibria arising from within-city migration, rather than in the effects of overseas capital flows on the housing market.

[10]Also see Claessens and Forbes (2001) and Karolyi (2003) for surveys of the literature on international financial contagion.

[11]See, 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).

The remainder of the paper is organized as follows. Section II describes our methodology, and Section III the data that we employ. Section IV presents the results from our empirical estimation. Section V explores how these safe-haven effects in the residential real estate market relate to overseas immigration. Section VI concludes. An online appendix contains supplementary results and robustness checks, and is available at the URL specified in the references section.

II. Methodology

A. *Outbound Capital Flows and Political Risk*

Our goal is to identify the effect of risk-induced foreign capital flows on the price of residential real estate in London. Directly testing for such effects is impossible, as data on cross-border transactions in residential real estate are generally not available.[12] We therefore rely on our novel identification strategy to extract information from the cross-section of London house prices conditional on risk movements in source countries.

We can however check whether measures of aggregate capital outflows from these source countries are driven by political risk, an essential building block for our analysis. To motivate our more involved analysis, therefore, we first employ a large cross-country dataset compiled by Broner et al. (2013), which is based on the Balance of Payments Statistics of the IMF over the past two decades, and estimate the following panel specification:[13]

$$Outflows_t^k = \mu^k + \delta_t + \rho Outflows_{t-1}^k + \tau^k t + \beta_0 z_t^k + \beta_1 z_{t-1}^k + \boldsymbol{\gamma} \mathbf{X}_t^k + \varepsilon_t^k. \qquad (1)$$

The dependent variable captures gross capital outflows from domestic residents in a set of relatively poor and risky countries k in year t, z_t^k is the ICRG index of political risk (the main measure of risk that we employ in our analysis), and which we describe below, and X_t^k is a vector of three control variables: capital inflows by foreign agents into country k, the trade balance, and the real GDP growth rate.

[12] In the online appendix, we report the evolution of foreign capital entering the London *commercial* real estate sector. During recent years, foreign demand for UK commercial real estate has also increased markedly. However, any further analysis of this phenomenon is hindered by the lack of information about the physical characteristics and the geographical location of the properties.

[13] In these data, capital flows are calculated as changes in the reporting country's assets and liabilities vis-a-vis non-residents. Positive values of gross capital outflows therefore correspond to an increase in the holdings of foreign assets by domestic agents. In the online appendix, we report summary statistics and illustrate the sample coverage of this dataset.

[Insert Table 1 about here]

Table 1 shows that capital outflows are positively related to increases in domestic political risk. In response to a one standard deviation increase in the value of the ICRG index, capital outflows rise by 0.43 percent of GDP. Interestingly, we find that capital flows respond quite rapidly to changes in the domestic political situation. When contemporaneous and lagged values of the ICRG index are in the specification simultaneously, the contemporaneous effect is strong, and the lag effect is weak. The second column of the table repeats the estimation procedure for the remaining set of relatively rich and less risky countries, and shows very little role for movements in political risk in driving capital outflows from such countries. Consistent with the political economy explanations for the Lucas puzzle, the presence of strong local institutions and lower levels of political conflict appears to damp down any motivation for capital flight from richer countries.

We now turn to our method for estimating the effects of such capital flight on "safe-haven" residential real estate in London.

B. A Hedonic Pricing Model with Foreign Demand

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

$$\ln P_{i,t} = \alpha + \Pi_{w,t} + \beta \mathbf{X}_{i,t} + u_{i,t}. \tag{2}$$

Here, $P_{i,t}$ is the price of property i (which is physically located in location w within London), measured in month t. $\mathbf{X}_{i,t}$ is a vector of hedonic characteristics for property i, and the second component on the right-hand side of equation (2), denoted $\Pi_{w,t}$, denotes unrestricted location (w) cross time (t) fixed effects.

Our identification strategy 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 $t-1$.[14] We model $\Pi_{w,t}$ in the following fashion:

$$\Pi_{w,t} = \rho_1 \Pi_{w,t-1} + \rho_2 \Pi_{w,t-2} + \delta_t + \phi_w + \sum_{k \in K} \gamma_0^k f_w^k z_{t-1}^k + \gamma_1 y_w \overline{z}_{t-1}. \tag{3}$$

[14] 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.

In equation (3), δ_t are time fixed effects which absorb common time-variation in London house prices,[15] ϕ_w are location-specific fixed effects which control for fixed price differentials across London locations, and f_w^k links location w with country k. In our empirical implementation, f_w^k is the share of residents in London electoral ward w (we often refer to these simply as "wards") that were born in country k. The remaining coefficients are purely estimated off simultaneous cross-time- and cross-ward variation in house prices, and we describe them more fully below.

In the appendix, we show that estimating the model described in equations (2) and (3) is equivalent to estimating the following specification:

$$\begin{aligned} \ln P_{i,t} &= \mu + \delta_t + \phi_w + \sum_{k \in K} \gamma_0^k f_w^k z_{t-1}^k + \gamma_1 y_w \overline{z}_{t-1} + \beta \mathbf{X}_{i,t} \\ &\quad + \rho_1 \ln \overline{P}_{w,t-1} + \rho_2 \ln \overline{P}_{w,t-2} + \nu_{i,t}. \end{aligned} \tag{4}$$

In equation (4), γ_0^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.

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 neighborhoods 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 show evidence later in the paper that patterns of growth in ward-level immigration shares in our data are consistent with this line of reasoning.[16] Second, 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, through direct communication between foreign-origin local London residents and overseas buyers, or because of specialty realtors, local legal firms, and

[15] 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.

[16] 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 also 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.

other soft infrastructure set up to match overseas purchasers of specific nationalities with property investment opportunities in specific London areas.[17] Third, connections between foreign purchasers and country-of-origin London residents may facilitate monitoring of the physical investment, that is, property maintenance or managing rental tenants. This is an especially useful service when purchasers are located in a different country.

If foreign buyers choose to purchase London properties in locations *not* highly populated by residents originating from their countries, we would estimate γ_0^k to be negative.[18] In this sense, the estimated coefficients allow us to derive interesting insights about specific countries in addition to identifying safe-haven effects. γ_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 (an effect we rarely find in our empirical results). Throughout, the obvious null hypothesis is that $\gamma_0^k = 0$.

We note that the share of foreign origin residents in an electoral ward f_w^k may be correlated with other attributes of the ward which attract foreign capital other than the familiarity channel that we outline above. An obvious possibility is that f_w^k may be high in relatively more desirable parts of London (such as Chelsea or Mayfair), if high-net worth foreigners generally settle in such 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 $\overline{z}_t$, a weighted average risk measure across countries, with weights given by country-k population shares in London.

C. Implications for Housing Transaction Volume

If we are able to identify 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.

[17] 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.

[18] 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.

We therefore estimate:

$$\ln V_{w,t} = \vartheta_w + \varsigma_t + \sum_{k\in K} \chi_0^k f_w^k z_{t-1}^k + \chi_1 y_w \overline{z}_{t-1} + \upsilon_{w,t}. \tag{5}$$

In equation (5), we test whether χ_0^k and χ_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.[19] 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 demand effects.

We therefore instrument for volume using $\widehat{\ln V_{w,t}} = \sum_{k\in K} \chi_0^k f_w^k z_{t-1}^k + \chi_1 y_w \overline{z}_{t-1}$, and estimate:

$$\ln P_{i,t} = \mu + \delta_t + \phi_w + \theta^S \widehat{\ln V_{w,t}} + \beta \mathbf{X}_{i,t} + \rho_1 \ln \overline{P}_{w,t-1} + \rho_2 \ln \overline{P}_{w,t-2} + u_{i,t}. \tag{6}$$

We expect θ^S to be positive and statistically significant.

D. *The Persistence of Price Impacts from Foreign Demand*

We use the dynamics of house prices in our hedonic regression specification (4) to estimate 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 interest of focusing on time-series variation rather than the cross-sectional variation in these effects.

We define s_t^k 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_w^k :

$$s_t^k \equiv \overline{\Pi}_t^{high,k} - \overline{\Pi}_t^{low,k}.$$

We then estimate:[20]

$$\begin{aligned} s_t^k &= \mu^k + \delta_t + \sum_{q=1}^{Q} \rho_q s_{t-q}^k + \sum_{q=1}^{Q} \zeta_q z_{t-q}^k + u_t^k, & (7) \\ z_t^k &= \theta^k + \sum_{q=1}^{Q} \pi_q z_{t-q}^k + \varepsilon_t^k. & (8) \end{aligned}$$

[19] Stein (1993) presents a rational model in which price increases and declines affect available downpayment 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.

[20] The appendix describes exactly how specification (7) derives from the benchmark model described in equations (2) and (3).

The second of these equations is a simple autoregressive model for the country-specific risk factors z_t^k. This specification relies on the assumption that there is no feedback from intra-London variation in house prices to foreign risk. We use equations (7) and (8) to generate an impulse response function of s_t^k to a shock ε_t^k to risk, to check whether the effects are permanent or temporary – as we might expect if foreign capital generates temporary dislocations from fundamentals in the London housing market.

E. 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 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 do not expect to observe safe-haven effects for these countries, so we test the null hypothesis $\zeta_q = 0$ in equation (7) separately for these countries.

In a more formal placebo test, we construct a synthetic variable $z_{t-1}^{\widetilde{k}}$ by drawing with replacement from the set $K \setminus k$ for each of the 41 high-risk countries k. That is, we match regions of London with foreigners originating from particular countries with movements in risk $z_{t-1}^{\widetilde{k}}$ from randomly selected *other* countries $\widetilde{k}$. We include this placebo variable $z_{t-1}^{\widetilde{k}}$ in our panel regression (7) and estimate:

$$s_t^k = \mu^k + \delta_t + \rho_1 s_{t-1}^k + \rho_2 s_{t-2}^k + \zeta z_{t-1}^k + \beta z_{t-1}^{\widetilde{k}} + u_t^k. \tag{9}$$

We run $2,000$ such regressions with random draws of $z_{t-1}^{\widetilde{k}}$, and in one version of the placebo test, we restrict $\zeta = 0$ and plot the distribution of the estimated β coefficient to assess the power of our identification approach, and in the other, we simply include both z_{t-1}^k and $z_{t-1}^{\widetilde{k}}$ to check the distributions of ζ and β when both are included.

F. Notes on Estimation

In our empirical implementation, we construct z_{t-1}^k 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, and zero otherwise.[21] We also transform the foreign-origin shares f_w^k, 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 (7) and (8), and in the small final note on country-specific effects. In these two cases, we simply estimate effects which are linear in z_{t-1}^k and f_w^k to maximize the number of observations utilized.[22]

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 γ_0^k 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.

In addition to the inclusion of lags, we deal with residual persistence in ward-level price changes 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 time, and borough (a broader geographical unit of aggregation than wards – there are 32 boroughs in London encapsulating 624 wards) as a way to capture spatial effects, i.e., cross-ward-time correlation. In specification (7), 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.[23]

We conduct extensive robustness checks on our results, which we discuss briefly in the text, leaving details for the online appendix to the paper. One worth mentioning here 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.

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[21] We verify using equation (7) with decile dummies for z_{t-1}^k that this approach is sensible (these results are in the online appendix) – the effects of z_{t-1}^k are large and statistically significant only when z_{t-1}^k is in its top few deciles.

[22] 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.

[23] 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).

III. Data

We employ a number of datasets in our study. We use two databases of housing transactions in the UK, from the UK Land Registry, and from the Nationwide Building Society. We use census data from the Office for National Statistics in the UK to identify the country of origin of London residents. Finally, we use time-series indexes of country-level economic and political risk measures from ICRG.

A. 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 the whole of England and Wales, with a few property characteristics including the type of house (e.g., whether it is an apartment), the tenure status (leasehold or freehold) and an indicator for new builds.

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 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 the share of foreign residents, and the income level in each electoral ward.

Table 2 describes the distribution of transaction volumes and prices in the data. The number of housing transactions dips in 2008 and remains low relative to historical averages during the credit crunch. Over the same period, however, average house prices increase.

The Registry dataset does not contain information about the provenance of buyers. We note that these data are likely to miss extremely highly-priced properties in premium areas of London, because of a higher tendency of these types of properties to be classified as commercial real estate and to be purchased using private trusts or specialized companies. Nevertheless, we still find that during the years characterized by large political crises around the world (1998 in Russia, 2001 in South- and Central America and 2012-2013 in the Middle East, Russia and Europe), residential transactions in the Registry data include exceptionally expensive houses.

For example, the overall most expensive residential property in London was sold for £55 million in 2012. In contrast, the distribution of transaction prices is less disperse during periods of relative calm, such as 2003 and 2004.[24] This is reassuring, as it suggests that these data are

[24] Market-level evidence collected by private agencies in London tends to support this hypothesis. For ex-

able to capture some part of the demand of wealthy (foreign and local) private investors in the London housing market.

B. 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, 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 a wider set of hedonic characteristics.[25] 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 spill over into the domestic housing market.

In Table 2, we show that the average price is equal to £265,400 across all properties lodged with the Land Registry and £203,300 for the ones which are subject to a mortgage loan from Nationwide. This is due to marked differences at the top of the distribution. The most expensive property used as collateral had a value of £4.4 million, which lies an order of magnitude below the most expensive residential property sold in London during the same period.

[Insert Table 2 about here]

ample, the estate agency Savills (Market Research Report, "The World in London") estimates that around 20% of low-end newly built properties were bought by foreigners.in 2013. This number rises gradually for more expensive properties and reaches 90% for luxury units.

[25] 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.

C. UK Office for National Statistics (ONS)

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 with wards using property postcodes and the 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_w^k. 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 Table 3.

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 have important differences with the remainder, but as described below, such static characteristics of wards are not allowed to affect our estimation as we include ward-fixed effects in our estimation.

We observe 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-origin people draw upon sources of funds which lie outside the UK mortgage system, and at least partially, may come from overseas.[26]

[Insert Table 3 about here]

[26] 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 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.

[Insert Figure 1 about here]

D. Measures of Foreign Risk

We measure economic and political risk in foreign countries each month using three separate variables. First, in the majority of our analysis we use the International Country Risk Guide (ICRG) indexes of political risk. These indexes rate each country each month 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 use weights based on the shares of respective populations in London to build our time series for world regions.[27]

Glaeser et al. (2004) note that the ICRG indicators do not perfectly describe the permanent state of country-level political institutions, but instead reflect changes in these institutions over

[27] 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.

time. This is well-suited to our purposes, as we are most interested in capturing changes in the political situation in a given country, rather than static differences across institutions. These authors also emphasize that the ICRG indicators are subjective assessments of risk, which we also find desirable, as we are interested in sentiment about the local political environment driving capital flight. The ICRG data are also used by Erb, Harvey, and Viskanta (1996 a,b), who show 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.

In our placebo test and other parts of our analysis we distinguish between high- and low-risk countries, classified based on their average levels of risk (which we report in the online appendix). 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.[28] 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. We describe these data more fully 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. This measure is only available for a small set of countries, and we use it for Spain, Italy, China, and India for the period between 2004 and 2013.[29]

[28] 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.

[29] We find that higher economic policy uncertainty is associated with increases in the ICRG indexes – 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.

IV. Results

A. *Time-Series Patterns*

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 those in 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.

[Insert Figure 2 about here]

While these patterns are striking, as mentioned earlier, these time-series do not allow us to attribute causality because of the possibility of common shocks affecting aggregate house prices and political risk. In addition, we might wish to separately identify any effects on London house prices by the country of origin of the safe-haven demand.

We turn next to better identification of foreign demand effects on London house prices, and a deeper characterization of the underlying economic motivations for this demand.

B. *Preliminary Results: Total Foreign-Origin Share*

Figure 3 provides a preliminary illustration of our main results. The panels in the figure show maps of London, with electoral wards color-coded in a "heatmap" – red shades indicate greater levels of particular variables, and blue shades indicates 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 as the crisis unfolded. Panel C of the figure shows that the areas of appreciation in Panel B do appear to be relatively 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 wards measured in the 2001 census. The relationship between Panel B and Panel D is visually striking – house price appreciation between 2007 and 2012 in Panel B appears to move away from the locations in Panel A, now lining up well with the share of foreign-origin people in London wards in Panel D. 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 we can estimate using our method. We first turn to describing the effect of hedonics on London house prices, and then turn to these cross-country results.

[Insert Figure 3 about here]

C. Hedonics

Table 4 shows the coefficients β 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.[30]

Finally, Panel C of Table 4 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.

[Insert Table 4 about here]

D. The Effects of Foreign Demand

Figure 4 shows the coefficients γ_0^k estimated from equation (4) for each of the nine world regions, as well as the coefficient γ_1, which captures safe-haven effects in wards with high net average income.

The figure shows that elevated political risk in Southern EU, the Asia-Pacific region, the Middle East, and South- and Central America is associated with subsequent movements in 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 foreign-origin shares will be elevated by 1.62 percentage points. Controlling for this foreign-born share effect, the effect of foreign risk amounts to 2.21 percentage points on wards with high average income. These numbers come from the Land Registry estimates; comparable magnitudes are evident from the Nationwide data.

[Insert Figure 4 about here]

[30] 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.

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 γ_0^k and γ_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 γ_0^k, 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, South America, and the Asia-Pacific region.

