Preferred Habitats and Safe-Haven Effects: Evidence from the London Housing Market*

Cristian Badarinza and Tarun Ramadorai†

April 17, 2014

Abstract

The infrequent nature of economic and political crises means that pure time-series methods struggle to distinguish safe-haven demand effects on asset prices from a wide range of alternative drivers. We present a new cross-sectional approach, motivated by the insight that investors may have different "preferred habitats" within a broad asset class. We employ this strategy on large databases of historical housing transactions in London, finding that economic and political risk in Southern Europe, China, the Middle East, Russia, and South Asia helps explain price and volume dynamics in the London housing market over the past two decades. Safe-haven effects on the London housing market are long-lasting and significant, but temporary. The method also uncovers intriguing insights about cross-country variation in preferred habitats within London.

*The first version of this paper was circulated under the title "Home Away From Home? Safe Haven Effects and London House Prices". 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, Cam Harvey, Ralph Koijen, Florencio Lopez-de-Silanes, and seminar participants at the Saïd Business School, the Oxford-Man Institute of Quantitative Finance, Goethe University, the Dutch National Bank, the Duisenberg Institute of Finance, and Cass Business School for useful comments, and Job Boerma and Shuang Zheng for dedicated research assistance.

†Badarinza: Saïd Business School and Oxford-Man Institute of Quantitative Finance, University of Oxford, Park End Street, Oxford OX1 1HP, UK. Email: cristian.badarinza@sbs.ox.ac.uk. Ramadorai: Saïd Business School, Oxford-Man Institute of Quantitative Finance, University of Oxford, Park End Street, Oxford OX1 1HP, UK, and CEPR. Email: tarun.ramadorai@sbs.ox.ac.uk.

1 Introduction

To paraphrase Jane Austen, it is a truth universally acknowledged that at times of market stress, global investors flee towards "safe-haven" assets. Such safe-haven or "flight-to-quality" demand has been used by press commentators and academics to explain unusual variation in the prices of a number of assets, including gold, international real estate, currencies, sovereign debt, and international equities, during crises.[1] A clearer understanding of safe-haven demand effects can help pin down the drivers of time-varying risk aversion. Moreover, such effects can have long-lasting consequences on macroeconomic aggregates such as global imbalances, especially when manifested in relatively illiquid and high unit-value assets such as real-estate.

Our paper contributes to this broad area of research in two ways. Our methodological contribution addresses an important identification challenge which arises because safe-haven demand generally appears during periods of country-specific or global crisis. Historical time-series data is short relative to the frequency of crises. During crises, returns, fundamentals, regulation, and market frictions generally tend to be synchronized. These issues make it problematic to use pure time-series methods to convincingly attribute asset price movements to price pressure from safe-haven demand, rather than to movements in price-relevant information, the impact of regulation, or crisis-induced market frictions.

The new strategy that we propose begins with the insight that safe-haven investors may be heterogeneous, and have different "preferred habitats" within a broad asset class.[2] This allows us to develop an empirical methodology to estimate safe-haven

[1] 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).

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

effects which uses the power of the cross-section of asset prices, in addition to the time-series. We view this methodology as being more generally applicable in identifying safe-haven effects in a range of asset prices.

Our empirical contribution is to employ this strategy on large databases of historical housing transactions in London, a city which has received substantial attention for its unusual house price appreciation during the great recession. To explain this phenomenon, media commentators have routinely invoked price pressure from the safe-haven demands of overseas purchasers.[3] We provide evidence that safe-haven demand effects from Southern Europe, China, the Middle East, Russia, and South Asia are indeed an important part of the explanation for the recent movements of London house prices.

Our empirical approach is to sub-divide London into smaller geographical areas, and to enumerate the strength of the links of each such London area with specific foreign countries. When there is an increase in political or economic risk in a given foreign country, our specifications forecast intra-London rates of price appreciation which differ according to the strength of the links between the London area and the foreign country.

A simple 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 safe-haven 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 safe-haven property purchases. If this conjecture is correct, following heightened uncertainty 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.

[3]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.

In our specifications, transactions prices in London areas are modelled as linear in the pre-existing population of Greek-born residents, but this is of course easily generalizable to any economically plausible functional form. Using this approach, we find strong effects from political and economic uncertainty in Southern Europe, China, and East Asia on the prices of houses in London areas with high shares of people originating from these regions of the world.

In our empirical implementation, we also consider the possibility that safe-haven demand is concentrated in desirable London regions, driven by high-net worth foreigners purchasing premium properties in response to uncertainty in their home countries. We use cross-sectional variation in net average income across London areas to capture this notion of desirability.

Using this additional identification, we find that London areas with high average income levels experience unusually high prices following increases in political risk in China, the Middle East, East Asia, and Russia. On the other hand, London areas with relatively lower average income levels experience higher prices following turbulence in Southern Europe and South Asia. These findings using London-locality-specific average income levels are over and above the safe-haven effects captured by the foreign-origin share. 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).[4]

We are also able to use our method to predict variation in London housing transactions volumes, and to explain part of the well-documented association between house prices and transactions volumes (see, for example, Stein, 1995).

---

[4] 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)).

Our estimated safe-haven effects are likely underestimates, as they rule out homogenous impacts across London areas. Nevertheless, we find that they are large. In London areas with a one standard deviation higher share of people originating from a particular country, we find that prices are approximately 40 basis points higher in months following a one standard deviation increase in that country's average risk over the previous year. Controlling for this effect, on average, house prices in London areas with one standard deviation higher average income levels are elevated by an additional 25 basis points. These estimated impacts of elevated foreign risk on London house prices are long-lived but transitory, becoming statistically indistinguishable from zero after roughly three years. We also find effects on relative housing returns across wards, and not just on prices.

Our specifications also allow us to uncover intriguing cross-country variation in the preferred habitat of foreign investors. Safe-haven capital flows from Russia and the Middle East appear to be directed towards premium areas of London, rather than towards areas in which there are pre-existing residents originating from these regions of the world. In contrast, investors from Southern European countries tend to favour regions of London with higher shares of residents originating from their countries, but are less likely to direct capital towards premium residential areas or luxury properties.

Our work is related to the literature on international capital flows and contagion. Using a wide range of methodologies and data sets, this literature finds that there is a strong relationship between risk and institutional quality in source countries, and their extent of outbound direct and portfolio investment (see, for example, Alfaro et al. 2008, and Forbes and Warnock 2011). Our findings confirm that economic and political risk in a number of world regions predict future movements of house prices in London over and above variation in hedonic characteristics, suggesting that these cross-border investments generate price pressure in destination country assets. Our results also suggest that cross-border capital flows and lending relationships can act as a channel for the international transmission of risk, joining a growing list of papers including Kaminsky et al. (2004), Broner et al. (2006), Jotikasthira et al. (2012), and Schnabl (2012).[5]

Our work is also related to the literature on the determinants and predictors of real estate prices (see, for example, Ghysels et al., 2012, who provide a comprehensive survey of the ability of local and aggregate variables to forecast real estate returns in the United States). Our methodological approach is similar in spirit to Mian and Sufi (2009), who use zipcode level geographical variation across the United States to measure the impact of credit supply factors on household mortgage indebtedness and residential property markets. Aizenman and Jinjarak (2009), and Jinjarak and Sheffrin (2011) also connect foreign capital flows with local house prices. Our results are consistent with these papers, but are distinguished by a focus on the identification of safe-haven effects using a new cross-sectional approach, and implementation using transactions-level data.