We further investigate the cross-country patterns of foreign demand effects from individual countries, allowing the income interaction on y_w^k to 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 5 plots these income and foreign-origin share coefficients (γ_0^k and γ_1^k) against one another. To aid understanding of this plot, consider a hypothetical country k for which γ_0^k and γ_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.

The figure shows a negative relationship between the two estimated coefficients, which we believe captures economically interesting cross-country variation in the particular form of the "home bias abroad" exhibited by 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 not to greatly favour 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.

[Insert Figure 5 about here]

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.

E. *Evidence on Transaction Volumes*

Table 5 shows results from estimating equations (5)-(6). 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 8.53 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 χ_0^k and χ_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 (6), 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.

[Insert Table 5 about here]

F. *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 equation (7) 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 6 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.

[Insert Table 6 about here]

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 of these specifications, we do not introduce time fixed effects as data on these alternative measures is available for a more limited cross-section of countries (13 for bond yield spreads, and 4 for economic policy uncertainty).

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 risk.

[Insert Figure 6 about here]

In Figure 6, we estimate the evolution of the spread in response to a one standard deviation shock to foreign risk, using equations (7) and (8).[31] The estimation results using the Registry dataset show that the predicted spread in house prices begins at 0.18 percent, and reaches a peak at 0.76 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 less 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 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

[31] 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 transfer of funds) initiated by overseas investors.

of risk, and substantial house price inertia, a likely consequence of the difficulty of arbitraging price discrepancies in this market.

This specification is also informative about the total extent of variation in price spreads across wards that are explained by foreign demand effects. In the entire sample of countries, these effects explain 6.56% of the variation in price spreads. For the sub-group of Southern European countries, the explanatory power reaches 51.50%.[32]

G. *Placebo Tests*

As discussed earlier, we identify 14 relatively wealthy and 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 the ICRG indexes. Our first test is to estimate (7) for this low-risk sample.

We report the results in the last column of Table 6, which shows 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" to London real estate from the high-risk countries, consistent with political economy explanations of the Lucas puzzle, 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.

[Insert Figure 7 about here]

Figure 7 reports the results of the placebo tests described in equation (9). In Panel A, we constrain the coefficient ζ to equal zero, and plot the estimated distribution of the coefficient β in equation (9) across a set of N=2,000 draws with replacement from the set of countries $K \setminus k$. The point estimate on ζ in Table 6 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 ζ and β simultaneously. The estimated coefficient ζ, plotted in the left panel varies between 0.077 and 0.148 percentage

[32] This number is calculated as the unadjusted R^2 statistic (coefficient of determination) from the regression specification (7), after filtering out the country fixed effects. In this estimation procedure, we do not allow for autoregressive dynamics and do not include time fixed effects, in order to isolate the overall long-run contribution of foreign risk to the price spreads across wards.

points, with a 5% lower percentile equal to 0.099 percentage points. In contrast, the coefficient β on randomly assigned risk, plotted in the right panel varies between -0.207 and 0.152 with far greater dispersion.

V. Economic Channels: Safe-Haven Effects and Immigration

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.[33] Simultaneously, there has been enormous public discussion of immigration, which may also be associated with movements of foreign capital into residential real estate. In this section we explore the extent to which our results are driven by these two related but distinct motivations for safe-haven residential real estate purchases driven by source-country risk.

We first check, using a version of equation (4) whether price impacts are highest in high-value properties, separately estimating coefficients γ_0^k and γ_1 conditional on levels of house prices. Figure 8 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, for the highest price category of houses, we now detect effects of Russian demand. Overall, it does appear that pure "safe-haven" capital preservation motivations drive some portion of foreign demand for London housing.

[Insert Figure 8 about here]

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

[33] There 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.

an implicit or explicit consideration of future London-bound immigration. To explore this channel, we estimate:

$$\Delta f^k_{w,2011} = \alpha + \rho^k f^k_{w,2001} + \beta^k \Delta f^{UK}_{w,2011} + e^k_{w,2011}. \tag{10}$$

We condition the change between 2011 and 2001 $\Delta f^k_{w,2011} = f^k_{w,2011} - f^k_{w,2001}$ in the share of people in ward w originating from country k on the starting level of this share in 2001, $f^k_{w,2001}$. The sign and significance of the coefficient ρ^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. $\Delta f^{UK}_{w,2011}$ 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. Equation (10) also provides further evidence on our identification strategy, if ρ^k is estimated to be significantly positive – 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.

[Insert Figure 9 about here]

Figure 9 shows estimates of equation (10) 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, for a smaller set of countries which includes Jamaica, Kenya, Bangladesh, and Zimbabwe, that we see the opposite pattern, namely that immigration to wards with high shares of residents originating from these countries is significantly lower than in other areas of London.

[Insert Figure 10 about here]

We complement this simple analysis by analyzing two sets of immigration statistics. First, we obtain the numbers of registrations for National Insurance from the Department for Work and Pensions.[34] Second, we obtain the numbers of visas granted to foreign citizens during the period between 2008 and 2013 from the UK Border Agency, with a particular focus on Tier 1 (Investor) visas. Such visas are granted to foreign investors investing more than £2 million in UK government bonds, share capital, or loan capital in active registered companies.[35] The issuance of such visas mean that foreign capital flows and flows of people to the UK can explicitly be linked to one another. Are they also linked to political risk in source countries? The top panel of Figure 10 shows that there is a strong positive correlation between changes in political risk in countries requiring a visa to enter the UK, and the number of investor visas granted by the UK to citizens of those countries. Complementing our analysis above, there is a remarkable similarity between the countries identified here – China, Russia, the Middle East – and using our methodology above. The online appendix shows that this positive association between country political risk and visas granted to residents of the country is restricted to the Tier 1 (Investor) type. That is, in the (poor) countries for which a visa is necessary in order to enter the UK, political risk appears to be associated with movements to the UK only for the wealthiest individuals (i.e. those holding at least £2 million of liquid and internationally transferrable assets). Interestingly, the bottom panel of Figure 10 shows that for countries for which a visa is not generally necessary to enter and work in the UK, high levels of political risk are also linked with a higher propensity to enter the UK labour market. This appears to hold especially for the Southern European countries, which we also find to be an important source of foreign demand in the London real estate market.

Overall, the evidence from visas points to two sources of effects. Political economy explanations of the Lucas puzzle concentrate on the motivations of those most likely to suffer expropriation or taxation consequences in poor countries, i.e., the wealthiest residents of those countries, who have the ability to achieve "escape velocity" from their home countries. We find evidence in the investor visas that this appears to be an important part of the story. The second source appears to arise from movements in political risk in countries which face relatively low impediments to moving capital (and labor) to London, such as the Southern European countries. This suggests that the political economy explanations of the Lucas puzzle interact with explanations for the puzzle which focus on the role of frictions in global capital markets.

[34] These data are available for the period between 2003 and 2013, for the countries with the largest immigrant populations in the UK, namely Poland, Spain, Italy, India, Portugal, France, Romania, Pakistan, China, Australia, Nigeria, the USA and Greece.

[35] In order to obtain an investor visa, foreign nationals have to prove their ability and express the intention to invest more than £2 million in the country. The visas are initially granted for a period of three years. Note that direct investments in companies engaged in property investment, property management or property development are not considered legitimate grounds for obtaining a Tier 1 visa.

VI. Conclusions

A hitherto untested set of explanations for the Lucas (1990) puzzle posits that private capital from poor countries may flee to safer rich country assets in response to changes in the domestic political environment. In this paper, we provide evidence in support of such political economy explanations for the puzzle, uncovering the effects of foreign capital flight into residential real estate in London, a contemporary issue of great public interest.

In order to show these effects, 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. 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 (2) implies that the average level of transaction prices across all properties in a ward in a time period can be written as:

$$\ln \overline{P}_{w,t} = \alpha + \Pi_{w,t} + \beta \overline{\mathbf{X}}_{w,t} + \overline{u}_{w,t}, \tag{11}$$

where $\overline{\mathbf{X}}_{w,t}$ 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 model ward-level prices as an $AR(2)$[36] in equation (3). Substituting equation (11) into equation (3), we get:

$$\begin{aligned} \Pi_{w,t} = {} & \rho_1(\ln \overline{P}_{w,t-1} - \alpha - \beta \overline{\mathbf{X}}_{w,t-1} - \overline{u}_{w,t-1}) + \rho_2(\ln \overline{P}_{w,t-2} - \alpha - \beta \overline{\mathbf{X}}_{w,t-2} \\ & - \overline{u}_{w,t-2}) + \delta_t + \phi_w + \sum_{k \in K} \gamma_0^k f_w^k z_{t-1}^k + \gamma_1 y_w \overline{z}_{t-1}. \end{aligned} \tag{12}$$

Substituting equation (12) into (2), our benchmark hedonic house price model can be expressed as:

$$\begin{aligned} \ln P_{i,t} = {} & \mu + \delta_t + \phi_w + \sum_{k \in K} \gamma_0^k f_w^k z_{t-1}^k + \gamma_1 y_w \overline{z}_{t-1} + \beta \mathbf{X}_{i,t} + \rho_1 \ln \overline{P}_{w,t-1} \\ & + \rho_2 \ln \overline{P}_{w,t-2} + b_1 \overline{\mathbf{X}}_{w,t-1} + b_2 \overline{\mathbf{X}}_{w,t-2} + \nu_{i,t}. \end{aligned} \tag{13}$$

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 (13), which produces very similar results.

The autoregressive nature of ward-level prices implies that $\nu_{i,t} = \rho_1 \overline{u}_{w,t-1} + \rho_2 \overline{u}_{w,t-2} + u_{i,t}$, leading to both serial- and ward-level correlation of error terms. We deal with this issue

[36] 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.

by reporting robust standard errors in all tables and figures, which are clustered along two dimensions, namely, borough and time.

Spreads Specification

Let $\overline{\Pi}_t^{high,k}$ 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 (3), we obtain:

$$\begin{aligned}\overline{\Pi}_t^{high,k} &= \rho_1 \overline{\Pi}_{t-1}^{high,k} + \rho_2 \overline{\Pi}_{t-2}^{high,k} + \delta_t + \overline{\phi}^{high,k} + \gamma_0^k \overline{f}^{high,k} z_{t-1}^k + \sum_{k' \in K, k' \neq k} \gamma_0^{k'} \overline{f}^{k'|high,k} z_{t-1}^{k'} \\ &\quad + \gamma_1 \overline{y}^{high,k} \overline{z}_{t-1},\end{aligned}$$

where $\overline{\phi}^{high,k}$ is the ward fixed effect, $\overline{f}^{high,k}$ is the average share of people born in country k, $f^{k'|high,k}$ is the average share of people born in country k' and $\overline{y}^{high,k}$ 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:

$$\begin{aligned}\overline{\Pi}_t^{low,k} &= \rho_1 \overline{\Pi}_{t-1}^{low,k} + \rho_2 \overline{\Pi}_{t-2}^{low,k} + \delta_t + \overline{\phi}^{low,k} + \gamma_0^k \overline{f}^{low,k} z_{t-1}^k + \sum_{k' \in K, k' \neq k} \gamma_0^{k'} \overline{f}^{k'|low,k} z_{t-1}^{k'} \\ &\quad + \gamma_1 \overline{y}^{low,k} \overline{z}_{t-1}.\end{aligned}$$

The spread s_t^k is the difference between $\overline{\Pi}_t^{high,k}$ and $\overline{\Pi}_t^{low,k}$:

$$\begin{aligned}s_t^k &= \rho_1 s_{t-1}^k + \rho_2 s_{t-2}^k + (\overline{\phi}^{high,k} - \overline{\phi}^{low,k}) + \gamma_0^k (\overline{f}^{high,k} - \overline{f}^{low,k}) z_{t-1}^k + \\ &\quad + \sum_{k' \in K, k' \neq k} \gamma_0^{k'} \left(\overline{f}^{k'|high,k} - \overline{f}^{k'|low,k} \right) z_{t-1}^{k'} + \gamma_1 \left(\overline{y}^{high,k} - \overline{y}^{low,k} \right) \overline{z}_{t-1}.\end{aligned} \tag{14}$$

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 order to obtain a parsimonious modeling framework and characterize the overall impact of foreign demand on the cross-ward distribution of house prices in London, we restrict the γ_0^k coefficients to be the same across countries. This implies that:

$$\begin{aligned}s_t^k &= \rho_1 s_{t-1}^k + \rho_2 s_{t-2}^k + (\overline{\phi}^{high,k} - \overline{\phi}^{low,k}) \\ &\quad + \gamma_0 \left[(\overline{f}^{high,k} - \overline{f}^{low,k}) z_{t-1}^k + \sum_{k' \in K, k' \neq k} \left(\overline{f}^{k'|high,k} - \overline{f}^{k'|low,k} \right) z_{t-1}^{k'} \right] \\ &\quad + \gamma_1 \left(\overline{y}^{high,k} - \overline{y}^{low,k} \right) \overline{z}_{t-1}.\end{aligned}$$

We show in the online appendix that the estimated γ_0 obtained in this framework is very similar to the result from a specification where we just consider the direct component and include time fixed effects, as reported in Table 6. The explanation is that the last two terms in equation (14) do not vary greatly across countries – the first principal component of the set of interaction terms $\sum_{k' \in K, k' \neq k} \left(\overline{f}^{k'|high,k} - \overline{f}^{k'|low,k} \right) z_{t-1}^{k'}$ explains 85% of their total variation across our full sample.

This derivation reveals what is intuitively clear, namely, 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.

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Table 1
Political risk and capital outflows: a cross-country view

The table reports results from the following panel regression specification:

$$Outflows_t^k = \mu^k + \delta_t + \rho Outflows_{t-1}^k + \tau^k t + \beta_0 z_t^k + \beta_1 z_{t-1}^k + \boldsymbol{\gamma} \mathbf{X}_t^k + \varepsilon_t^k, \tag{15}$$

where $Outflows_t^k$ are capital outflows by domestic agents in country k and year t, z_t^k is the average ICRG index level in country k and year t, and $\mathbf{X}_t^k$ is a vector of three control variables: the capital inflows by foreign agents, the trade balance and the real rate of GDP growth. In our estimation, we distinguish between two country groups, according to the average level of political risk, with low-risk countries having an average ICRG index below 20. We normalize the ICRG indexes by subtracting the mean and dividing by the standard deviation. In parentheses, we report heteroskedasticity-robust standard errors. *, **, *** denote statistical significance at the 10%, 5%, and 1% level, respectively.

	Capital outflows relative to GDP (in percent)	
	High-risk countries	Low-risk countries
β_0	0.43**	0.09
	(0.21)	(0.21)
β_1	-0.44	-0.19
	(0.28)	(0.20)

Table 2
Summary statistics

The table reports selected statistics which capture the distribution of observed prices and volumes, 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.

	Registry dataset					Loans dataset				
	Transaction prices (in £'000)				Transaction	Transaction prices (in £'000)				Transaction
Year	Min	Mean	Max	St. dev.	volume	Min	Mean	Max	St. dev.	volume
1995	50.1	113.6	3,750.0	107.7	84,525	-	-	-	-	-
1996	50.0	120.6	8,000.0	120.4	108,735	10.5	84.6	1,000.0	51.7	12,243
1997	30.0	132.7	7,500.0	139.0	133,598	17.5	97.4	920.0	57.6	13,108
1998	50.1	143.5	11,300.0	162.4	133,784	21.0	118.6	935.0	71.6	12,463
1999	50.0	163.1	32,500.0	190.9	162,930	24.0	135.4	980.0	77.3	19,330
2000	50.3	190.7	10,000.0	213.5	147,865	28.5	149.8	960.0	82.8	8,958
2001	50.0	205.7	24,800.0	222.7	162,705	35.0	169.8	971.0	89.3	7,522
2002	50.2	233.9	8,300.0	217.8	174,667	42.5	197.9	998.0	92.6	9,386
2003	50.3	250.0	9,250.0	217.1	155,115	57.2	221.3	950.0	94.6	7,371
2004	52.0	273.4	7,950.0	236.5	165,471	65.0	243.9	3,340.0	106.0	7,565
2005	52.0	289.9	15,200.0	266.6	137,660	83.0	256.4	1,955.0	108.3	6,649
2006	50.8	315.5	12,400.0	305.5	172,360	48.1	273.9	3,125.0	132.4	10,712
2007	52.5	352.4	19,000.0	361.7	166,417	68.0	312.8	2,500.0	154.1	9,143
2008	51.0	360.6	19,800.0	423.8	81,763	94.5	297.6	1,625.0	149.4	3,977
2009	50.0	361.2	12,500.0	385.8	75,379	70.0	294.2	1,850.0	148.4	4,394
2010	51.0	406.0	16,200.0	466.1	91,877	50.0	320.0	1,850.0	172.0	5,068
2011	50.8	418.6	19,300.0	493.0	89,762	81.0	347.5	4,400.0	205.5	6,356
2012	50.5	435.5	55,000.0	620.0	92,151	66.5	321.2	2,550.0	171.6	9,892
2013	10.0	475.7	39,000.0	642.5	108,383	-	-	-	-	-
Full sample	10.0	265.4	55,000.0	336.1	2,445,057	10.5	203.3	4,400.0	142.4	154,137

Table 3
List of countries and world regions

The table reports 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.

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		
Asia-Pacific		
China	Hong Kong	New Zealand
Malaysia	Singapore	Australia
Japan		
South Asia		
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 4
Estimated coefficients in the hedonic regression framework

The table reports estimated coefficients from the following hedonic regression:

$$\ln P_{i,t} = \mu + \phi_w + \delta_t + \beta \boldsymbol{X}_{i,t} + \nu_{i,t},$$

where $\boldsymbol{X}_{i,t}$ 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

Registry dataset

Semi detached house	-0.426***
Terraced house	-0.580***
Flat	-0.744***
New property	0.232***
Leasehold indicator	-0.257***

Loans dataset

Semi detached house	-0.141***
Terraced house	-0.191***
Detached bungalow	0.012
Semi detached bungalow	-0.065***
Purpose built flat	-0.232***
Purpose built maisonette	-0.274***
Flat conversion	-0.178***
Maisonette conversion	-0.208***
New property	0.117***
Leasehold indicator	-0.115***

Panel B

Loans dataset

First purchase	-0.056***
Two bedrooms	0.173***
Three bedrooms	0.229***
Four or five bedrooms	0.277***
More than five bedrooms	0.377***
Two bathrooms	0.057***
Three bathrooms	0.024***
More bathrooms	-0.005
Parking space	0.042***
Single garage	0.080***
Double garage	0.098***

Construction date	
1900 to 1920	-0.004
1920 to 1940	-0.060***
1940 to 1960	-0.133***
1960 to 1980	-0.141***
1980 to 2000	-0.026***
after 2000	0.001
Floor area	
50 to 70 m^2	0.112***
70 to 90 m^2	0.189***
90 to 110 m^2	0.282***
110 to 130 m^2	0.380***
130 to 150 m^2	0.471***
150 to 170 m^2	0.570***
above 170 m^2	0.738***

Table 4
Estimated coefficients in the hedonic regression framework
(continued)

Panel C

	Estimated time fixed effects	
	Registry dataset	Loans dataset
1996	0.052***	-
1997	0.184***	0.217***
1998	0.311***	0.354***
1999	0.436***	0.511***
2000	0.635***	0.699***
2001	0.742***	0.810***
2002	0.902***	0.982***
2003	1.027***	1.080***
2004	1.108***	1.149***
2005	1.139***	1.171***
2006	1.203***	1.255***
2007	1.342***	1.391***
2008	1.327***	1.337***
2009	1.233***	1.255***
2010	1.340***	1.350***
2011	1.358***	1.386***
2012	1.409***	1.418***
2013	1.490***	-

Table 5
Foreign risk and transaction volumes

In Panel A, we report the coefficients χ_0^k and χ_1 from the following regression:

$$\ln V_{w,t} = \vartheta_w + \varsigma_t + \sum_{k \in K} \chi_0^k f_w^k z_{t-1}^k + \chi_1 y_w \overline{z}_{t-1} + \upsilon_{w,t},$$

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

$$\begin{aligned} \ln P_{i,t} &= \mu + \delta_t + \phi_w + \theta \ln V_{w,t} + \rho_1 \ln \overline{P}_{w,t-1} + \rho_2 \ln \overline{P}_{w,t-2} + \beta \mathbf{X}_{i,t} + \nu_{i,t}, \text{ and} \\ \ln P_{i,t} &= \mu + \delta_t + \phi_w + \theta^S \widehat{\ln V_{w,t}} + \rho_1 \ln \overline{P}_{w,t-1} + \rho_2 \ln \overline{P}_{w,t-2} + \beta \mathbf{X}_{i,t} + \nu_{i,t}, \end{aligned}$$

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 3. 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

Southern Europe	8.53***	Russia	7.81***
	(2.68)		(2.31)
Asia-Pacific	10.58***	Eastern Europe	-8.50***
	(2.35)		(2.29)
Middle East	-1.85	South Asia	-3.31
	(3.81)		(4.15)
Africa	8.91**	Sth. and C. America	6.06**
	(3.54)		(2.76)
North EU and America	8.96**	Net average income	-3.71
	(3.69)		(2.85)

F-test for joint significance of safe-haven effects: $F(10, 7295) = 7.51$.
P-Value: 0.00

Panel B
Hedonic house price regression

	Benchmark	IV spec.
Transaction volume	2.98***	23.23***
	(0.42)	(4.27)

Table 6
Cross-country panel analysis

The table reports the estimated coefficients ζ from panel regressions where the dependent variable is the price spread s_t^k 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:

$$s_t^k = \mu^k + \delta_t + \rho_1 s_{t-1}^k + \rho_2 s_{t-2}^k + \zeta z_{t-1}^k + u_t^k,$$

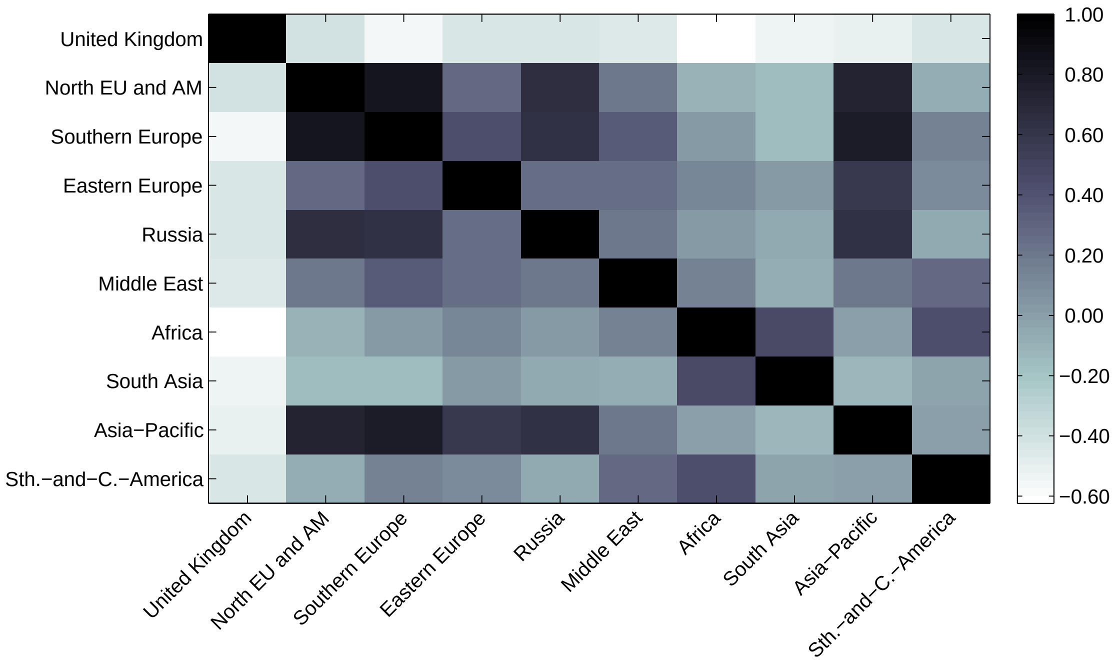
where z_t^k is an indicator variable which takes a value of one if the respective risk measure from country k is in the top 20% of its time series distribution. In parentheses, we report Driscoll-Kraay (1998) standard errors with a lag length of 12 months. *, **, *** denote statistical significance at the 10%, 5%, and 1% level respectively.

	ICRG index		**Yield spread vs. the UK**	**Policy uncertainty**	**ICRG index**
	(1)	(2)			Low-risk countries
Registry dataset	0.17***	0.11**	0.25**	0.42**	-0.05
	(0.06)	(0.04)	(0.11)	(0.20)	(0.09)
Loans dataset	0.27	0.02	0.70**	1.35***	-0.57***
	(0.23)	(0.13)	(0.33)	(0.50)	(0.20)
Country fixed effects	Yes	Yes	Yes	Yes	Yes
Time fixed effects	No	Yes	No	No	Yes

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. 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 3.

Panel A



Panel B

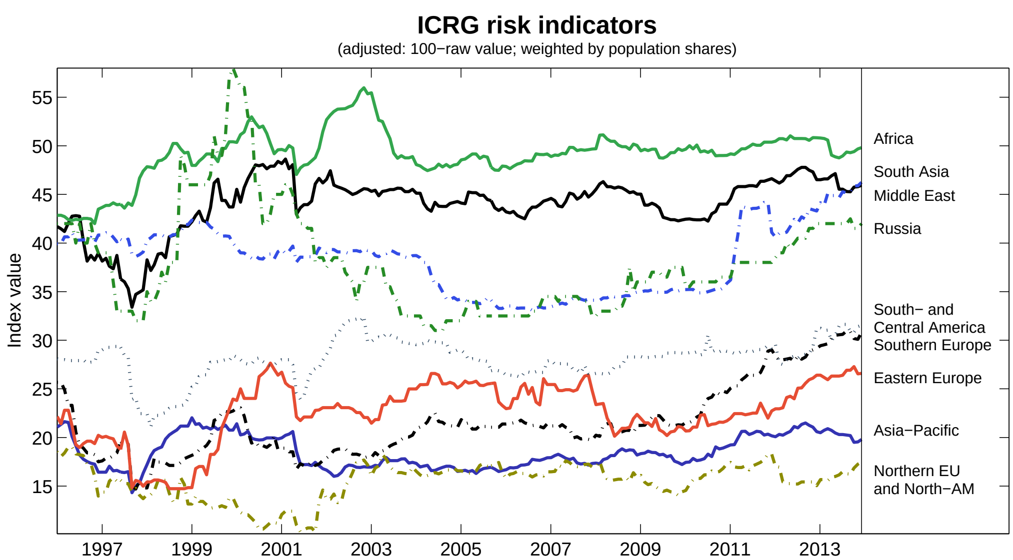
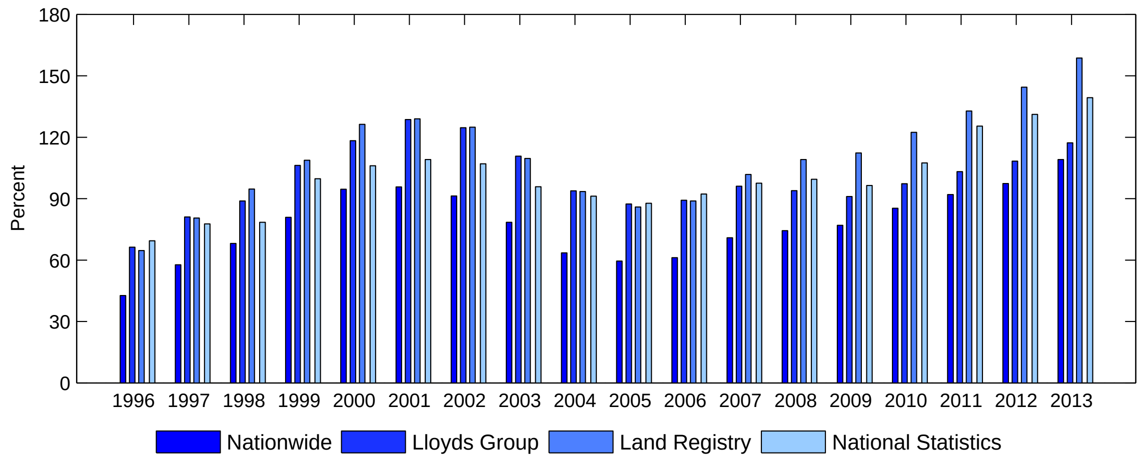


Figure 2
Evolution of the London housing market

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.

Panel A



Panel B

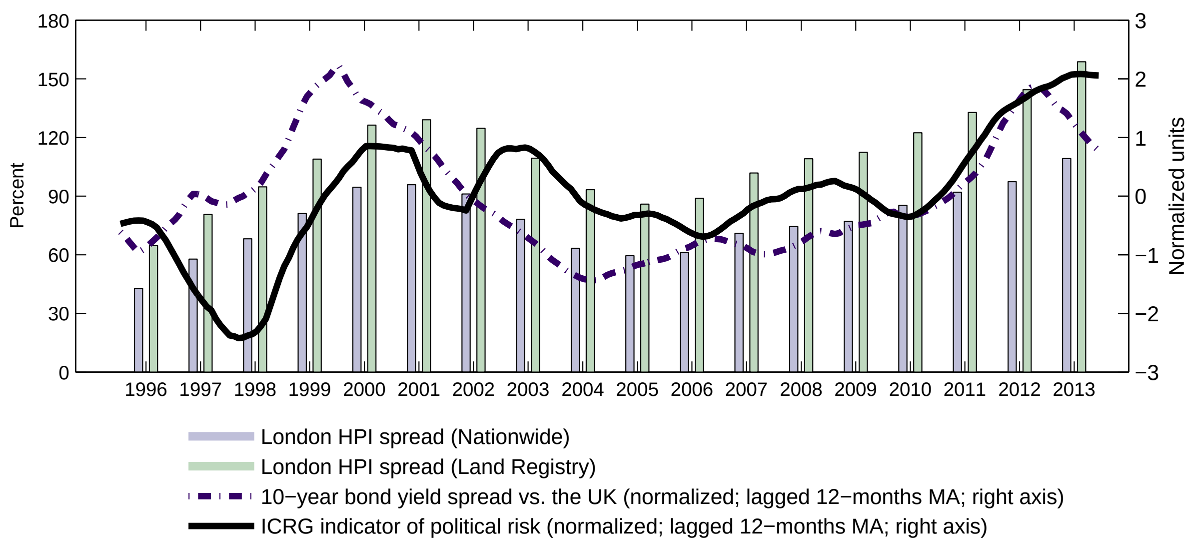


Figure 3
House prices, demographic structure and economic conditions at ward level

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.

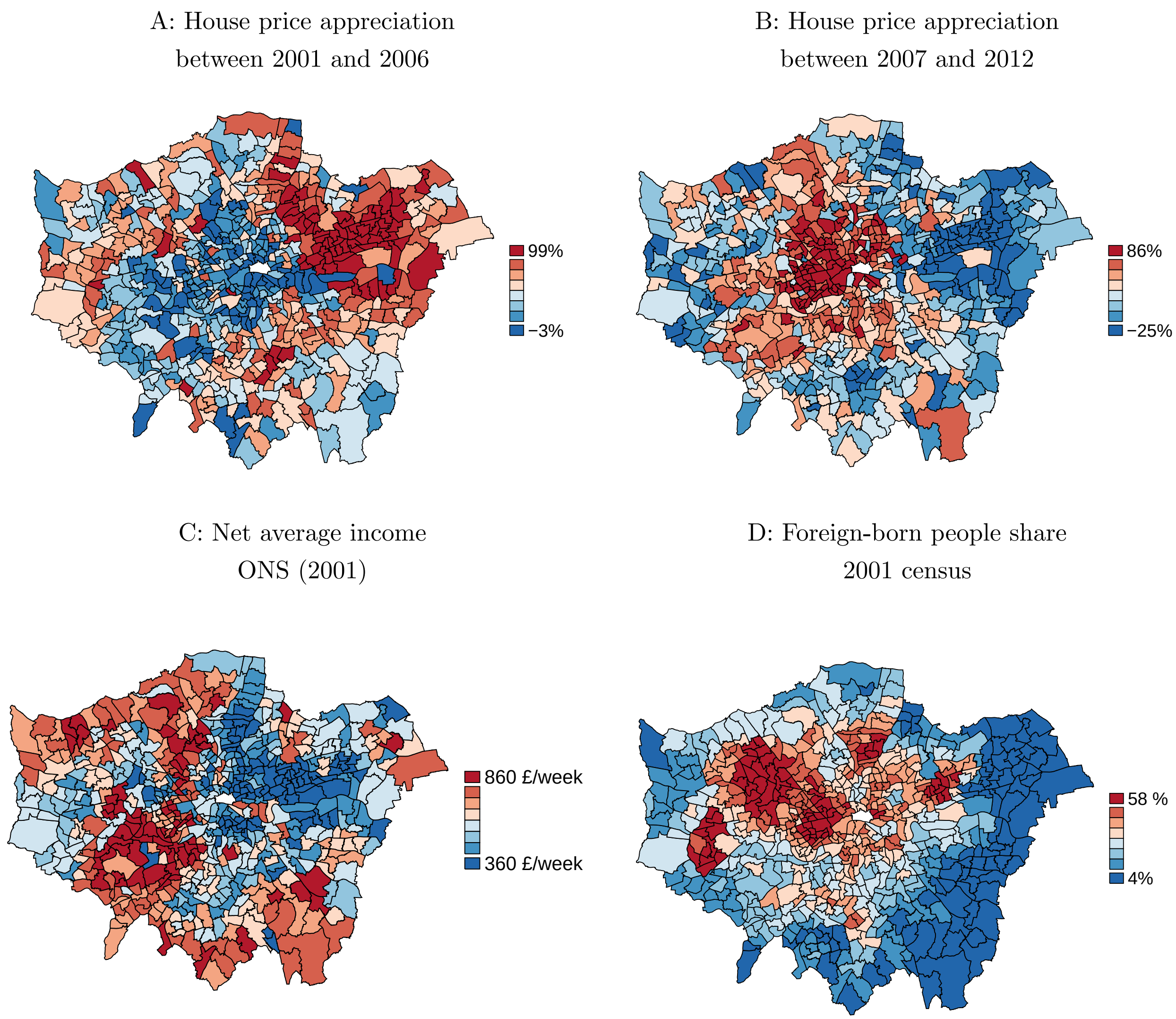


Figure 4
Safe haven effects across world regions

The figure reports the coefficients γ_0^k and γ_1 from the following hedonic regression:

$$\begin{aligned}\ln P_{i,t} &= \mu + \delta_t + \phi_w + \sum_{k \in K} \gamma_0^k f_w^k z_{t-1}^k + \gamma_1 y_w \overline{z}_{t-1} + \beta \mathbf{X}_{i,t} \\ &\quad + \rho_1 \ln \overline{P}_{w,t-1} + \rho_2 \ln \overline{P}_{w,t-2} + \nu_{i,t},\end{aligned}$$

where f_k^w are the shares of people in ward w born in world region k, y_w is average net income and $\overline{P}_{w,t}$ are average transaction prices in ward w in period t. In this specification, z_t^k is a top-quintile indicator of the ICRG index of political risk, corresponding to the nine world regions listed in Table 3. 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. The total length of the bars indicates point estimates and the shaded areas correspond to 95% confidence intervals. The standard errors are double-clustered at the borough and time level.