Finally, our results provide some insights into the famous Lucas (1990) puzzle about the reason for the limited flow of capital from relatively rich to relatively poor countries.[6] One reason for the "uphill" flow of capital from poor to rich countries may be the desire of a subset of poor-country residents to insure themselves against economic and political risk in their home countries.

The remainder of the paper is organized as follows. Section 2 describes our methodology, and Section 3 the data that we employ. Section 4 discusses the results from our empirical estimation, and Section 5 concludes.

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

[6] See, for example, Prasad, Rajan, and Subramanian (2007), and Carroll and Jeanne (2009) among others for more recent empirical evidence and theoretical rationales for this puzzle.

2 Methodology: Identifying Safe-Haven Effects

This section describes our methodology for identifying the impacts of safe-haven demand effects on asset prices. Within a broad asset class, there may be specific assets which are "preferred habitats" for certain groups of investors. If we can measure variation in investor-group-specific demand, and identify the preferred habitat asset for each investor group, we can employ cross-sectional information in addition to time-series information to identify the impacts of demand on asset prices.

2.1 Safe-Haven Effects in London's Housing Market: Hedonic Regressions

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

$$\ln P_{i,t} = \alpha + \beta \mathbb{X}_i + \Pi_{w,t} + u_{i,t}. \tag{1}$$

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

Our approach is to restrict these location-time fixed effects $\Pi_{w,t}$ in an economically meaningful fashion. These restrictions allow us to identify the impacts of political and economic uncertainty, which varies across regions of the world, and time, on the cross-location time-variation in London house prices.

Specific sub-regions of London (of the total set of 624 London electoral wards) may be more attractive investment destinations for capital originating from particular parts of the world. One way in which we link London wards with specific regions of the world is to use the share of the population of each London ward which originates from a particular country or part of the world, but this is of course generalizable to any measure of the "preferred habitat" for demand shocks emanating from specific regions of the world. The choice of the appropriate links is an important one, as non-rejection of the null of no safe-haven effects is possible either because there are indeed no identifiable price impacts of safe-haven demand, *or* because the identification of preferred habitat is not correct. We discuss this issue in greater detail below.

Why is it sensible to link particular foreign countries to London by using the share of London residents in sub-regions of the city who originate from these countries? First, at least some portion of safe-haven housing demand is likely to be driven by calculations of subsequent *physical* safe-haven movements, i.e., the potential for subsequent immigration to London. If so, the cultural affinity of prospective immigrants for their future neighbourhoods in London is very likely to affect their real-estate demand. This line of reasoning does not require the immigration to materialize, only that it is a factor influencing property selection. That said, we find evidence that patterns of growth in ward-level immigration shares in our data are consistent with this line of reasoning.

Second, even if an implicit or explicit future immigration motive is not a factor affecting purchasing decisions, there may be social network effects associated with foreign-origin settlement in particular London areas. These may help to lower informational asymmetries in property purchases. This could occur through direct communication between foreign-origin local London residents and overseas safe-haven buyers. This could also occur because of specialty realtors, local legal firms, and other soft infrastructure set up to match overseas purchasers of specific nationalities with property investment opportunities in specific London areas.[7]

[7] 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.

More formally, let $z_{k,t-1}$ be an indicator of economic and political conditions for a specific country or world region k, in period $t-1$.[8] We model $\Pi_{w,t}$ in the following fashion:

$$\Pi_{w,t} = \delta_t + \phi_w + \sum_{k \in K} \gamma^k f_w^k z_{t-1}^k. \qquad (2)$$

In equation (2), δ_t are time fixed effects, ϕ_w are electoral ward fixed effects, and f_w^k is the fraction of people in ward w who were born in country or world region k. In our empirical implementation, we estimate variants of this specification in which we sequentially set k to be a specific country, as well as ones in which we include multiple countries or regions k simultaneously. The latter specifications allow for the possibility that political and economic uncertainty may be correlated across countries, and that country-of-origin population shares may also be correlated across London wards. We also check robustness in our country-specific estimation to orthogonalizing each z_{t-1}^k measure by regressing it on "global uncertainty," which we measure as the first principal component of all country-specific z measures.

The inclusion of the time fixed effects in equation (2) eliminates common time-variation in London house prices, departing from the usual strategy of identifying safe-haven effects in the time-series. 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.

The ward fixed effects control for the possibility that some London wards may have higher prices on average than other wards, also eliminating this source of variation. The remaining coefficients are purely estimated off simultaneous cross-time- and cross-ward variation in London house prices.

[8] 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 our empirical approach, we employ moving averages of indicator variables z over a year prior to t. In this sense, our approach is predictive, and not just purely explanatory.

If London wards with relatively high pre-existing shares of population from country k are preferred locations for safe-haven demand emanating from country k, and if there are price impacts on London house prices as a result of this demand, γ^k will be estimated positive. It is of course possible that safe-haven demand for London properties is driven by buyers who wish to locate in London regions that are *not* highly populated by residents originating from their countries. If so, we would estimate γ^k to be negative. In this sense, the estimated coefficients allow us to derive interesting insights about specific countries in addition to identifying safe-haven effects. Throughout, the obvious null hypothesis is that $\gamma^k = 0$.

This identification strategy is susceptible to the possibility that there may be omitted London ward-level characteristics which attract foreign safe-haven demand (i.e., in the space of the ward cross time effects), and which we might mistakenly attribute to the country-of-origin share.

An obvious possibility along these lines is that safe-haven demand is concentrated in relatively more desirable London wards (such as King's Road in Chelsea or Mayfair), driven by high-net worth foreigners in risky parts of the world moving capital into premium London real estate. In order to control for this possibility, we add a ward-level indicator of desirability y_w as an additional conditioning factor in the interaction term (in empirical estimation, we use net average ward-level income as y_w to capture this notion of desirability):

$$\Pi_{w,t} = \delta_t + \phi_w + \sum_{k \in K} (\gamma_0^k f_w^k + \gamma_1^k y_w) z_{t-1}^k. \tag{3}$$

We also estimate a specification in which we include the total non-UK-origin ward share (i.e., the percentage of total foreign-origin population in each ward) in addition to the country-specific shares f_w^k, as an additional interaction with z. This is to check whether a more accurate measure of the "tilt" of a ward towards a specific country results in more precise estimation of safe-haven effects.

We also check whether the effects that we estimate using our ward-level identification strategy are simply capturing effects that occur at a broader level of geographical aggregation within London. For example, it might be the case that all foreign-born residents or high-income regions of London are concentrated in certain boroughs of London (of which there are 32),[9] such as Kensington and Chelsea, or Westminster, and increases in foreign political risk might affect all wards within a same borough similarly. To check whether we benefit from the more granular identification conferred by the 624 wards, we also include borough-time fixed effects in our regressions and check whether we continue to find statistically significant safe-haven effects γ_0^k and γ_1^k which are identified using only within-borough, cross-ward variation in foreign-born origin shares and income levels interacted with variation in political and economic risk.

All standard errors in our tables and figures from the hedonic regression approach are clustered at the ward-time level (using the robust clustering procedure described in Cameron and Trivedi (2007)) to account for unexplained commonalities in the residuals.[10]

[9] We do not include the "City of London" in our analysis because it is a separate administrative and ceremonial entity, with mostly commercial properties.

[10] In the Land Registry sample, we cluster at the ward-month level, and in the Nationwide sample of mortgage loans, we cluster at the ward-year level on account of more sparse data in this sample. We report ward-level volume statistics in the online appendix.