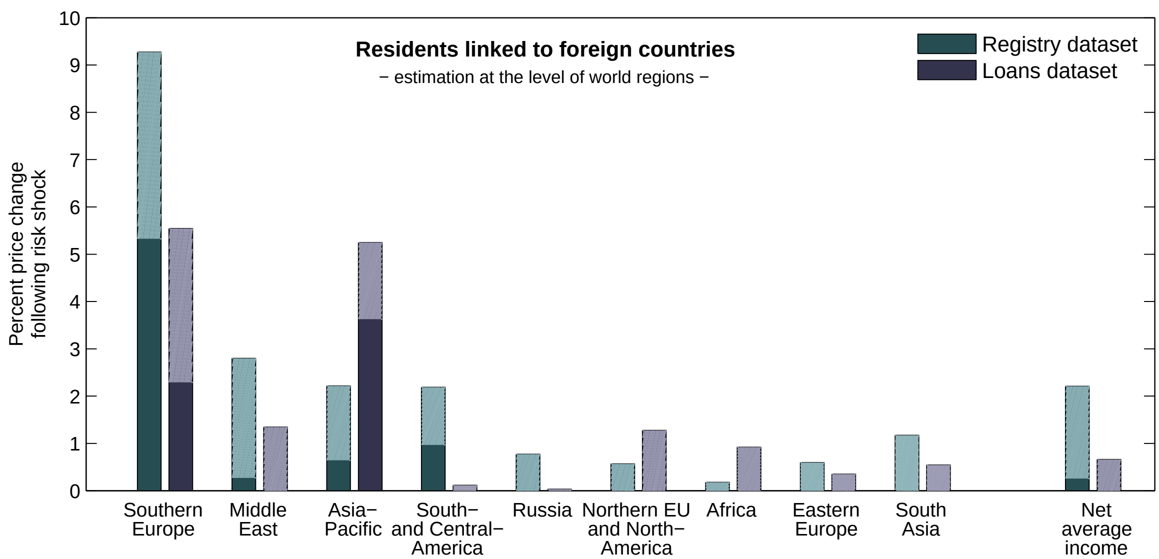


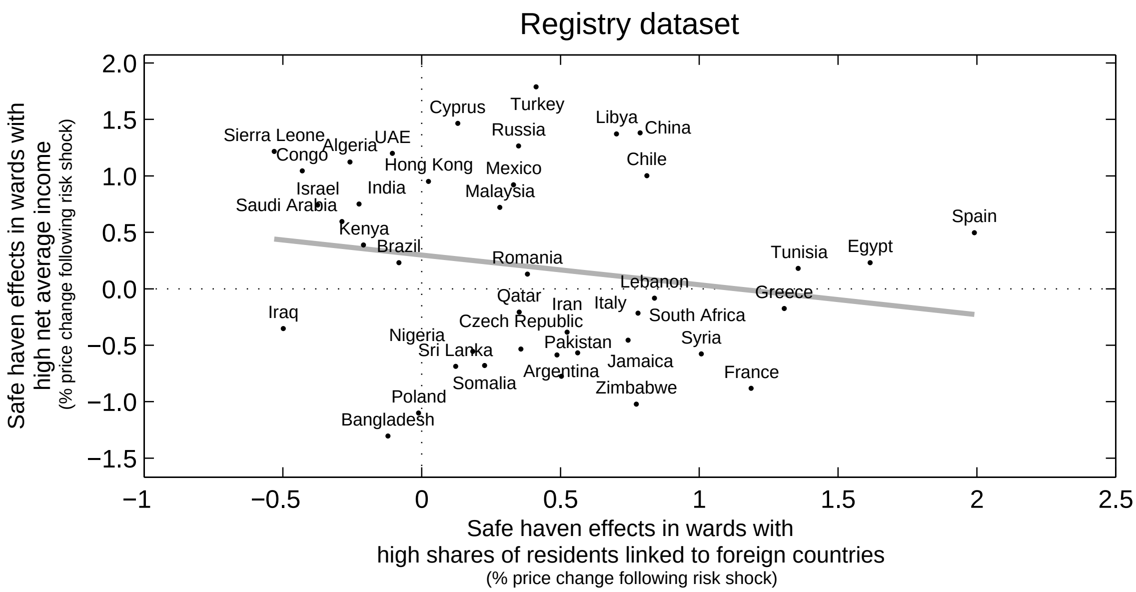
Figure 5

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 γ_0^k from country-by-country versions of equation (4):

$$\ln P_{i,t} = \mu + \delta_t + \phi_w + \left(\gamma_0^k f_w^k + \gamma_1^k y_w\right) z_{t-1}^k + \beta X_{i,t} + \rho_1 \ln \overline{P}_{w,t-1} + \rho_2 \ln \overline{P}_{w,t-2} + \nu_{i,t},$$

where f_w^k are the shares of people in ward w born in country k, y_w is average net income and z_t^k is the ICRG index of political risk. On the vertical axes, we report the corresponding point estimates of the coefficients γ_1^k, for each country k. z_t^k is normalized by subtracting the in-sample mean and dividing by the standard deviation. The coefficients are multiplied by a factor of 100, for easier interpretation as percentage points relative price appreciation. The lines indicate univariate cross-country fitted values.



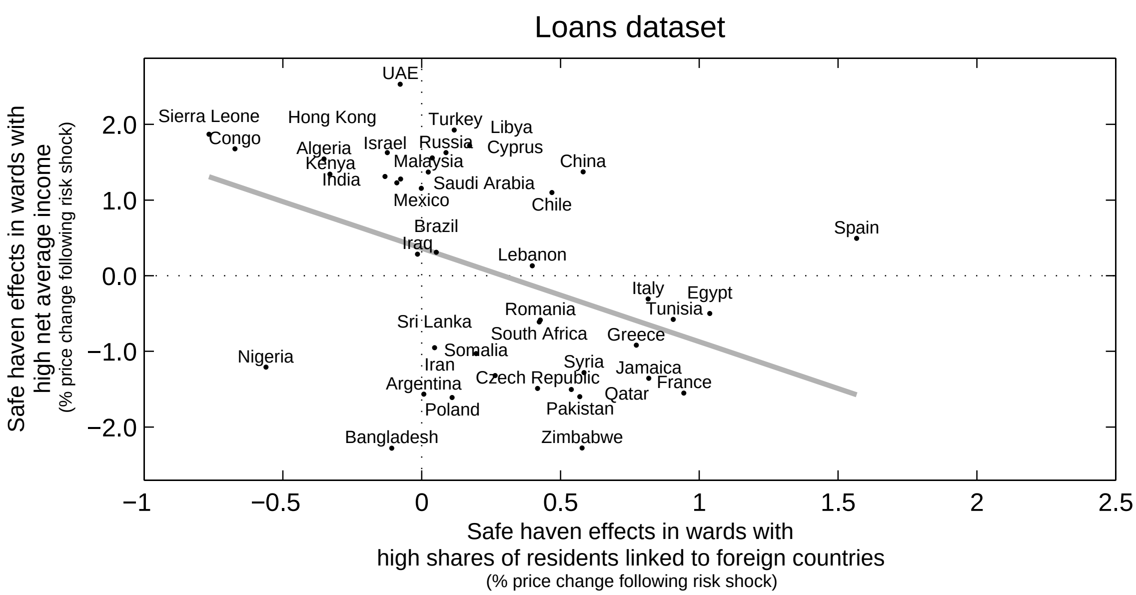


Figure 6

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:

$$s_t^k = \mu^k + \sum_{q=1}^{Q} \rho_q s_{t-q}^k + \sum_{q=1}^{Q} \zeta_q z_{t-q}^k + u_t^k, \quad \text{and} \quad z_t^k = \varphi^k + \sum_{q=1}^{Q} \pi_q z_{t-q}^k + \varepsilon_t^k,$$

where s_t^k 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_t^k 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.

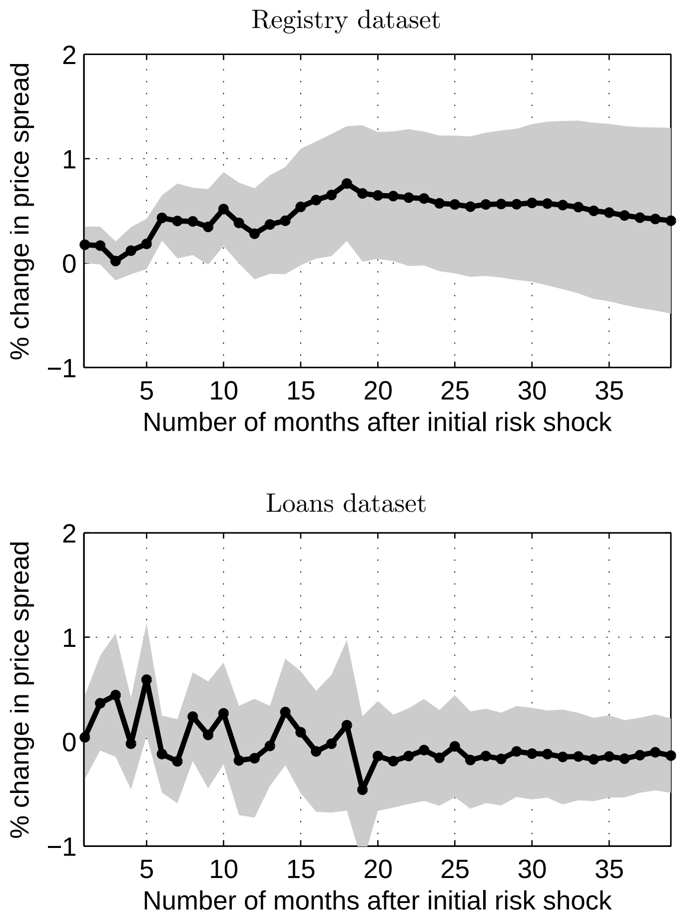


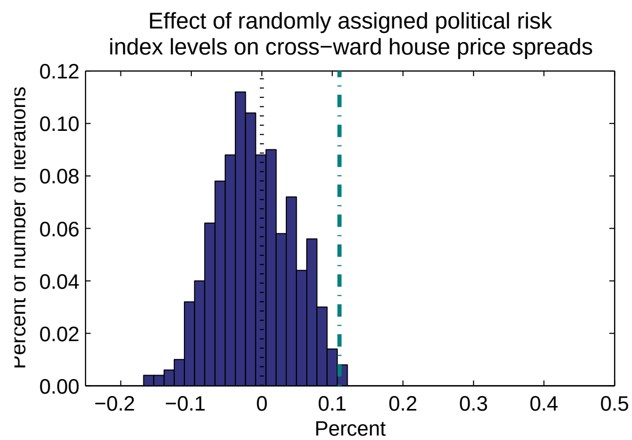
Figure 7
Placebo tests

The figure reports estimated distributions of the coefficients ζ and β from the following dynamic monthly panel regression specification:

$$s_t^k = \mu^k + \delta_t + \rho_1 s_{t-1}^k + \rho_2 s_{t-2}^k + \zeta z_{t-1}^k + \beta z_{t-1}^{\widetilde{k}} + u_t^k, \text{ with } \widetilde{k} \in K \text{ and } \widetilde{k} \neq k,$$

across a set of N=2,000 draws during which we pair each country k with a randomly drawn country $\widetilde{k}$ (with replacement). In each draw, we consider the political risk indicator in country $\widetilde{k}$ 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 ζ and β for each complete cross-country set of draws. In this specification, z_t^k is a top-quintile indicator of the ICRG index of political risk. In Panel A, we report the distribution of estimated coefficients β when restricting $\zeta = 0$. The estimation is carried out in the sample of high-risk countries. The dashed line indicates the estimated coefficient ζ, as reported in Table 6, in a specification which includes time fixed effects.

Panel A



Panel B

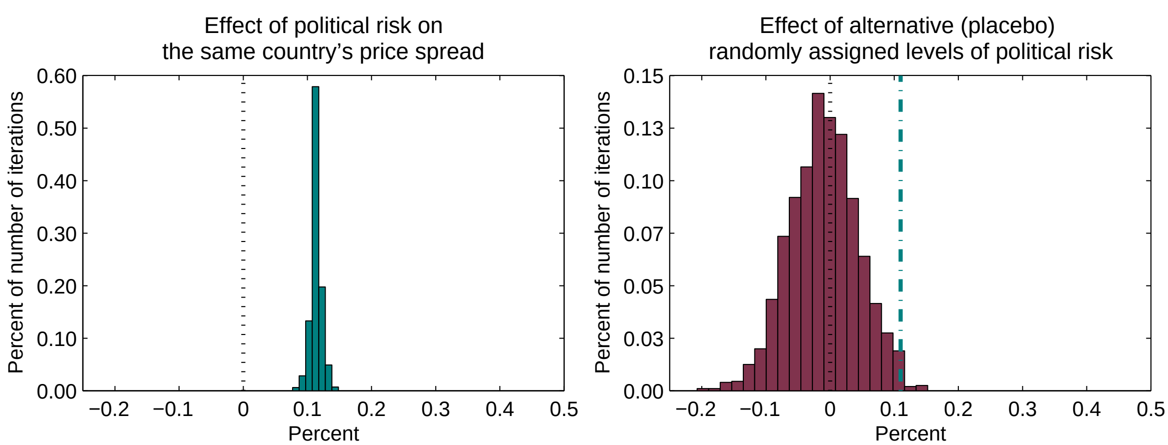
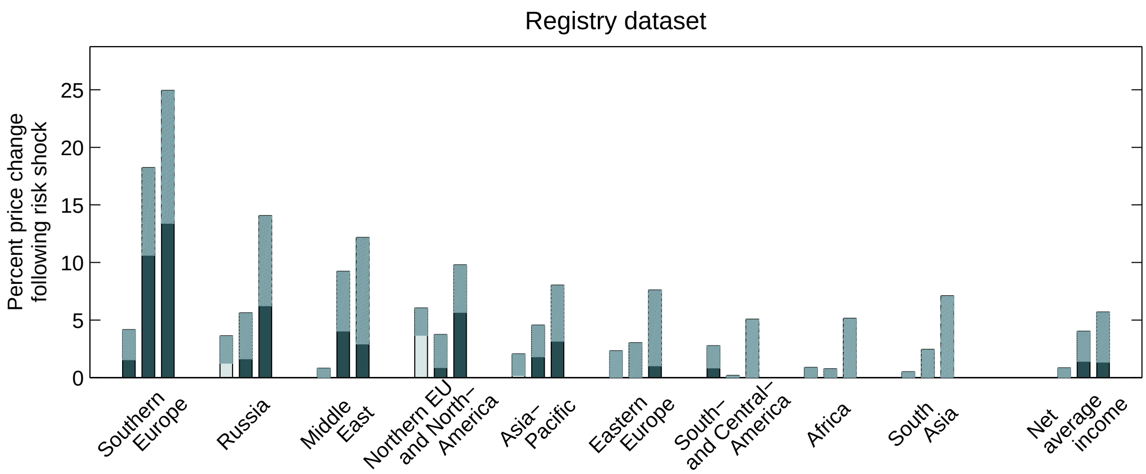


Figure 8
Safe haven effects across price categories

The figure reports the estimated coefficients $\gamma_0^{k,\eta}$ and γ_1^{η} from the following hedonic regression:

$$\begin{aligned}\ln P_{i,t} &= \mu + \delta_t + \phi_w + c^{\eta} + \sum_{\eta=1}^{3}\left(\sum_{k\in K}\gamma_0^{k,\eta} f_w^k z_{t-1}^k + \gamma_1^{\eta} y_w \overline{z}_{t-1}\right) + \beta \mathbf{X}_{i,t} \\ &\quad + \rho_1 \ln \overline{P}_{w,t-1} + \rho_2 \ln \overline{P}_{w,t-2} + \nu_{i,t},\end{aligned}$$

where f_w^k are the shares of people in ward w born in world region k and y_w is average net income. The parameter η 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 γ correspond thus to a triple interaction term between the ward-level characteristics f_w^k or y_w, the external factor z_t^k and price category dummies. In this specification, z_t^k is a top-quintile indicator of the ICRG index of political risk, corresponding to the nine world regions listed in Table 3. 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. The total length of the bars indicates point estimates and the shaded areas correspond to 95% confidence intervals. The standard errors are double-clustered at the borough and time level.