2.1.1 A Note on Magnitudes

In our empirical implementation, we cross-sectionally demean f_w^k and y_w, and time-demean z_{t-1}^k. We also standardize these variables by their cross-sectional (f_w^k and y_w) and time-series (z_{t-1}^k) standard deviations. In all tables and figures, we also multiply the estimated coefficients by 100, to allow for the easy interpretation of magnitudes as percentages of relative house price appreciation.

To illustrate the total safe-haven price impact that we estimate, consider a ward in which the share of people born in country k, and the level of net average income are both one standard deviation higher than their cross-ward means. An increase of one standard deviation in the 12-month moving average of political risk in country k predicts house price appreciation of $\gamma_0^k + \gamma_1^k$ in the subsequent month. Note that safe-haven effects are constrained by our functional form to be linear in f_w^k and y_w, but we could easily generalize this to any plausible non-linear form.[11]

[11] The online appendix also shows the results when using terciles of z rather than the level, i.e., linearity in z is also something we can easily change in our specifications.

2.2 Sources of Safe-Haven Effects

The discussion thus far has tended to emphasize that one channel for cross-border property investments into London is a desire to move capital away from regions or countries with high political and economic uncertainty. Here we note that property transactions in London have high unit value, meaning that only a subset of country residents may be able to participate in these markets.

To investigate whether safe-haven price impacts are driven by the wealthiest residents of particular countries, we split our identification of price impacts by interacting the γ_0^k and γ_1^k variables with dummy variables that capture the relative price of properties in London, checking whether foreign-origin safe-haven demand is most prevalent at the very top of the market.[12] In order to do this, we simply sort properties within each of 32 London boroughs each year on the basis of their transaction price. A transaction is then classified as either belonging to the bottom (below the 70^{th} percentile), high-value (70^{th} to 90^{th} percentile), or ultra-high value (above the 90^{th} percentile) group on the basis of its rank in the borough-year distribution.

Of course, it is possible that residents of countries that are further down the wealth distribution also move capital towards London or other safe-haven investments in response to uncertainty. However, given the high unit-value involved in residential real estate investments, it is more likely that these investments incorporate an implicit or explicit future consideration of future London-bound immigration.

To explore the immigration channel, we estimate*:*

$$\Delta f_{w,2011}^k = \alpha + \rho^k f_{w,2001}^k + e_{w,2011}. \tag{4}$$

[12] 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.

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. This specification allows us to provide further evidence on our identification strategy in our hedonic pricing regressions in the previous subsection, if ρ^k is estimated to be statistically different from zero – although it is worth noting that our identification strategy is not conditional on immigration flows actually materializing. Given data availability, we are constrained to using the U.K. Office for National Statistics census information recorded in 2001 and 2011 in these tests.

The immigration channel also raises the issue of 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 uses relatively sparse data to investigate whether price increases in wards with higher shares of foreign-born people are a signal of increased future immigration into those wards.

2.3 Safe-Haven Effects in London's Housing Market: Spreads

In the absence of multiple repeat sales of individual properties, it is difficult to see how best to control for the (potentially autoregressive) dynamics of house prices in our hedonic regression specification (1). Moreover, it is potentially of interest to understand the time-pattern of safe-haven effects on *changes* in prices. While the issue of short time-series relative to the infrequent incidence of crises continues to be an important one in this context, we attempt to use the cross-sectional insights conferred by our identification strategy in a complementary dynamic panel modeling approach.

We define $s_{k,t}$ as the period t "price spread" between the logarithm of average house prices in the top and bottom quantiles of London wards (in our empirical application we use quintiles), ranked by the strength of their links to country k. $s_{k,t}$ captures relative house price differentials between wards with strong and weak links to specific countries, and should be unaffected by aggregate variation in London house prices. We estimate the following distributed-lag specification to explain variation in $s_{k,t}$:

$$s_{k,t} = \mu_k + \sum_{q=1}^{Q} \rho_q s_{k,t-q} + \sum_{q=1}^{Q} \zeta_q z_{k,t-q} + u_{k,t}, \tag{5}$$

where $z_{k,t-q}$ is the foreign risk factor for country k measured at lag q, and μ_k are country fixed effects. In our empirical application, we generally estimate coefficients ζ_q in an unrestricted fashion, and occasionally restrict them to correspond to an equal-weighted moving average of past z_t. These ζ_q are the counterparts to the estimated γ_0^k in our hedonic specification (2), and rejections of the null that $\zeta_q = 0 \ \forall Q$ imply a role for safe-haven demand effects on the London housing market.

To accurately capture the dynamics of safe-haven effects, we also capture the time-series behaviour of the country-specific risk factors $z_{k,t}$ using an autoregressive distributed-lag panel specification. This imposes the (to our mind plausible) restriction that there is no feedback from London house prices to political and economic uncertainty in foreign countries:

$$z_{k,t} = \theta_k + \sum_{q=1}^{Q} \pi_q z_{k,t-q} + \varepsilon_{k,t}, \tag{6}$$

We can use equations (6) and (5) to evaluate the impact of a shock $\varepsilon_{k,t}$ to political risk on the relative price level between London wards with high and low levels of people originating from country k.

Finally, we modify these specifications to explain the growth rates of price spreads rather than the levels of these spreads:

$$s_{k,t} - s_{k,t-12} = \mu_k + \zeta(z_{k,t-12} - z_{k,t-24}) + u_{k,t}. \tag{7}$$

In our empirical implementation, we estimate equation (7) for the full sample using overlapping year-on-year changes on both left- and right-hand sides (with the necessary correction for standard errors in the presence of overlapping left-hand-side variables). We also estimate the regression specification annually, using non-overlapping annual first differences of log spreads, and lagged annual changes in z. Despite the inevitable loss of power at this lower frequency, this specification constitutes a useful robustness check of our main results.

2.4 Safe Haven Effects and Transaction Volumes

To complement the evidence obtained from prices, we investigate whether the foreign uncertainty measures and our identification strategy are related to London housing market volume.

We first ask whether our estimated safe-haven mechanism can predict London housing transaction volumes, by estimating the following specification, in which we try to explain the number of transactions $V_{w,t}$ in each ward w and period t as follows:

$$\ln V_{w,t} = \mu + \vartheta_w + \varsigma_t + \sum_{k \in K} \left(\chi_0^k f_w^k + \chi_1^k y_w \right) z_{t-1}^k + \upsilon_{w,t}. \tag{8}$$

In equation (8), we check if χ_0^k and χ_1^k are greater than zero. Our posited mechanism works through the channel of cross-border housing transactions, so if we are indeed able to identify these demand movements, they should show up our ability to predict the variation in housing transactions across ward-months.

Many authors, such as Stein (1995), and Ortalo-Magne and Rady (2006), have shown that aggregate housing market transactions volume is positively associated with prices. Stein (1993) presents a rational model in which he explains this association with the fact that 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.

If our approach is able to capture safe-haven cross-border housing transactions in London, we should observe attenuation in the correlation between volume and prices in each period of time, once we include our safe-haven demand measures on the right-hand side. If so, this would provide an additional explanation for the correlation between house prices and volume, namely, demand fluctuations generated by safe-haven effects.

To investigate this issue, we estimate the following two equations, with the same notation as before:

$$\ln P_{i,t} = \alpha^N + \beta^N \mathbb{X}_i + \phi_w^N + \delta_t^N + \theta^N \ln V_{w,t} + u_{i,t}^N, \text{ and} \tag{9}$$

$$\ln P_{i,t} = \alpha + \beta \mathbb{X}_i + \phi_w + \delta_t + \theta \ln V_{w,t} + \sum_{k \in K} \left(\gamma_0^k f_w^k + \gamma_1^k y_w \right) z_{t-1}^k + u_{i,t}. \tag{10}$$

The correlation between prices and volume is captured in these specifications by θ^N, which, going by the observations in prior literature, we should expect to be statistically significant and positive. If our safe-haven effects are able to explain some of the association of contemporaneous ward-level housing transactions volume on prices, then we should observe that $\theta < \theta^N$.

3 Data

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

3.1 Registry data

We obtain data on the complete set of house purchases in London from HM Land Registry. All purchasers of houses are required to report transactions to the Land Registry, and the data cover 2,254,590 transactions over the period from 1996 to 2012. This amounts to 12% of the roughly 18.5 million residential property sales the Land Registry has lodged for England and Wales. Property characteristics reported with these data include the type of house (whether it is an apartment, semi-detached or terraced house, for example), the tenure status (whether the property is a leasehold or a freehold property) and an indicator of whether the property is newly built.

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

3.2 Loan-level data

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

The reason that we use Nationwide data in addition to the Land Registry data is to allow us to better control for hedonic characteristics of properties, which can be important for house-price determination, as seen in recent work such as Campbell et al. (2011). For each individual property in the Nationwide London sample, 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. The data provider also indicates the date at which the loan was approved, and the purchase price of the property. Due to the sensitive nature of the mortgage credit transactions, the information pertaining to the borrower is restricted to whether he or she is a first-time home buyer.

However, the Nationwide data do come with some limitations. The online appendix shows that ward-month transactions volume in this dataset is quite sparse, which is not surprising given the far smaller set of transactions that Nationwide covers relative to the Land Registry. Therefore in our analysis of volume, we restrict ourselves to the Land Registry dataset.

3.3 UK Office for National Statistics

There are 1.8 million postcodes active in the UK, corresponding to 29 million postal addresses, an average of roughly 16 buildings per postcode. The UK Office for National Statistics (ONS) publishes a postcode directory, allowing the establishment of unique relationships between postcodes and other geographical units such as electoral wards. For each individual housing transaction from both Nationwide and Land Registry samples, we identify the associated geographical unit in which the house is located. We then match this to information on the demographic and economic characteristics of that unit, also available from the 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.

For each electoral ward, we use data from the ONS corresponding to the year 2001. A key variable in our analysis is the share of each ward's population that was born in foreign countries. For China, Japan, Malaysia, Singapore, India, Pakistan, Sri Lanka, Italy, Portugal, Greece, and Spain, the ONS reports this share precisely in the 2001 census. However, for Russia and the Middle Eastern countries in our sample, these data are not available, so we use the number of people in the ward who speak Russian and Arabic instead, as reported in the 2011 census. We also use average income at the ward-level as reported in 2001. The online appendix shows histograms of these and other ward-level variables.

The complete set of countries and world regions which we consider in our analysis is shown in Panel A of table 1. For the world regions East Asia, South Asia, and Southern Europe, we generate regional aggregates of all variables in the analysis by weighting individual countries' data by (lagged) GDP expressed in US Dollars, obtained from the World Bank.

3.4 Economic and Political Risk Measures

We characterize economic and political risk in foreign countries and world regions using three separate variables. First, we obtain a list of indicators from the International Country Risk Guide (ICRG), which are monthly data on the political situation around the world. These indicators rate each country along 12 dimensions, each of which contributes to the total country rating according to the number of points indicated in parentheses: government stability (12 points), socioeconomic conditions (12 points), investment opportunities (12 points), the potential for internal and external conflicts (24 points), the level of corruption (6 points), the influence of military in politics (6 points), the prevalence of religious tensions (6 points), general conditions in terms of law and order (6 points), ethnic tensions (6 points), democratic accountability (6 points), and the quality of bureaucracy (4 points). For each country, we build a composite index by simply summing across these 12 risk categories (by ICRG construction, this adds up to 100), and we again use GDP weights to build our time series for world regions.[13]

Glaeser et al. (2004) note that the ICRG indicators do not describe the permanent state of country-level political institutions, but rather reflect actual changes that happen through time. For the purposes of our analysis, it is precisely this time series dimension in which we are interested – capturing changes in the political situation in a given country, rather than solely capturing differences across country-level institutions. They also emphasize that the ICRG indicators capture subjective, not solely objective assessments of risk, which is also relevant, as we are most interested in capturing prospective safe-haven investors' sentiment about their local political environment. These 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.

[13] 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.

We also use two alternative measures of country-level economic and political 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.[14] 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 2012. For some countries long-term bond data are unavailable; we describe the specific data restrictions for these variables in the online appendix.

Second, we use the economic policy uncertainty indexes of Baker, Bloom, and Davis (2013). These are available at monthly frequency for China, Spain, Italy, and India, and cover different time periods, as detailed in the online appendix. Baker, Bloom, and Davis (2013) show that economic policy uncertainty is an important factor in determining the allocation of capital, and that it spills over negatively into investment, output, and employment dynamics in a variety of countries.

We find that higher economic policy uncertainty is associated with increases in our other political risk measure – in the online appendix, we report short-run co-movement between 12-month changes in the ICRG indexes of political risk and the economic policy uncertainty measure, country-by-country. Since the Baker, Bloom and Davis (2013) measure is based on news coverage and forecaster disagreement, it spikes during periods in which the initial turmoil occurs and tends to revert quickly. In contrast, the ICRG indexes tend to be more persistent.

In our empirical implementation, we subtract the time-series mean for each of the above economic and political risk variables, and divide them by the time-series standard deviation across the entire sample period, as mentioned earlier. We then take moving

[14] 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.

averages over 12 months of each of these variables to capture periods of increasing and decreasing uncertainty, making sure to lag them by at least one month relative to the left-hand side house price variables.[15]

[15] We also use one-month changes in these indexes, lagged twelve months as an alternative to the lagged averages, with very little qualitative effect on our results. These results are available in the online appendix.

4 Results

4.1 Time-Series Patterns

The starting point of our analysis is the widely documented emergence of a house price gap between London and the rest of the UK. In the top panel of figure 1, we illustrate this using a set of UK regional house price indexes, reported by Nationwide, Halifax (now owned by Lloyds), the Land Registry, and the UK ONS. For each set of indexes, we plot the percentage spread between the price of the average house in London relative to a (population in 2001-) weighted average house price in the remainder of the UK. All four series clearly show that the spread between London house prices and those in the remainder of the UK is very large on average. Moreover, this spread fluctuates substantially over time. There is a pronounced increase in this spread beginning in 1998, a period of heightened international political and economic uncertainty owing to the Asian and Russian financial crises. Following this period, London prices appear to grow at roughly the same rate (or even slightly lower) as the remainder of the UK during the early part of the decade beginning in 2000. Finally, there is strong growth in the spread following the onset of the financial crisis beginning in 2008.

How much of this increase is attributable to external political and economic uncertainty? The bottom panel of figure 1 shows the relationship between the London house price spread and a 12-month trailing moving average of the two economic and political indicators (bond yield spreads over the UK, and the ICRG political risk index). These indicators are GDP-weighted across all of the non-UK countries in our sample available in each time period. In the time-series, these indicators appear very closely related with the level of London house prices relative to the remainder of the UK.