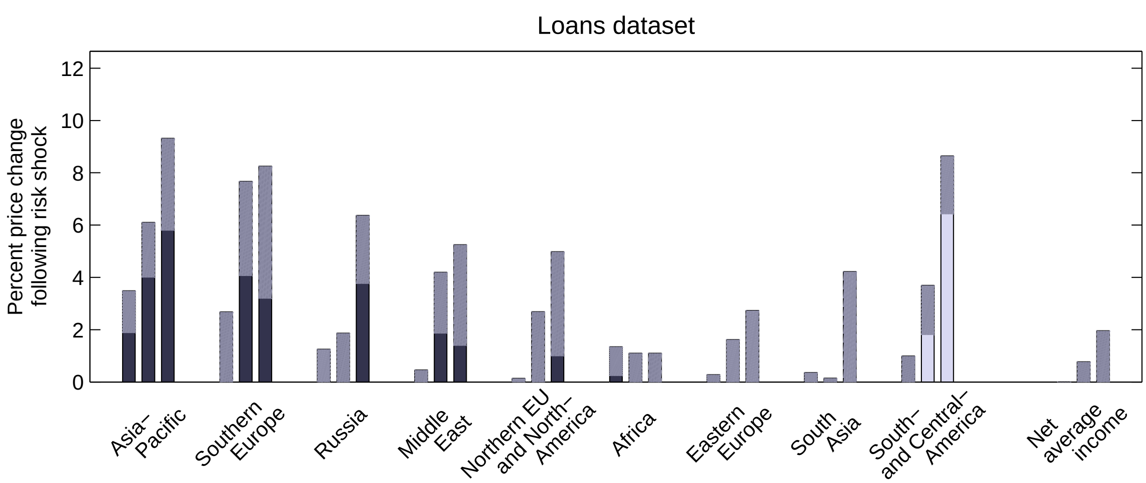


Figure 9
The evolution of foreign-born people shares through time

The figure reports the coefficients ρ^k from the regression:

$$\Delta f^k_{w,2011} = \alpha + \rho^k f^k_{w,2001} + \beta^k \Delta f^{UK}_{w,2011} + e^k_{w,2011},$$

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. The total length of the bars indicates point estimates and the shaded areas correspond to 95% confidence intervals. The estimated standard errors are White heteroskedasticity-robust.

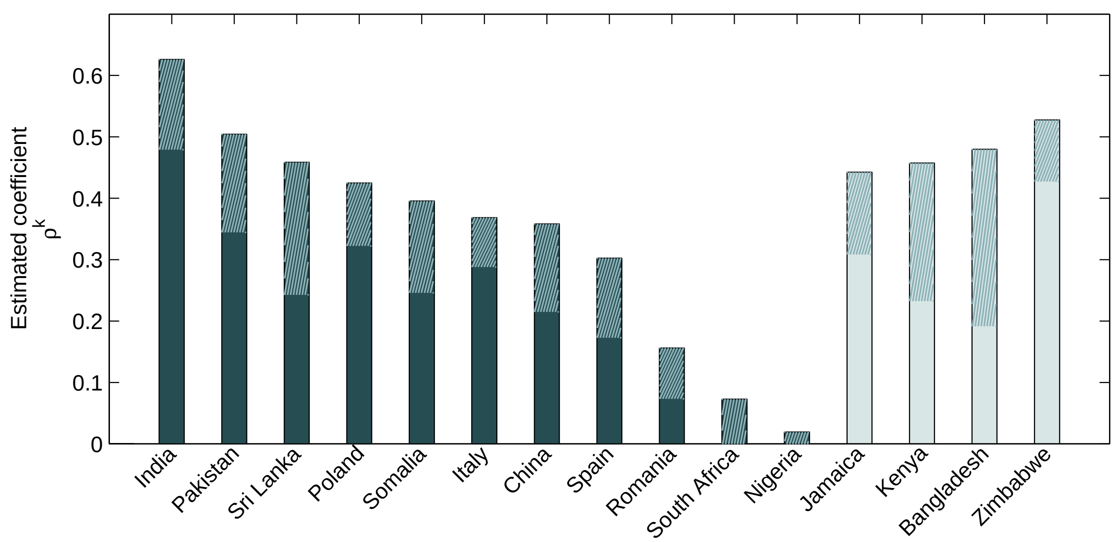
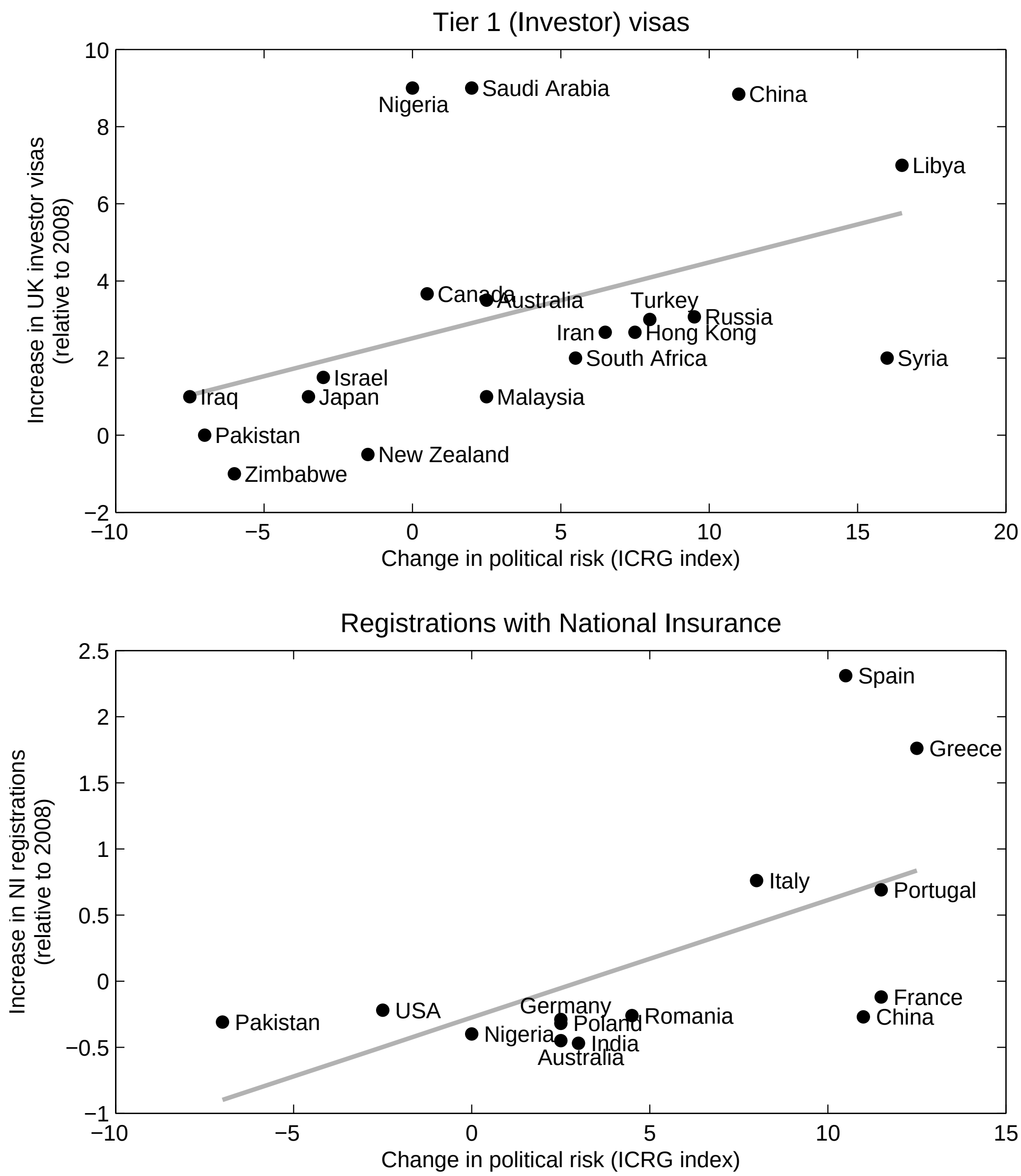


Figure 10

Foreign political risk, investment flows and migration into the UK

In the top panel, we report the number of additional Tier 1 (Investor) visas granted in 2013, relative to 2008. Tier 1 visas are granted to citizen from outside the European Economic Area and Switzerland which invest more than £2 million in the UK. In the bottom panel, we report the change in total migration into the UK for selected population groups, between 2008 and 2013. The lines indicate univariate cross-country fitted values. On the horizontal axis, we report the change in political risk (measured by the ICRG index) during the same period. In this representation, we exclude countries which are not in our main estimation sample, and countries for which the number of visas or the number of people which enter the UK are equal to zero.



Online Appendix for

Home Away From Home? Foreign Demand and London House Prices

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List of Tables

List of Figures

Table A.1
Data availability, risk indicators

The table reports the time periods for which risk and bond yield data is available, for each of the countries in our sample. A missing entry indicates that the respective series is not available or only covers a too short time span to be included in the analysis.

	ICRG index of political risk	10-year gov. bond yields	Economic policy uncertainty
Algeria	1995 - 2013	.	.
Argentina	1995 - 2013	.	.
Australia	1995 - 2013	1995 - 2013	.
Austria	1995 - 2013	1995 - 2013	.
Bangladesh	1995 - 2013	.	.
Belgium	1995 - 2013	1995 - 2013	.
Brazil	1995 - 2013	.	.
Canada	1995 - 2013	1995 - 2013	.
Chile	1995 - 2013	.	.
China	1995 - 2013	1995 - 2013	1995 - 2013
Congo	1995 - 2013	.	.
Cyprus	1995 - 2013	.	.
Czech Republic	1995 - 2013	2000 - 2013	.
Denmark	1995 - 2013	1995 - 2013	.
Egypt	1995 - 2013	.	.
Finland	1995 - 2013	1995 - 2013	.
France	1995 - 2013	1995 - 2013	.
Germany	1995 - 2013	1998 - 2013	.
Greece	1995 - 2013	1995 - 2013	.
Hong Kong	1995 - 2013	1995 - 2013	.
India	1995 - 2013	1996 - 2013	2003 - 2013
Iran	1995 - 2013	.	.
Iraq	1995 - 2013	.	.
Israel	1995 - 2013	1995 - 2013	.
Italy	1995 - 2013	1995 - 2013	1997 - 2013
Jamaica	1995 - 2013	.	.
Japan	1995 - 2013	1995 - 2013	.
Kenya	1995 - 2013	.	.

Table A.1
Data availability, risk indicators
(continued)

	ICRG index of political risk	10-year gov. bond yields	Economic policy uncertainty
Lebanon	1995 - 2013	.	.
Libya	1995 - 2013	.	.
Malaysia	1995 - 2013	1996 - 2013	.
Mexico	1995 - 2013	.	.
Netherlands	1995 - 2013	1995 - 2013	.
New Zealand	1995 - 2013	1995 - 2013	.
Nigeria	1995 - 2013	.	.
Pakistan	1995 - 2013	.	.
Poland	1995 - 2013	2001 - 2013	.
Portugal	1995 - 2013	1995 - 2013	.
Qatar	1995 - 2013	.	.
Romania	1995 - 2013	.	.
Russia	1995 - 2013	1999 - 2013	.
Saudi Arabia	1995 - 2013	.	.
Sierra Leone	1995 - 2013	.	.
Singapore	1995 - 2013	1998 - 2013	.
Somalia	1995 - 2013	.	.
South Africa	1995 - 2013	1995 - 2013	.
Spain	1995 - 2013	1995 - 2013	2001 - 2013
Sri Lanka	1995 - 2013	.	.
Sweden	1995 - 2013	1995 - 2013	.
Syria	1995 - 2013	.	.
Tunisia	1995 - 2013	.	.
Turkey	1995 - 2013	.	.
UAE	1995 - 2013	.	.
USA	1995 - 2013	1995 - 2013	.
Zimbabwe	1995 - 2013	.	.

Table A.2
Details of ONS variables

The ward-level values of the variables are obtained from the Office for National Statistics (ONS). This table reports the title descriptions and corresponding ONS data sources.

Variable	ONS variable description	Dataset
Shares of foreign-born people	Number of people (Count)	Country of Birth, 2001 (UV08)
Median age	Median age of population in the area (Years)	Age Structure, 2001 (KS02)
Population density	Density (Number of Persons per Hectare)	Population Density, 2001 (UV02)
Net average income	Average weekly household net income (Pounds Sterling)	Income: Model Based Estimates, 2001
Higher prof. occupations	People aged 16-74: Higher professional occupations (Percentage)	Socioeconomic Classification - All People, 2001 (KS14A)
Percent of detached houses	In an unshared dwelling: House or Bungalow: Detached (Persons)	Accommodation Type - People, 2001 (UV42)
Percent of flats	In an unshared dwelling: Flat, maisonette or apartment (Persons)	Accommodation Type - People, 2001 (UV42)
Long term unemployed	People aged 16-74: Long-term unemployed (Percentage)	Socioeconomic Classification - All People, 2001 (KS14A)
Mortgage ownership	Owned (Households, Count)	Tenure - Households, 2001 (UV63)
	Owned: Owns with a mortgage or loan (Households, Count)	Tenure - Households, 2001 (UV63)
Cars per household	All Households (Count)	Cars or Vans, 2001 (KS17)
	All cars or vans in the area (Vehicles)	Cars or Vans, 2001 (KS17)

Table A.3

Summary statistics across wards

The table reports mean values for selected variables, calculated for the wards in the top quintile of the respective distributions, according to the share of people born in our set of country regions. The population density is calculated using the usual resident population and the size of the area in hectares. The market share of flats indicates all people who were usually resident in the area at the time of the 2001 census, who lived in an unshared dwelling, that was a flat, maisonette or apartment, as a percent of the total ward population. Net average income levels are estimated by the UK Office for National Statistics and expressed in pounds sterling per week. The information on vehicle ownership is based on the number of cars or vans owned, or available for use, by one or more members of a household, including company cars or vans available for private use. The share of people in higher professional occupations is reported as classified by the UK Office for National Statistics. The ward-level degree of mortgage ownership is given by the number of households in the area at the time of the 2001 census, who are holders of a residential mortgage, as a fraction of the total number of homeowners.

		Population density *(no/ha)*	Market share of flats *(percent)*	Net income *(£/week)*	Cars per household *(no/hh.)*	Higher prof. occupations *(percent)*	Mortgage holders *(percent)*
Top 20% of wards with highest shares of people born in:	Nothern Europe and North America	96.23	62.09	640.89	0.74	12.64	54.47
	Southern Europe	110.99	68.51	580.80	0.63	10.86	55.71
	Eastern Europe	85.56	50.59	603.28	0.81	10.60	55.41
	Russia	94.08	59.84	590.08	0.73	10.85	54.81
	Middle East	90.17	54.38	537.10	0.73	8.51	57.16
	Africa	77.09	40.36	490.24	0.82	6.27	61.35
	South Asia	76.17	34.04	497.10	0.86	6.22	61.24
	Asia-Pacific	97.42	62.38	641.28	0.73	12.74	54.92
	South and Central America	89.76	48.87	484.24	0.68	6.52	63.81
	UK	38.04	14.95	553.20	1.15	4.92	58.86
Full sample of wards		70.68	39.41	546.14	0.88	7.62	59.88

Table A.4
Safe haven effects: Global risk and total foreign population shares

The table reports estimated coefficients γ_0 from the following hedonic regression:

$$\ln P_{i,t} = \mu + \delta_t + \phi_w + \gamma_0 \overline{f}_w \overline{z}_{t-1} + \beta \boldsymbol{X}_{i,t} + \rho_1 \ln \overline{P}_{w,t-1} + \rho_2 \ln \overline{P}_{w,t-2} + \nu_{i,t}.$$

where $\boldsymbol{X}_{i,t}$ are property-level characteristics. In this specification, $\overline{f}_w$ is the total ward-level share of people born outside the UK. The explanatory variables are normalized by subtracting the in-sample mean and dividing by the standard deviation. The estimated coefficients are multiplied by 100, for easier interpretation as percentage points. *, **, *** denote statistical significance at the 10%, 5%, and 1% level respectively. In parentheses, we report standard errors, double-clustered at the borough and time level.

		Registry dataset	Loans dataset
Residents linked to foreign countries	γ_0	1.25***	1.53***
		(0.18)	(0.21)

Table A.5
Cross-country panel analysis

The table reports estimated coefficients γ_0 and γ_1 from the following cross-country specification:

$$\begin{aligned} s_t^k &= \rho_1 s_{t-1}^k + \rho_2 s_{t-2}^k + (\overline{\phi}^{high,k} - \overline{\phi}^{low,k}) \\ &+\gamma_0 \left[(\overline{f}^{high,k} - \overline{f}^{low,k}) z_{t-1}^k + \sum_{k' \in K, k' \neq k} \left(\overline{f}^{k'|high,k} - \overline{f}^{k'|low,k} \right) z_{t-1}^{k'} \right] \\ &+\gamma_1 \left(\overline{y}^{high,k} - \overline{y}^{low,k} \right) \overline{z}_{t-1} + u_t^k. \end{aligned}$$

Here, the dependent variable is the price spread s_t^k in period t between the top and bottom 20% of wards with respect to the share of people born in country k,and z_t^k 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. $\overline{f}^{high,k}$ is the average share of people born in country k, $f^{k'|high,k}$ is the average share of people born in country k' and $\overline{y}^{high,k}$ is the average income level across these wards. In parentheses, we report Driscoll-Kraay (1998) standard errors with a lag length of 12 months.