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

4.2 Hedonics

Panel B of table 1 focuses on micro-data on housing transactions, and shows the total number of transactions in each year, that are reported by the Land Registry as having occurred in London, as well as the number of mortgages issued by Nationwide which are associated with housing transactions in London. Interestingly, mortgage volumes and housing transactions both dip in 2008 and remain low relative to historical averages during the credit crunch, while London house prices are increasing over the same period, relative to prices in the remainder of the UK. While these broad time-series trends are not allowed by our approach to affect our identification of safe haven effects, we investigate the impacts of ward-time variation in housing market illiquidity on prices, and the relationship of this variation to our identification approach.

In our empirical analysis, as described earlier, we present both country-by-country results (z^k_{t-1} and f^k_w pertaining to a single country k), and regionally aggregated results.

Figure 2 shows the correlation between the shares of people born in specific countries (darker shaded areas of the matrix represent higher correlations), and confirms that London residents that come from culturally and geographically proximate countries tend to live near one another. For example, London wards with high shares of residents originating from India also have high shares of residents originating from Pakistan and Sri Lanka, and relatively low shares of people originating in Greece, Italy, Portugal, and Spain. We group the former set into world region "South Asia," and the latter set into world region "Southern Europe," as seen in Panel A of table 1.

Figure 3 shows estimates of equation (4) 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 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, as the result is strong and robust for all of the ones for which data is available.

The online appendix shows that wards vary in a number of characteristics, and that wards more densely inhabited by people of foreign origin do have important differences with the remainder. However, these static characteristics of wards are not allowed to affect our estimation as we include ward-fixed effects in our estimation. It is worth noting here that the fraction of owners with mortgages appears to be lower in wards with high shares of foreign-born people. While it is of course possible that foreigners' access to the UK mortgage system is less straightforward than access for UK-born London residents, we see this fact as consistent with the argument that foreign-born people draw upon sources of funds which lie outside the UK mortgage system, and at least partially, may come from overseas.[16]

[16] We tried including a control for this variation in mortgage usage by interacting the total foreign-born origin share with the level of the UK mortgage interest rate, and found that our results were

The ward fixed effects do not completely eliminate variation across properties arising from property-specific hedonic characteristics. To control for any such variation, we employ the Nationwide dataset, which has data on a far more comprehensive set of hedonic characteristics than the Land Registry dataset. In all of the tables and figures in the paper, we report results from both of these datasets. For the most part, results are strongly consistent across the two datasets, despite the fact that the Land Registry data cover all London transactions, whereas the Nationwide data cover only a fraction, i.e., only those associated with mortgages from this financial institution. We view the consistency between the two sets of results as evidence that price effects from safe-haven flows spill over into more general local price increases that are experienced by London households.

Table 2 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 dataset.

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

unaffected by the use of this control.

builds), and detached houses are worth more than any other category of houses, once floor-space area is controlled for.

Our empirical specifications focus on interactions between *ward-level* characteristics and foreign political and economic uncertainty; 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 – for example, we could check whether the prices of newly-built properties are more strongly associated with safe-haven demand relative to other types of residences by interacting a dummy for new-builds with the lagged political and economic uncertainty measures.[17]

Finally, panel C of table 2 shows the estimated time fixed effects, averaged across all months in each year. The time fixed effects show steady appreciation in London house prices relative to their 1996 level, with the only recorded declines in 2008 and 2009, but strong recovery in 2010 and 2011.

We next turn to presenting the main point of our analysis, which is the empirical identification of the safe-haven demand effects from our specifications.

[17] See, for example, "Foreign investors snap up 70% of all central London new build homes fuelling a surge in prices.", www.thisismoney.co.uk, 16 August 2013.

4.3 Preliminary Results: Total Foreign-Born Share

A preliminary illustration of our results, which aggregates country-specific information into a single foreign-born share, is provided in figure 4. The figures show maps of London, with electoral wards color-coded in a "heatmap" – red shades indicate greater levels of particular variables, and blue shades 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 uncertainty. When juxtaposed, the figures clearly reveal that the London wards with the greatest house price appreciation shifted towards the centre and northwest of London. Panel C of the figure shows that this appears to be correlated with the high net-income areas of London as measured in the 2001 census, although the high-income areas in the South and South-West of London appear not to have experienced particularly high house price appreciation over the 2007 to 2012 period. Finally, Panel D of the figure shows the total share of foreign-origin people in London measured in the 2001 census. The pattern is visually striking – the house price appreciation between 2007 and 2012 appears to line up well with the share of foreign-origin people in London wards. Indeed, the simple cross-ward correlation between 2001-2006 price appreciation and 2001 foreign-origin shares is -12%, while the cross-ward correlation between 2007-2012 price appreciation and 2001 foreign-origin shares increases to $+38\%$.

The online appendix shows that the use of the single foreign-born origin share and a measure of global uncertainty yields strong and statistically significant results in our specifications. However, this masks the important and economically interesting cross-country variation in safe-haven effects which it is possible to estimate using our method. This is the topic of the next subsection.

4.4 Cross-Country Results, Single Country

In our specification (3), the coefficients γ_0^k and γ_1^k constitute our identification of safe-haven demand effects on London house prices. Figure 5 plots these coefficients when we employ the ICRG political risk indicator as z_{t-1}^k, and when country-by-country analysis is employed, i.e., the set K consists of a single country at a time, and the reported coefficients come from separate country-by-country regressions. In these figures, the two bars for each country show the results from the Land Registry and Nationwide datasets, and the comparison between these bars indicates that the results are quite similar despite being estimated on different sets of data, and with different sets of hedonic controls in these regressions.

In these figures, the total length of each bar shows the total coefficient estimate, with the (smaller) statistically significant portion of the coefficient estimate also highlighted within each bar. Across bars, the shading of the bars shows estimated positive coefficients in a dark shade, and estimated negative coefficients in a lighter shade. There are several possible interpretations of negative coefficients. One possibility, which we think is most likely, is that price appreciation associated with political risk in particular countries is concentrated in wards with relatively *low* average income levels, or *low* foreign-born shares.

The second possibility is that *declines* in political risk in particular countries are associated with house price appreciation in high average income level or high foreign-origin share wards. This would be consistent with a wealth effect – if declines in political risk occurring over the period are associated with increases in prosperity. Finally, it is possible, of course, that high political risk in particular countries is associated with relative price declines in London wards with high income or foreign-origin shares, which could be associated with the imposition of capital controls or other restrictions on the ability of country residents to move capital overseas at such times, and the consequent reduction in capital flows to their preferred London areas.[18]

Panel A of the figure shows the results when z is the ICRG index. The left and right hand parts of this panel plot coefficients γ_0^k (interaction with the ward-share of foreign-born people) and γ_1^k (coefficient with the ward-level average income level) respectively. The figure shows that there are strong positive effects on house prices in wards with high shares of people originating from a particular country, following periods of elevated political risk in those countries. For example, for Egypt, following a one-standard deviation shock to average political risk measured over the previous year by the ICRG index, the (statistically significant) coefficient shows that house prices in wards with a one-standard deviation higher level of Arabic speaking people experience a one percentage point appreciation in house prices over the subsequent month, over and above variation in hedonics and other ward characteristics, and over and above the average rise in London house prices. Controlling for this effect, there is also a separate effect of political risk in Egypt on wards with high levels of average income – of roughly half the size of the effect on wards with high shares of Arabic speakers. These magnitudes are large, but also potentially an overstatement, since in these country-specific regressions, the impacts of political risk in country j not included in the regression are also absorbed in country k's coefficient, to the extent that the z's and the f_w's are correlated across countries j and k (which is likely to be the case).

Panels B and C of figure 5 show the results estimated when z is the long-term bond yield spread of countries measured relative to the UK, and the economic policy uncertainty index of Baker, Bloom, and Davis (2013), and the results seem broadly

[18] These last two explanations would be consistent with a negative coefficient in a forecasting regression of prices in wards populated with a high share of country residents by the political uncertainty measures that we employ. Despite the obvious econometric issues with forecasting highly-persistent price levels, we investigate this issue in the online appendix, and find evidence that the forecasting coefficient in such regressions is estimated positive. Wards with a low share of foreign-born residents on the other hand, have prices which are not well forecasted by our uncertainty measures.

consistent across countries. What is more clearly evident comparing the two pictures of γ_0^k and γ_1^k here is that there appears to be an inverse relationship (controlling for the other determinant) of income and foreign origin, which is also evident in Panel A.

For example, when Greece and countries in South Asia experience high political risk, this appears to be associated with higher future prices in London wards with high shares of people originating from these countries, but controlling for this tendency, one interpretation is that the price appreciation is relatively more concentrated in wards with low average incomes. To illustrate this point with a specific example, following heightened political 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.

In figure 6, we explore this issue further, plotting γ_0^k and γ_1^k against one another. This allows us to explore interesting cross-country variation. To aid understanding of this plot, consider a hypothetical country for which γ_0^k and γ_1^k are both estimated to be high. The estimates suggest that for this country, in response to political risk, and controlling for ward desirability, capital is more likely to be directed towards regions of London with higher pre-existing shares of population originating from the country. Also, controlling for the foreign-born origin share, purchasers from the country are more likely to direct investment towards more desirable regions of London.

The actual estimates from figure 6 show a negative relationship between the two estimated coefficients, which captures economically interesting cross-country variation in the preferred habitat of foreign investors. The position of the countries in this plot suggests that safe-haven investors from Russia and the Middle East prefer premium areas of London, and holding this tendency constant, appear to disfavour areas of London in which there are pre-existing residents originating from these regions of the world. On the other end of the spectrum are investors from Southern European countries, who generally tend to favour areas of London in which residents originating from their countries are based, and are less likely to direct capital towards premium London residential real estate.

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 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.

Table 3 reports country-by-country estimation results in which we include an additional interaction term between the country-specific level of foreign political risk, and the total share of foreign-born people in each ward. The estimated coefficients γ_0^k now isolate the safe-haven impacts associated with the "excess tilt" of London wards towards residents originating from a specific country, over and above the total share of people originating from any country other than the United Kingdom. This specification delivers the additional benefit of controlling for any unobservable ward-specific factor that may be correlated with places in which non-UK-origin residents reside in London. We find that the estimated safe-haven effects γ_0^k are still strong and statistically significant in this specification.

While these effects are indeed interesting, it is clear that our single-country specifications suffer from the issue that measures of uncertainty and foreign-born origin shares may be correlated across countries.[19] As a result, the country-by-country estimation in these panels may be more likely to capture the impacts of global rather than country-specific uncertainty. In the next sub-section, we more reliably estimate country-specific safe-haven effects, by including all countries in our sample simultaneously, grouped into world regions as described earlier.

[19] See, for example, the evidence reported by Longstaff et al. (2011) that sovereign yields are highly correlated across countries.

4.5 Cross-Country Results, World Regions

Table 4 shows the results of estimating our hedonic regressions with K world regions included simultaneously, i.e., we include $\sum_{k \in K} (\gamma_0^k f_w^k + \gamma_1^k y_w) z_{t-1}^k$ in our regressions. We report the estimated coefficients γ_0^k and γ_1^k for each of the six world regions in the first column of the table. In these specifications, we use the ICRG measure of risk as our z_{t-1}^k variable, GDP-weighted within each region.

This specification shows that political risk in Southern EU (Italy, Greece, Spain, Portugal) and China are strongly associated with future price appreciation in London wards with high shares of people born in these regions. Controlling for this tendency, there is relatively greater price appreciation in high-income wards following Chinese political risk, and relatively greater price appreciation in low-income wards following Southern European uncertainty. There is roughly the reverse tendency in response to political risk in Russia – high average income London wards appreciate following shocks to this variable, but controlling for this tendency, price increases appear concentrated in wards with relatively low shares of Russian-born people.

In terms of magnitudes, the effect appears large, even though it is likely an underestimate of the total effect taking into account the common impacts of foreign risk on London wards. Averaging across world regions, a foreign risk shock of one standard deviation over a year predicts that house prices in the subsequent month in wards which have a one standard deviation higher share of foreign-origin people will be elevated by 40 basis points. Controlling for this foreign-born share effect, the effect of foreign political risk is 30 basis points on wards with a one-standard deviation higher average income level. These numbers come from the Land Registry estimates; comparable magnitudes are evident from the Nationwide data.

Our identification of safe haven effects uses both the time series dynamics of foreign risk and the cross-sectional variation of house prices across the 624 London wards.

The second column of table 4 checks 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^k in a specification which includes borough-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 China, South Asia, and Southern Europe. Interestingly, though there is attenuation of the γ_1^k coefficients in this specification, all of them remain statistically significant – within-borough cross-ward income variation seems important in identifying safe-haven effects.

In figure 7, we extend the analysis to allow for the possibility that the impacts of foreign political risk differ by the price of the property, as described in the methodology section. This is to investigate whether the safe-haven effects are driven by the wealthiest purchasers from foreign nations seeking to preserve capital in response to elevated political and economic risk.[20]

Figure 7 shows that the impacts of foreign political risk on house prices are always increasing in price, regardless of the world region or ward-level variable. Of course, the reduction in sample size in the very highest price category means that some of these effects are not very precisely estimated, but the economic magnitudes of the effects are substantial. In some cases, the effects cause the interactions to change sign in a direction which is more consistent with our initial hypothesis – for example, controlling for the impact of Russian political risk on the prices of houses in high net average income wards, we now find that for the highest price category of houses, there

[20] 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.

is also price appreciation of houses in London wards with a high share of Russian-born people. For South Asia, the picture is also interesting – there is a tendency for greater price appreciation within wards with relatively low average income levels following high political risk, but within these relatively low income wards, the price appreciation appears to be the largest for the most expensive properties.

4.6 Price Dynamics Across Wards

We identify safe haven effects through variation in the dynamics of house prices across London electoral wards. Using hedonic regressions, we show that a price gap opens up between London areas with higher shares of foreign-origin residents and those with weaker international ties, following elevated economic and political risk in foreign countries. However, our inferences from our hedonic regression specifications may be subject to a number of issues. First, our identification of safe haven effects may, especially in our country-specific estimation, be vulnerable to common shocks to political risk over our 17 year sample. Second, foreign capital flows may have persistent effects on house prices which are difficult to account for in a hedonic setup with relatively low levels of repeat sales. Third, we would like to be able to estimate safe-haven impacts on housing rates of return rather than simply on their levels.