	(1)	(2)
Residents linked with foreign countries		
- total effect of foreign demand	0.09***	0.13***
	(0.03)	(0.04)
Net average income	-1.51**	-1.16*
	(0.75)	(0.64)
Time fixed effects	No	Yes

Table A.6
Capital flows: Summary statistics

The table reports selected summary statistics characterizing the distribution of capital flows in our sample of countries.

Capital outflows by domestic agents (*mil. USD*)				Capital outflows by domestic agents (*in percent, relative to GDP*)			
	Mean	Min	Max		Mean	Min	Max
United States	505,564	96	1,472,128	Belgium-Lux.	94.94	12.33	241.82
Belgium-Lux.	315,711	31,285	1,030,690	Singapore	51.65	22.90	95.44
Germany	311,253	59,090	830,418	Cyprus	35.89	5.31	151.97
France	271,536	25,795	717,797	Hong Kong	35.72	-85.94	159.94
Japan	214,055	-13,407	494,244	Netherlands	25.86	-4.50	47.52
China	168,939	8,880	632,496	Libya	18.90	-5.94	57.19
Netherlands	134,592	-37,362	334,206	Austria	15.85	-7.32	51.14
Italy	102,735	-14,927	248,119	Sweden	15.80	1.27	33.01
Spain	94,482	-3,573	212,138	Finland	14.61	3.23	35.04
Russia	70,761	10,518	264,898	Denmark	14.16	-5.72	32.04
Hong Kong	69,842	-135,126	316,806	France	14.12	1.95	29.47
Canada	69,731	20,815	160,794	Portugal	12.23	3.14	19.11
Singapore	54,970	19,016	145,279	Germany	12.17	2.93	26.61
Sweden	50,425	2,936	134,238	Spain	10.72	-0.68	18.83
Austria	44,540	-28,759	155,195	Malaysia	9.82	-5.34	26.32
Denmark	32,480	-8,650	92,825	Russia	9.26	2.51	24.01
Saudi Arabia	30,662	-13,584	146,950	Chile	8.69	-0.07	18.40
Australia	30,385	-2,226	94,379	Saudi Arabia	8.61	-10.84	37.67
Finland	24,550	3,782	49,475	Czech Republic	7.79	1.28	23.40
Brazil	22,829	-9,158	112,816	China	7.77	1.81	18.60
India	18,547	-659	90,520	Jamaica	7.73	-0.49	18.66
Portugal	18,118	7,017	30,986	Israel	7.64	2.69	21.88
Greece	13,216	-5,495	49,472	Canada	7.53	3.72	12.81
Malaysia	12,463	-3,664	45,600	Italy	6.97	-0.71	14.22
Israel	10,496	1,957	34,377	Greece	5.45	-4.32	16.91
Mexico	9,988	-12,728	40,296	Australia	5.15	-0.33	12.75
Chile	9,550	-32	26,410	United States	4.82	0.00	10.99
Libya	9,317	-1,647	38,079	Japan	4.81	-0.30	10.66
Argentina	9,281	-10,214	26,313	Argentina	3.84	-4.44	10.68
Turkey	8,577	-1,862	24,492	Tunisia	3.84	1.43	6.52
Poland	6,689	-6,638	26,819	South Africa	3.83	-0.76	9.67
South Africa	6,685	-2,095	18,506	Egypt	3.62	-2.74	11.52
Cyprus	6,574	407	33,707	Romania	2.84	-3.00	7.58
Czech Republic	6,018	1,305	14,468	Congo	2.79	-5.00	17.33
Egypt	3,692	-2,164	14,203	New Zealand	2.78	-4.45	8.30
Romania	2,772	-1,062	9,008	India	2.62	-0.20	8.87
New Zealand	2,035	-5,491	9,617	Poland	2.50	-2.01	6.87
Iran	1,449	-5,990	9,340	Turkey	2.47	-1.18	4.74
Tunisia	918	211	2,093	Brazil	2.28	-1.45	9.31
Jamaica	778	-35	2,026	Mexico	1.44	-3.66	4.56
Pakistan	754	-5,554	4,617	Iran	1.36	-6.51	8.25
Congo	173	-177	1,205	Pakistan	1.06	-3.71	5.37
Syria	-55	-1,073	1,290	Syria	-0.64	-7.45	10.53
Full sample	65,569	-135,126	1,472,128	Full sample	11.85	-85.94	241.82

Figure A.1
Foreign-born people shares in London

.
The figure reports the overall shares of foreign-born people in London. We use these shares in order to construct weighted averages of variables.

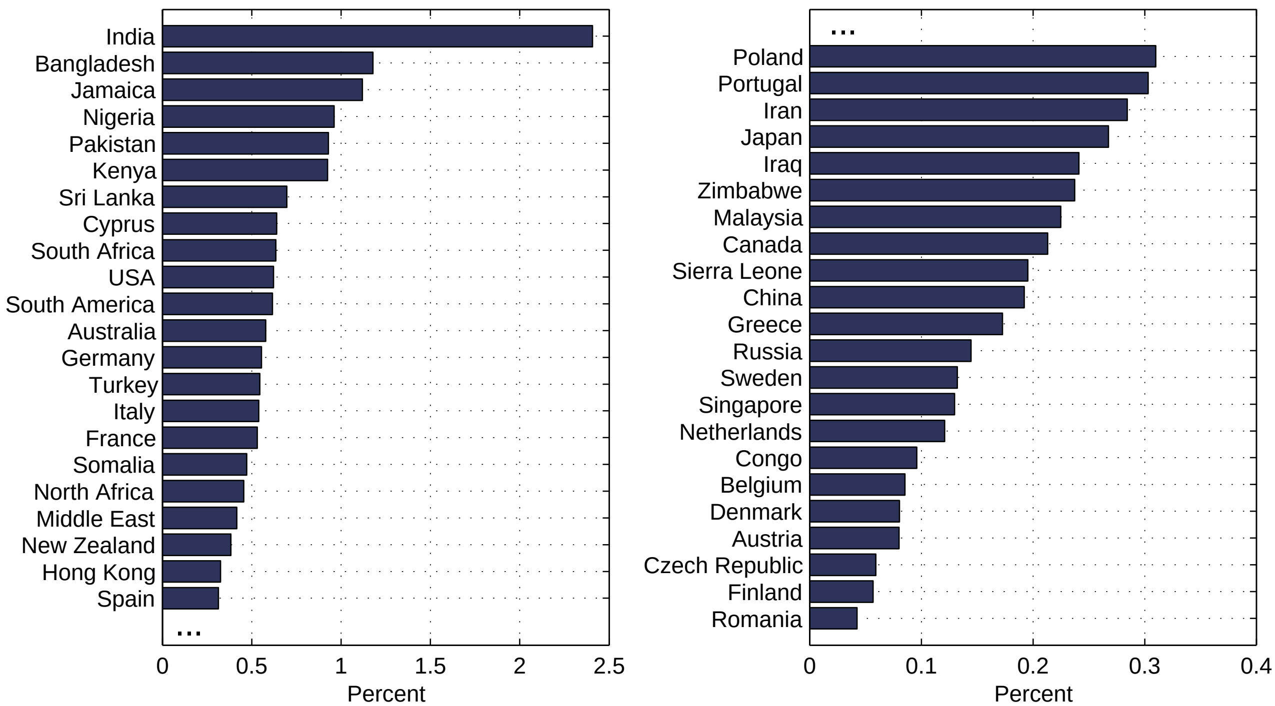


Figure A.2
Average levels of political risk

The figure reports average levels of political risk, as captured by the ICRG indexes. In our estimation, we distinguish between countries with high political risk (higher than a threshold of 20) and low political risk (lower than 20). We adjust the raw index series reported by the PRS Group by subtracting them from a total value of 100. This insures that we can interpret higher index values as increases in risk.

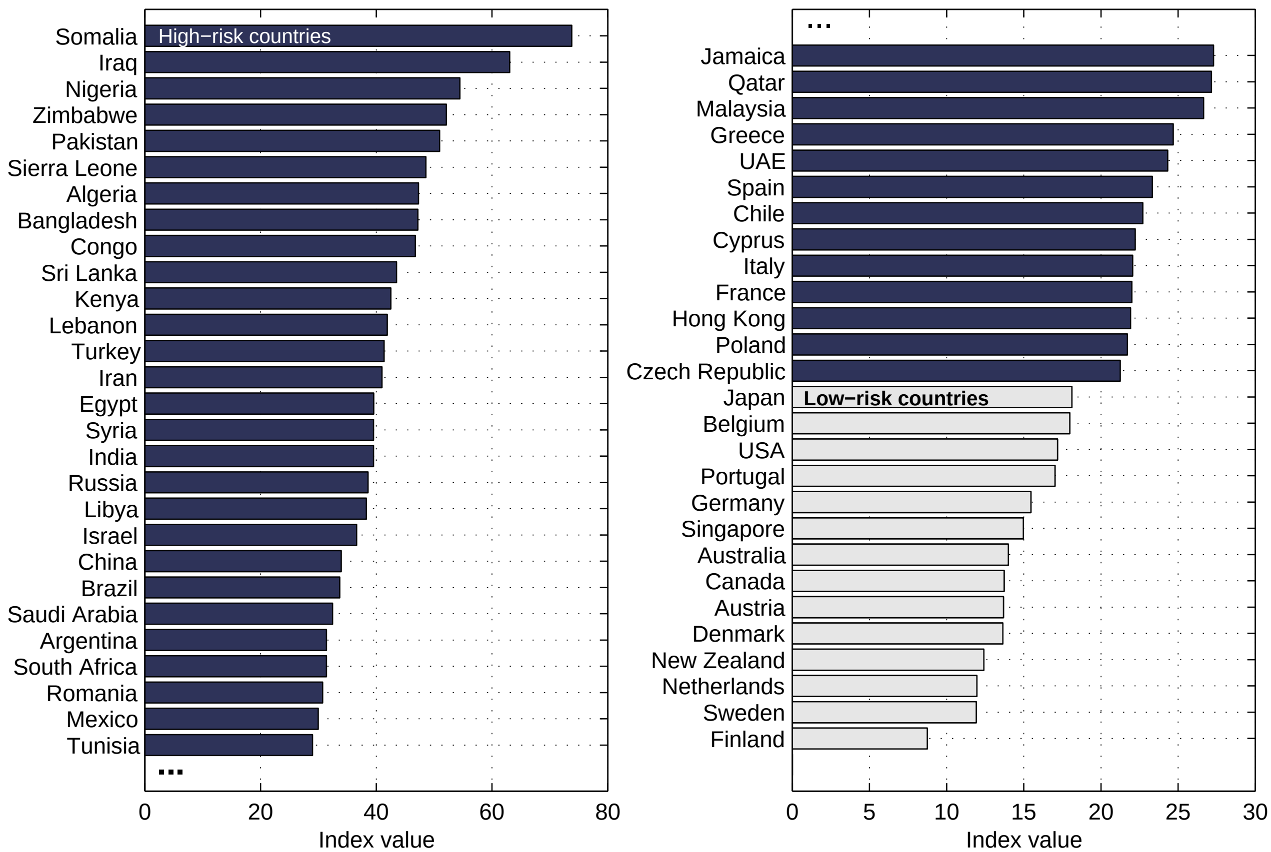


Figure A.3
Cross-ward distribution of demographic, social and economic variables

The figure shows the distribution of selected variables, across the set of 624 London wards. We report the unit of measurement in parentheses, below the variable name.

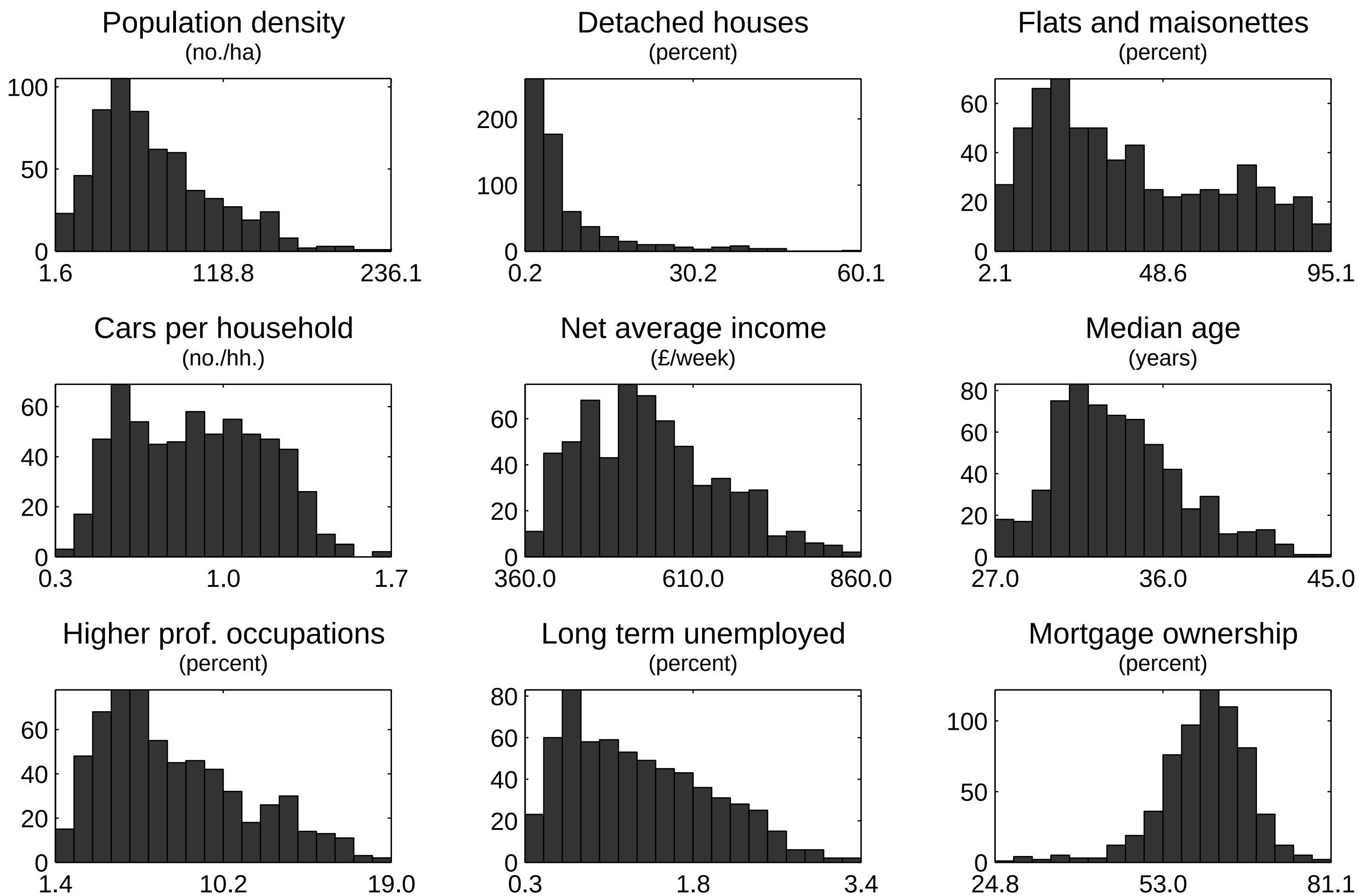


Figure A.4

Cross-ward distribution of the foreign-born people shares

The figure shows the distribution of the shares of people born in respective countries or country groups, across the set of 624 London wards. We report the shares in percent of the total ward population.