As described in the methodology section, we complement our hedonic regressions by computing price "spreads" $s_{k,t}$ 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. We then stack these spreads estimated for all countries each month into a panel.[21] We introduce time fixed effects into this specification to control for the possibility

[21] Since we use the same f_w^k for our entire group of seven Arabic countries, the resulting price spreads are identical for these countries. In our empirical implementation, we pool this group of countries into a single entity, weighting right-hand side variables by the level of US Dollar-denominated GDP. All panel spread specifications are thus estimated for a set of 13 distinct entities.

of common shocks, and estimate an autoregressive distributed-lag specification in the panel, to account for the persistence of house prices. We also correct the standard errors for any residual serial correlation, computing them using the Driscoll-Kraay (1998) approach, and a lag-length of 12 months.[22] Finally, we also estimate results on ward-level housing returns in a panel which conditions cross-ward return spreads on lagged changes in foreign economic and political risk.

Table 5 summarizes our findings. The first column presents estimated coefficients from a simple specification in which we condition the panel of spreads on lagged political uncertainty. The results show that a one standard deviation increase in the 12-month moving average of political risk overseas is associated with prices that are 3.2 percentage points higher (from the Registry dataset estimates) in London wards with high foreign-origin shares, than in wards with low foreign-origin shares. When we eliminate any identification from common shocks using monthly time dummies, the second column shows that the estimated safe haven effects reduce to 1.2 percentage points, but continue to be highly statistically significant.

The third column shows results using year-on-year changes in foreign risk to explain year-on-year changes in price spreads (spread or relative returns), and the fourth column uses non-overlapping one-year changes in the specifications. In both cases, we find that safe-haven effects are still important when we don't include time fixed-effects. This is a particularly stringent test, since it doesn't allow periods of elevated political risk in different countries to be used to predict variation in London house prices, only allowing upticks and downticks in political risk to do so (which are very likely to be correlated across countries). Nevertheless, the table shows that strong safe-haven effects are still

[22] 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).

estimated in the full Registry sample of transactions using changes in both price spreads and political risk.

All of the above patterns are also visible in the two right-most panels, in which we replace the measures of political risk with 10-year bond yield spreads, and the Baker, Bloom and Davis (2013) economic policy uncertainty indexes, respectively.

Figure 8 shows the estimated evolution of the spread in response to a one standard deviation shock to foreign political risk. This is estimated from a system of equations in which we condition spreads on lagged spreads and lagged political risk,[23] and political risk on its own lags. The figure estimated using the Registry dataset shows that the predicted spread in house prices begins at 0.98 percent, and plateaus between 2.57 and 2.62 percentage points at a horizon of 18-29 months. At 38 months after the shock, the effect becomes statistically indistinguishable from zero. Similar patterns are seen in the Loans dataset.

The estimated impulse response functions suggest that safe-haven effects on the London housing market are temporary. Properties in wards with higher foreign-origin shares trade at higher prices following increases in foreign political risk, but after 38 months, the predicted price gap arising from this source is statistically indistinguishable from zero. The relatively long lag before the effects dissipate are consistent with the high persistence of political risk, and substantial house price inertia which is likely a consequence of the difficulty of arbitraging price discrepancies in this market.

[23] We choose the lag length of 18 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.

4.7 Evidence on Transaction Volumes

In table 6 we report estimation results from specifications (8), (9), and (10). Panel A shows that our foreign risk measures are associated with an increase in transaction volumes. For example, following a one standard deviation shock to political uncertainty in Southern Europe, transaction volumes are 0.84 percent higher in a month in wards with a one standard deviation higher share of people born in Southern Europe.

Panel B of the table 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. The panel also shows that this relationship is partially explained by the evolution of political risk outside the UK, as the coefficient on volume attenuates following the inclusion of the foreign-risk safe-haven channel. It appears that the price-volume relationship in the London housing market seems to be partially driven by the higher demand generated by purchases from overseas investors.

5 Conclusions

In this paper, we propose a novel empirical method to identify the impact of safe-haven demand on asset prices. The approach takes as its starting point the possibility that investors are heterogeneous, and might have different "preferred habitat" assets within a broad asset class. Using this method and large micro-datasets from the U.K., we empirically estimate the safe-haven demand effects of foreign economic and political risk on the London housing market. We find economically large, statistically significant, and robust effects of overseas risk on house prices in locations in which the share of foreign-origin London residents are high, and controlling for this effect, on desirable (high average income) London areas.

We view our results as providing a more general contribution to the analysis of macro-variation in prices, by using an identification strategy that is grounded in agents' microeconomic motivations. Our approach could be used, with potentially minimal customization, in any situation in which the demand for safe-haven investments demonstrates cross-sectional as well as time-series variation.

Our empirical results document that increases in risk in Southern Europe, China, Russia, and the Middle East are associated with well-identified increases in London house prices in specific sub-regions of London, providing a more rigorous analysis of the recent rise in London house prices, a phenomenon that has been widely commented upon in the press, and been the subject of numerous policy debates.

In subsequent versions of this paper, we hope to provide a clearer insight into the specific mechanisms involved in the transmission of price spillovers in housing markets, using London as our laboratory.

6 References

Aizenman, J. and Y. Jinjarak, 2009, Current Account Patterns and National Real Estate Markets, *Journal of Urban Economics*, 66, 75-89.

Alfaro, L., S. Kalemli-Ozcan and V. Volosovych, 2008, Why Doesn't Capital Flow from Rich to Poor Countries? An Empirical Investigation, *Review of Economics and Statistics*, 90(2), 347-368.

Baker, S., N. Bloom and S. J. Davis, 2013, Measuring Economic Policy Uncertainty, Chicago Booth Research Paper No. 13-02.

Baur, D. G. and T. K. McDermott, 2010, Is Gold a Safe Haven? International Evidence, *Journal of Banking and Finance*, 34(8), 1886-1898.

Beber A., M. W. Brandt and K. Kavajecz, 2009, Flight-to-Quality or Flight-to-Liquidity? Evidence from the Euro-Area Bond Market, *Review of Financial Studies*, 22, 925-957.

Bollerslev, T., Patton, A., and W. Wang, 2013, Daily House Price Indexes: Construction, Modeling, and Longer-Run Predictions. Unpublished working paper.

Broner, F., Gelos, R. G., and C. Reinhart, 2006, When in Peril, Retrench: Testing the Portfolio Channel of Contagion, *Journal of International Economics,* 69, 203-230.

Caballero, R. J. and A. Krishnamurthy, 2009, Global Imbalances and Financial Fragility, *American Economic Review*, 99(2), 584-88.

Caballero, R. J., E. Farhi, and P. O. Gourinchas, 2008, An Equilibrium Model of "Global Imbalances" and Low Interest Rates, *American Economic Review*, 98(1), 358-393.

Campbell, J. Y., S. Giglio, and P. Pathak, 2011, Forced Sales and House Prices, *American Economic Review*, 101(5), 2108-2139.

Campbell, J. Y., K. Serfaty-de Medeiros and L. M. Viceira, 2010, Global Currency Hedging, *Journal of Finance*, 65(1), 87-121.

Carroll, C. and O. Jeanne, 2009, A Tractable Model of Precautionary Reserves, Net Foreign Assets, or Sovereign Wealth Funds, NBER Working Paper No. 15228.

Case, K. E., and R. J. Shiller, 1989, The Efficiency of the Market for Single-Family Homes, *American Economic Review*, 79(1), 125-137.

Culbertson, J., 1957, The Term Structure of Interest Rates, *Quarterly Journal of Economics*, 71, 485-517.

Dorn, D., and G. Huberman, 2010, Preferred Risk Habitat of Individual Investors, *Journal of Financial Economics*, 97(1), 155-173.