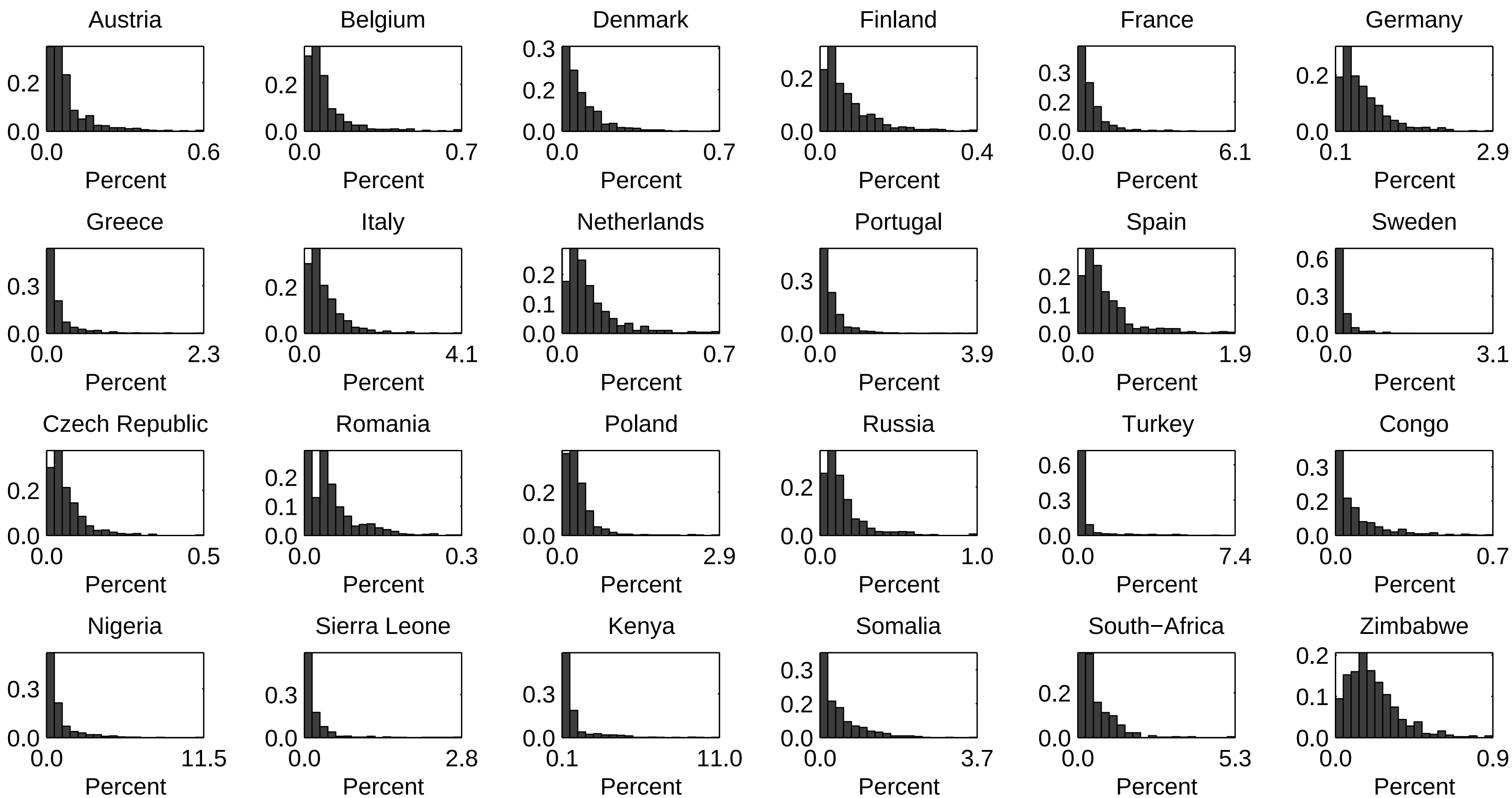


FIGURE A.4

Cross-ward distribution of the foreign-born people shares

(continued)

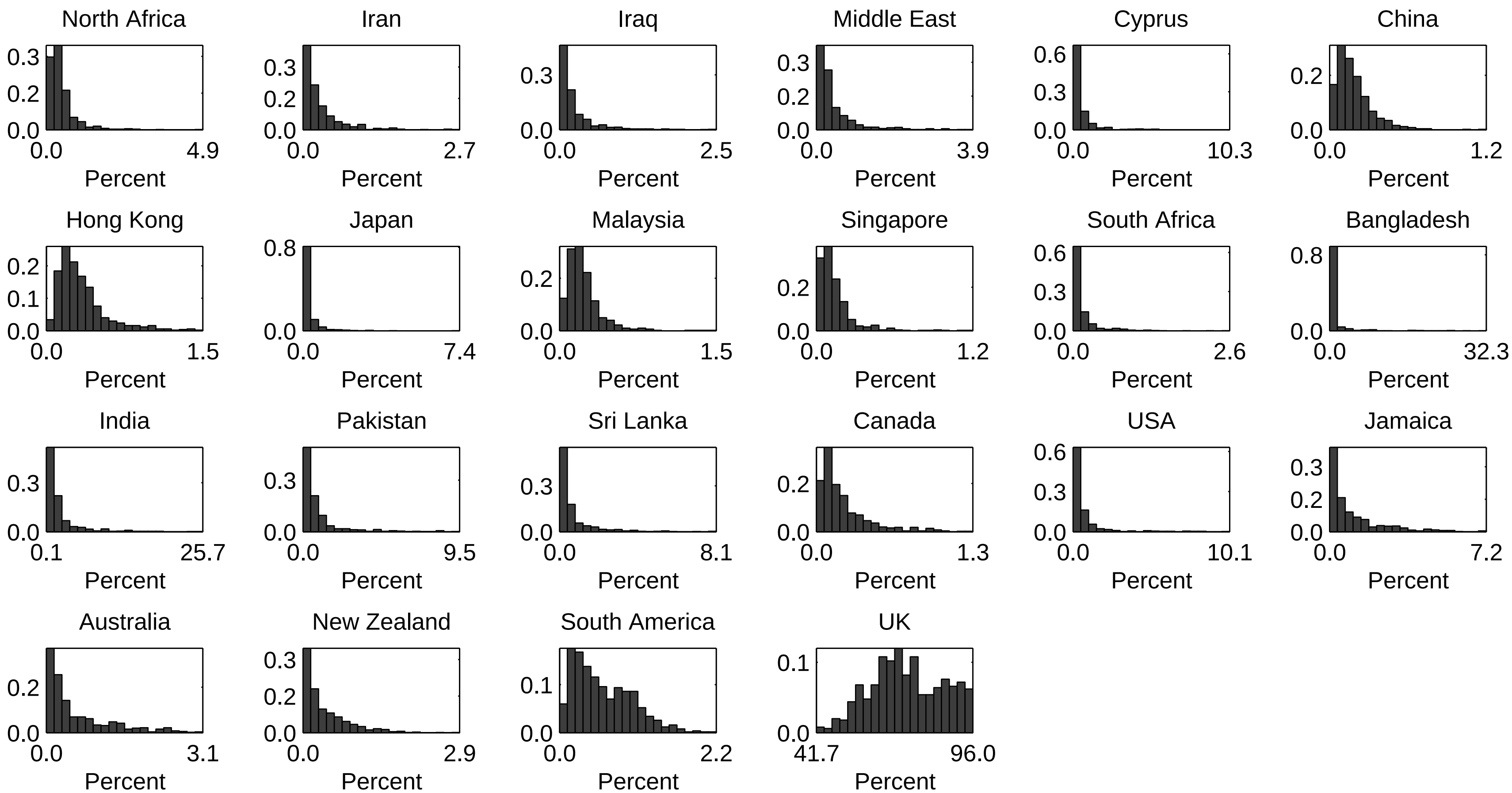


Figure A.5
Yield data: China

The figure compares the yield on 10-year Chinese government bonds from two data sources. We use the first, retrieved through Datastream, in our main analysis. The second is obtained from Global Financial Data (identified through the symbol IGCHN10D) and only available for a shorter time period.

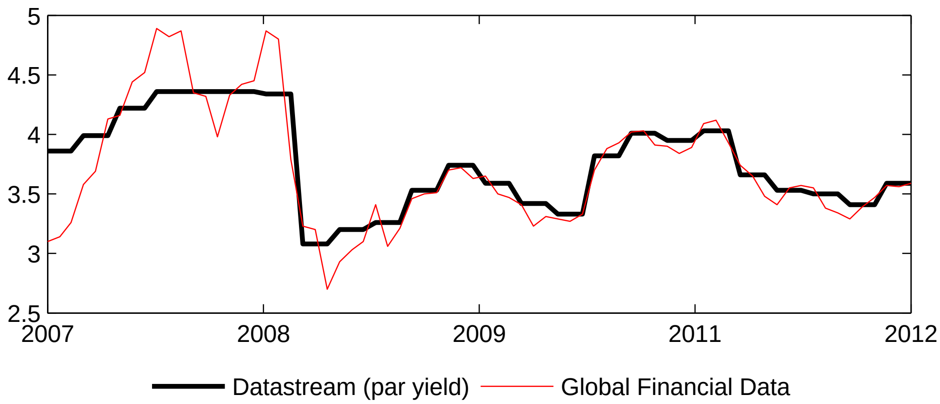


Figure A.6

The price impact of foreign demand: effects across risk deciles

The figure reports the estimated coefficients ζ_d from the following dynamic monthly panel regression specification:

$$s_t^k = \mu^k + \rho_1 s_{t-1}^k + \rho_2 s_{t-2}^k + \sum_{d=1}^{10} \zeta_d decile_d(z_{t-1}^k) + \varepsilon_t^k,$$

where s_t^k 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. The estimation is carried out in the sample of high-risk countries. Statistical significance is reported through error bars, indicating 90% confidence intervals based on Driscoll-Kraay (1998) standard errors with a lag length of 12 months.

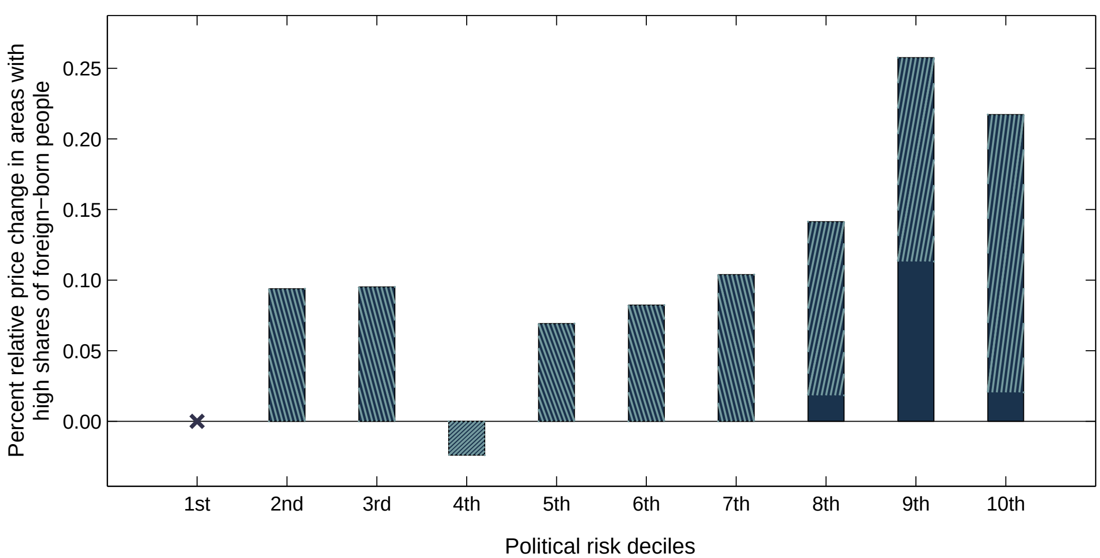


Figure A.7

Within-borough identification of safe haven effects

The figure reports the coefficients γ_0^k and γ_1 from the following hedonic regression:

$$\begin{aligned}\ln P_{i,t} \quad = \quad & \mu + \delta_t + \phi_w + \tau_{b,q} + \sum_{k \in K} \gamma_0^k f_w^k z_{t-1}^k + \gamma_1 y_w \overline{z}_{t-1} + \beta \mathbf{X}_{i,t} \\ & + \rho_1 \ln \overline{P}_{w,t-1} + \rho_2 \ln \overline{P}_{w,t-2} + \nu_{i,t}.\end{aligned}$$

where f_k^w are the shares of people in ward w born in world region k, y_w is average net income and $\overline{P}_{w,t}$ are average transaction prices in ward w in period t. In this specification, z_t^k is a top-quintile indicator of the ICRG index of political risk. $\tau_{b,q}$ are borough-year fixed effects, for borough b in year q. 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. The total length of the bars indicates point estimates and the shaded areas correspond to 95% confidence intervals. The standard errors are double-clustered at the borough and time level.

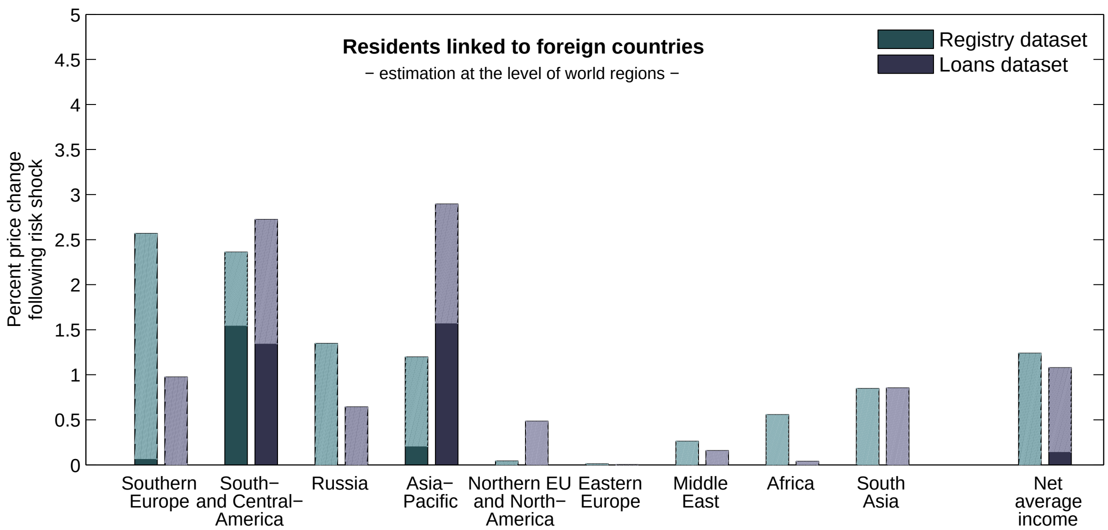


FIGURE A.8

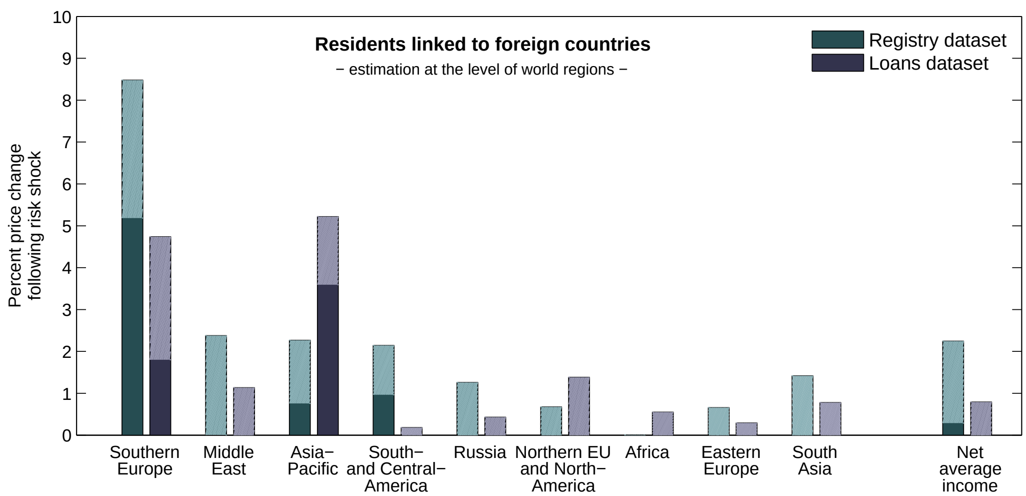
Robustness check: The role of aggregate factors

The figure reports the coefficients γ_0^k and γ_1 from the following regressions:

$$\begin{aligned}\ln P_{i,t} &= \mu + \delta_t + \phi_w + r_t^{UK}\overline{f}_w + \sum_{k\in K}\gamma_0^k f_w^k z_{t-1}^k + \gamma_1 y_w \overline{z}_{t-1} + \beta \mathbf{X}_{i,t} \\ &\quad + \rho_1 \ln \overline{P}_{w,t-1} + \rho_2 \ln \overline{P}_{w,t-2} + \nu_{i,t}, \quad \text{(Panel A)} \\ \ln P_{i,t} &= \mu + \delta_t + \phi_w + \tau_t\overline{f}_w + \sum_{k\in K}\gamma_0^k f_w^k z_{t-1}^k + \gamma_1 y_w \overline{z}_{t-1} + \beta \mathbf{X}_{i,t} \\ &\quad + \rho_1 \ln \overline{P}_{w,t-1} + \rho_2 \ln \overline{P}_{w,t-2} + \nu_{i,t}, \quad \text{(Panel B)}\end{aligned}$$

where f_k^w are the shares of people in ward w born in world region k, y_w is average net income and $\overline{P}_{w,t}$ are average transaction prices in ward w in period t. In this specification, z_t^k is a top-quintile indicator of the ICRG index of political risk and $\overline{f}_w$ is a top-quintile indicator of the total ward-level share of people born outside the UK. r_t^{UK} is the Bank of England policy rate. 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. The total length of the bars indicates point estimates and the shaded areas correspond to 95% confidence intervals. The standard errors are double-clustered at the borough and time level.

Panel A



Panel B

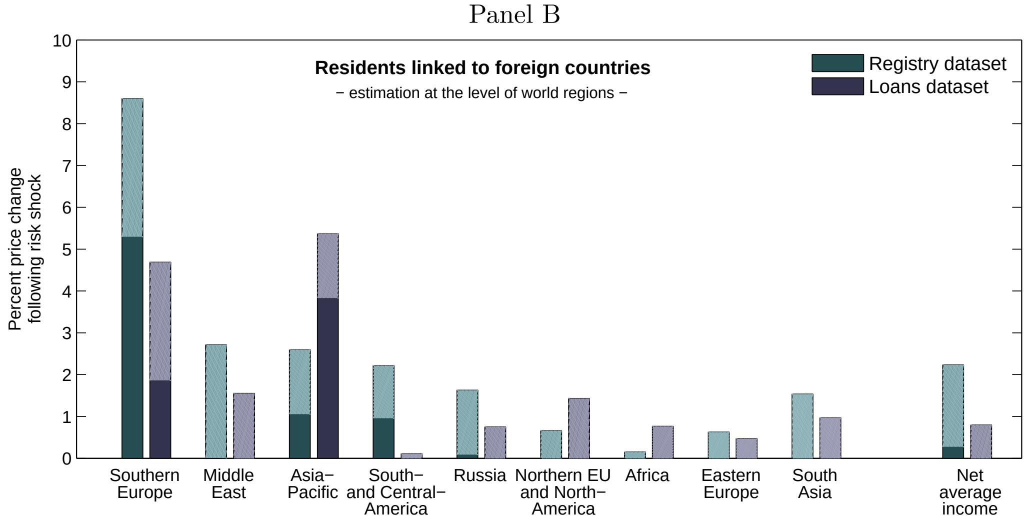


Figure A.9
Robustness check: Hedonic characteristics

The table reports the coefficients γ_0^k and γ_1 from the following hedonic regression:

$$\begin{aligned}\ln P_{i,t} \quad = \quad & \mu + \delta_t + \phi_w + \sum_{k \in K} \gamma_0^k f_w^k z_{t-1}^k + \gamma_1 y_w \overline{z}_{t-1} + \beta \mathbf{X}_{i,t} + \rho_1 \ln \overline{P}_{w,t-1} \\ & + \rho_2 \ln \overline{P}_{w,t-2} + b_1 \overline{\mathbf{X}}_{w,t-1} + b_2 \overline{\mathbf{X}}_{w,t-2} + \nu_{i,t}.\end{aligned}$$

where f_k^w are the shares of people in ward w born in world region k, y_w is average net income and $\overline{P}_{w,t}$ are average transaction prices in ward w in period t. In this specification, z_t^k is a top-quintile indicator of the ICRG index of political risk. $\overline{\mathbf{X}}_{w,t}$ are average hedonic characteristics at ward level. 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. The total length of the bars indicates point estimates and the shaded areas correspond to 95% confidence intervals. The standard errors are double-clustered at the borough and time level.

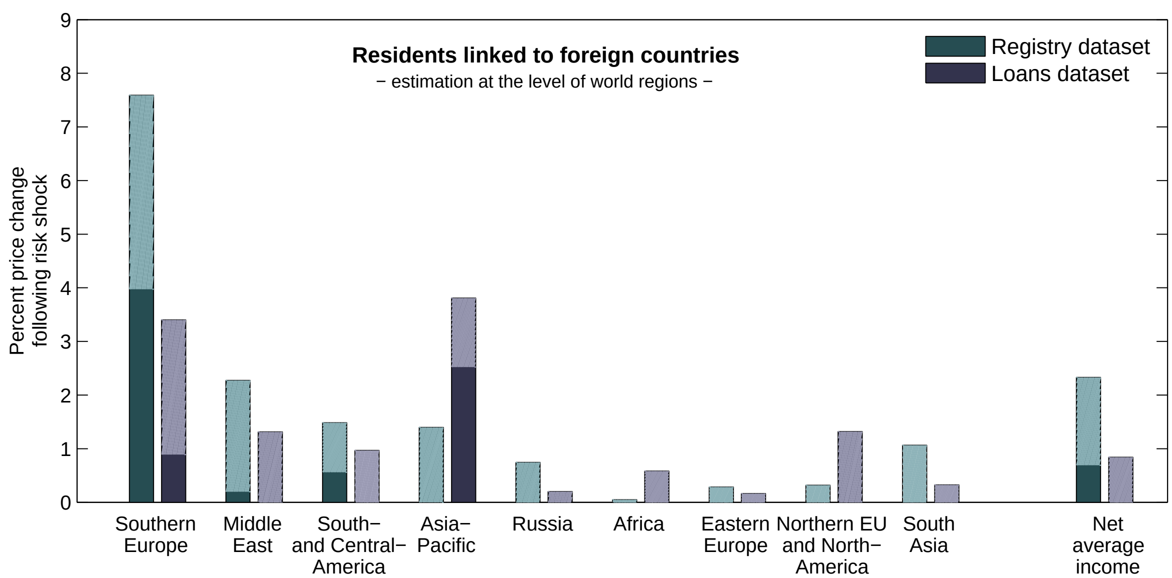


Figure A.10

Time series of capital flows into London's commercial real estate market

The figure reports the evolution of capital inflows into the London commercial real estate market and their relationship with political risk. The data source is Real Capital Analytics. We report the sum of the total inflows from our sample countries with relatively high levels of political risk, as listed in Figure A.2.

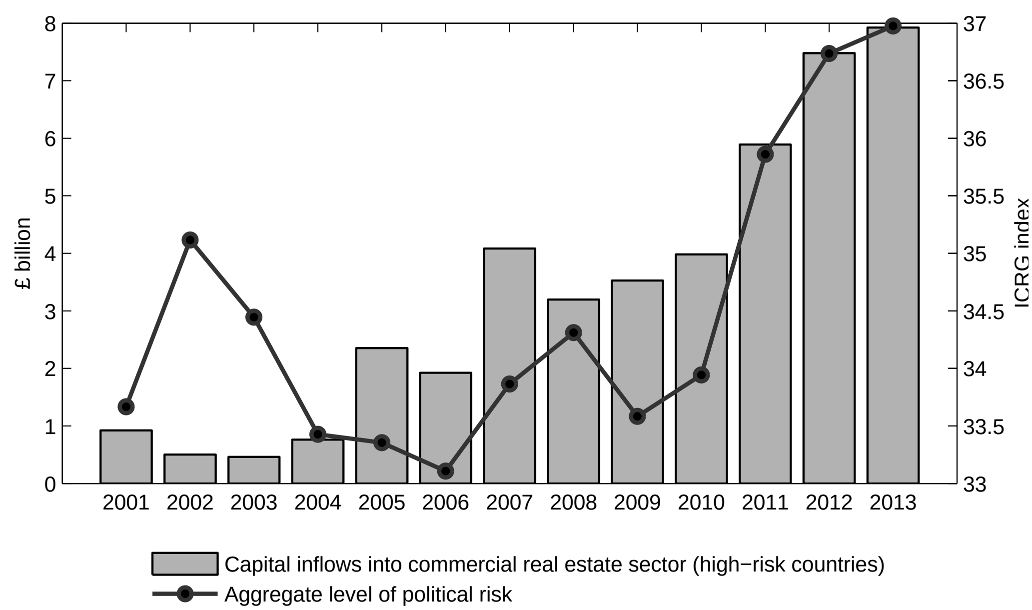


Figure A.11
Foreign political risk and migration into the UK

In this figure, we report the number of additional visas granted by the UK in 2013, relative to 2008. The line indicates univariate cross-country fitted values. On the horizontal axis, we report the change in political risk (measured by the ICRG index) between 2008 and 2013. In this representation, we exclude countries for which the number of visas or the number of people which enter the UK are equal to zero.

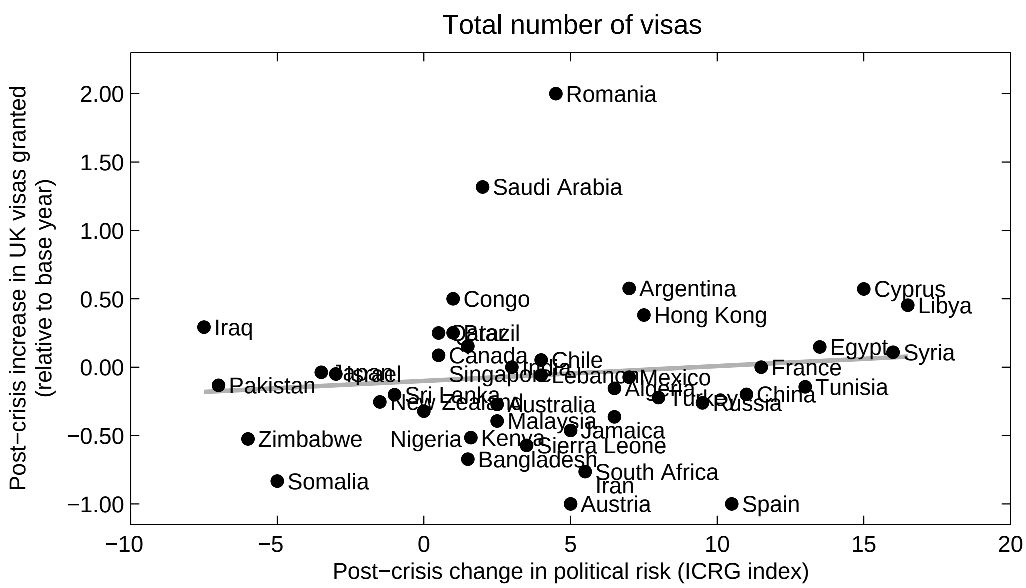
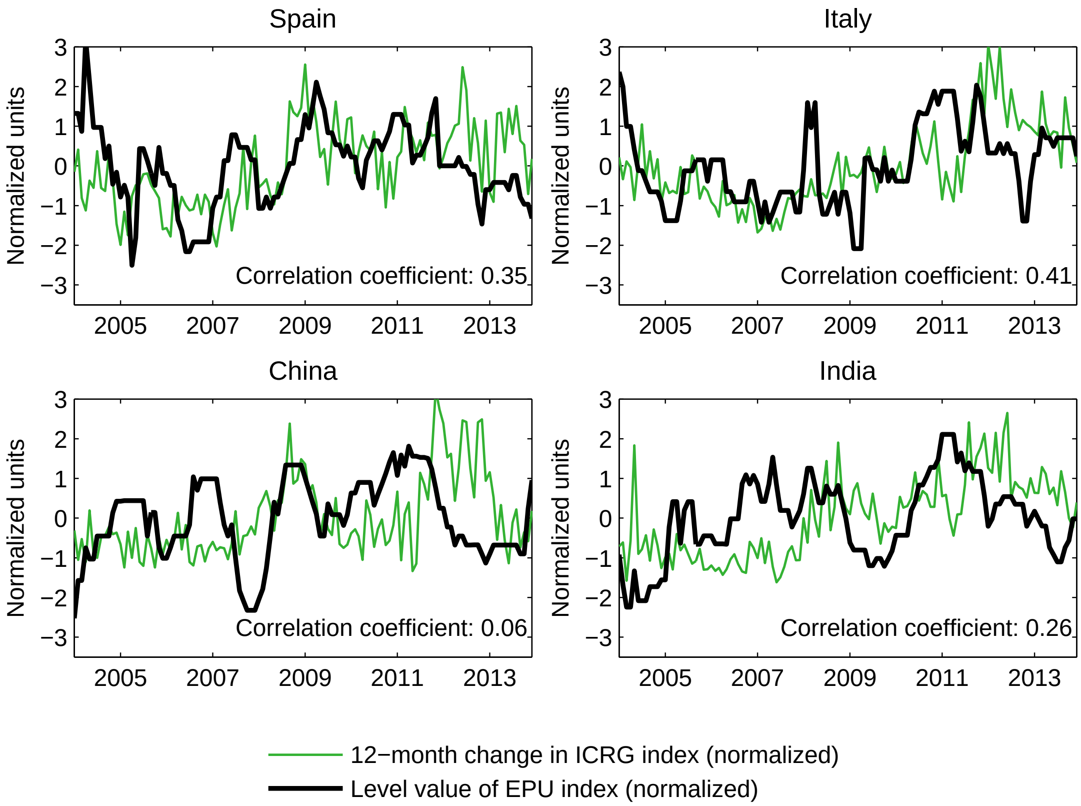


Figure A.12

Relationship between the ICRG and EPU indexes

The figure compares the time series patterns of the ICRG indexes of political risk and the Baker, Bloom and Davis (2013) economic policy uncertainty indexes at country level. The variables are normalized by subtracting the mean and dividing by the in-sample standard deviation.



Immigration and House Prices

One of the possibilities we consider in our specifications is that cross-border property investments into London are driven purely by a desire to move capital away from regions with high political and economic uncertainty, without any associated immigration of foreign purchasers into London. Yet another possibility is that safe-haven property investments incorporate an implicit or explicit future consideration by purchasers of future London-bound immigration. If this is indeed the case, 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. We therefore investigate whether price increases in wards with higher shares of foreign-born people are a signal of increased future immigration into those wards.

Any such immigration might be expected to occur at a much lower frequency than the safe-haven price effects, with longer-lasting effects on the demographic structure of London. Given data availability, we use the U.K. Office for National Statistics census information recorded in 2001 and 2011 to test this hypothesis.

We estimate the following regressions:

$$\Delta f^k_{w,2011} = \alpha + \rho^k f^k_{w,2001} + \pi^k_1 \Delta \ln P_{w,2001} + e^k_{w,2011}, \tag{A.1}$$
$$\Delta f^k_{w,2011} = \alpha + \rho^k f^k_{w,2001} + \pi^k_2 \Delta \ln \bar{P}_{w,2001} + \pi^k_3 \Delta u_{w,2001} + e^k_{w,2011}. \tag{A.2}$$

In these regressions, $\Delta \ln P_{w,2001}$ is the actual log price change between 1996 and 2001 in ward w, computed by equal-weighting prices of all properties transacted in ward w in each of those years. $\Delta \ln \bar{P}_{w,2001}$ and $\Delta u_{w,2001}$ are constructed by controlling for variation in price-impacting hedonic characteristics of properties at the ward level. $\Delta \ln \bar{P}_{w,2001}$ is the change in the fitted value of the price arising from hedonic price regressions in 1996 and 2001 and $\Delta u_{w,2001}$ is the difference in the residuals from these regressions between these two time periods.

In our interpretation of the results, we identify the coefficient π^k_3 with safe-haven demand effects for the purposes of this auxiliary exercise. We are limited by the fact that we only have two available vintages of the census data, from 2001 and 2011. Consequently, we are only able to run a cross-sectional regression to explain variation in the immigration share between these two vintages. This means that we cannot use time-variation in economic and political risk in our attribution of the impacts of safe-haven demand effects on price, and hence, we simply attribute unexplained-by-hedonics variation in prices between 1996 and 2001 ($\Delta u_{w,2001}$) to safe-haven demand effects. If other factors are responsible for this unexplained variation in prices, as long as they are uncorrelated with future immigration, we would expect them to act as classical measurement error, biasing π_3^k towards zero.

Together, specifications (A.1) and (A.2) allow us to check whether price changes have a role in predicting subsequent changes in future immigration over and above the lagged level of immigrants from country k residing in ward w. These regressions, while interesting, are only able to provide suggestive evidence on the interplay between house prices and immigration patterns, both across wards and through time. Figure A.13 shows estimates of equations (A.1) and (A.2). The figure shows that price changes in wards occurring between 1996 and 2001 are a statistically significant and positive predictor of immigration occurring thereafter from Spain, Italy, Portugal, and China. The first bar in these plots corresponds to actual pre-2001 price changes, while the second bar corresponds to the component of the price changes which is unexplained by property and ward characteristics. It is clear from these plots that the variation in hedonic characteristics between 1996 and 2001 is not responsible for the predictive power of prices for the immigration shares. These results are consistent with safe-haven demand causing price pressure in ward-level house prices which subsequently results in immigration flows from these countries. However, it is worth noting here that we view this part of the analysis as far less precise than our earlier specifications which explain house price movements.

The figures also show that these unexplained price changes are negative forecasters of immigration from the South Asian countries. This highlights another important limitation of this analysis of immigration, namely, that unexplained changes in ward-level prices may be generated by a number of potential determinants, including safe haven flows from other countries. This in turn might act as a deterrent to relatively less well-off immigrants from other regions of the world. So, for example, if certain wards experienced unusual price increases from 1996 to 2001 on account of safe-haven demand from, say, Russia, and if immigrants from, say, Sri Lanka shied away from wards with high price increases not caused by their own house purchases, then this would explain the negative coefficients π_3^k that we detect for Sri Lanka.

FIGURE A.13

Relationship between house prices and immigration shares

The figure reports the coefficients π_1^k and π_3^k from the regressions:

$$\begin{aligned}\Delta f_{w,2011}^k &= \alpha + \rho^k f_{w,2001}^k + \pi_1^k \Delta \ln P_{w,2001} + e_{w,2011}^k, \text{ and} \\ \Delta f_{w,2011}^k &= \alpha + \rho^k f_{w,2001}^k + \pi_2^k \Delta \ln \bar{P}_{w,2001} + \pi_3^k \Delta u_{w,2001} + e_{w,2011}^k.\end{aligned}$$

Here, $\Delta \ln P_{w,2001}$ is the actual log price change between 1996 and 2001 in ward w, computed by equal-weighting prices of all properties transacted in ward w in each of those years. $\Delta u_{w,2001}$ is the residual price change in ward w, constructed by controlling for variation in price-impacting hedonic characteristics of properties at the ward level. $\Delta \ln \bar{P}_{w,2001}$ is the component of total price changes which can be attributed to changes in characteristics between the two time periods. The price variables are normalized by subtracting the in-sample mean and dividing by the standard deviation. The estimation sample consists of the 624 London wards. The total length of the bars indicates point estimates and the shaded areas correspond to 95% confidence intervals. The estimated standard errors are White heteroskedasticity-robust.