Driscoll, J. C., and A. C. Kraay, 1998, Consistent Covariance Matrix Estimation with Spatially Dependent Panel Data, *Review of Economics and Statistics,* 80, 549-560.

Erb, C. B., C. R. Harvey and T. E. Viskanta, 1996, The Influence of Political, Economic, and Financial Risk on Expected Fixed-Income Returns, *Journal of Fixed Income*, 6(1), 7-30.

Erb, C. B., C. R. Harvey and T. E. Viskanta, 1996, Political Risk, Economic Risk, and Financial Risk, *Financial Analysts Journal*, 52(6), 29-46.

Forbes, K. J., and F. E. Warnock, 2012, Capital Flow Waves: Surges, Stops, Flight, and Retrenchment, *Journal of International Economics*, 88(2), 235-251.

Genesove, D. and C. Mayer, 2001, Loss Aversion And Seller Behavior: Evidence From the Housing Market, *Quarterly Journal of Economics*, 116(4), 1233-1260.

Ghysels, E., Plazzi, A., Torous, W. N. and R. I. Valkanov, 2012, Forecasting Real Estate Prices, Handbook of Economic Forecasting: Vol II, (G. Elliott and A. Timmermann, eds.), Elsevier.

Glaeser, E. L., R. La Porta, F. Lopez-de-Silanes and A. Shleifer, 2004, Do Institutions Cause Growth?, *Journal of Economic Growth*, 9(3), 271-303.

Jinjarak, Y. and S. M. Sheffrin, 2011, Causality, Real Estate Prices, and the Current Account, *Journal of Macroeconomics,* 33, 233-246.

Jotikasthira, C., Lundblad, C., and T. Ramadorai, 2012, Asset Fire Sales and Purchases and the International Transmission of Funding Shocks, *Journal of Finance*, 67, 2015-2050.

Kaminsky, G., Lyons, R. K., and S. Schmukler, 2004, Managers, investors, and crises: Investment strategies of mutual funds, *Journal of International Economics,* 64, 113-134.

Longstaff, F. A., 2004, The Flight-to-Liquidity Premium in U.S. Treasury Bond Prices, *Journal of Business*, 77(3), 511-526.

Longstaff, F. A., Pan, J., Pedersen, L. H., & K. J. Singleton, 2011, How Sovereign Is Sovereign Credit Risk?, *American Economic Journal: Macroeconomics*, 3(2), 75-103.

Lucas, R., 1990, Why Doesn't Capital Flow from Rich to Poor Countries?, *American Economic Review,* 80 (2), 92–96.

Mian, A. and A. Sufi, 2009, The Consequences of Mortgage Credit Expansion: Evidence from the U.S. Mortgage Default Crisis, *Quarterly Journal of Economics*, 124(4), 1449-1496.

Meese, R. A. and N. E. Wallace, 1997, The Construction of Residential Housing Price Indices: A Comparison of Repeat-Sales, Hedonic-Regression, and Hybrid Approaches, *Journal of Real Estate Finance and Economics*, 14, 51–73.

Modigliani, F. and R. Sutch, 1966, Innovations in Interest-Rate Policy, *American Economic Review*, 56, 178-197.

Ortalo-Magné, F. and S. Rady, 2006, Housing Market Dynamics: On the Contribution of Income Shocks and Credit Constraints, *Review of Economic Studies*, 73(2), 459-485.

Prasad, E., Rajan, R., and A. Subramanian, 2007, Foreign Capital and Economic Growth, *Brookings Papers on Economic Activity*, 153–230.

Ranaldo, A. and P. Söderlind, 2010, Safe Haven Currencies, *Review of Finance*, 10, 385-407.

Rosen, S., 1974, Hedonic Prices and Implicit Markets: Product Differentiation in Pure Competition, *Journal of Political Economy*, 82(1), 34-55.

Schnabl, P., 2012, The International Transmission of Bank Liquidity Shocks: Evidence from an Emerging Market, *Journal of Finance*, 67(3), 897-932.

Stein, J., 1995, Prices and Trading Volume in the Housing Market: A Model with Down-Payment Effects, *Quarterly Journal of Economics*, 110, 379-406.

Vayanos, D., and J-L Vila, 2009, A Preferred-Habitat Model of the Term Structure of Interest Rates, Unpublished working paper.

Table 1
Summary statistics

In Panel A, we list the countries which are part of our analysis concerning the effects of external factors on the London housing market. We group the countries in six regions and weight them according to their US Dollar-denominated GDP level, as reported by the World Bank. In Panel B, we report the number of observations per year, for both the transaction-level data from the Land Registry and the loan-level data from the Nationwide Building Society.

Panel A

List of countries

1.	**China**	
2.	**East Asia**	
	Japan	Malaysia
	Singapore	
3.	**South Asia**	
	India	Pakistan
	Sri Lanka	
4.	**Russia**	
5.	**Middle East**	
	Algeria	Egypt
	Libya	Qatar
	S. Arabia	Tunisia
	UAE	
6.	**Southern Europe**	
	Italy	Greece
	Portugal	Spain

Panel B

	Number of transactions	
	Registry dataset	Loans dataset
1996	108,431	12,243
1997	133,223	13,108
1998	133,317	12,463
1999	162,342	19,330
2000	147,370	8,958
2001	162,241	7,522
2002	174,237	9,386
2003	154,603	7,371
2004	165,041	7,565
2005	137,306	6,649
2006	172,021	10,712
2007	166,120	9,143
2008	81,508	3,977
2009	75,207	4,394
2010	91,615	5,068
2011	89,462	6,356
2012	91,885	9,892
Total	2,245,929	154,137

Table 2

Estimated coefficients in the hedonic regression framework

This table reports estimated coefficients from the following hedonic regression:

$$\ln P_{i,t} = \alpha + \beta \mathbb{X}_i + \phi_w + \delta_t + u_{i,t},$$

where $\mathbb{X}_i$ are property-level characteristics. All right-hand side variables are normalized by subtracting the in-sample mean and dividing by the standard deviation. *, **, *** denote statistical significance at the 10%, 5%, and 1% level respectively. The standard errors are clustered at the ward-time level.

Panel A

Registry dataset

Semi detached house	-0.428***
Terraced house	-0.582***
Flat	-0.743***
New property	0.237***
Leasehold indicator	-0.338***

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 2

Estimated coefficients in the hedonic regression framework
(continued)

Panel C

	Estimated time fixed effects	
	Registry dataset	Loans dataset
1997	0.166***	0.217***
1998	0.293***	0.354***
1999	0.417***	0.511***
2000	0.616***	0.699***
2001	0.723***	0.810***
2002	0.883***	0.982***
2003	1.008***	1.080***
2004	1.089***	1.149***
2005	1.120***	1.171***
2006	1.184***	1.255***
2007	1.323***	1.391***
2008	1.308***	1.337***
2009	1.214***	1.255***
2010	1.322***	1.350***
2011	1.339***	1.386***
2012	1.390***	1.418***

Table 3
Country-by-country estimation results

This table reports estimated coefficients γ_0^k from the following hedonic regression:

$$\ln P_{i,t} = \alpha + \beta \mathbb{X}_i + \phi_w + \delta_t + (\gamma_0^k f_w^k + \gamma_1^k y_w + \gamma_2^k \overline{f}_w) z_{t-1}^k + u_{i,t},$$

where $\mathbb{X}_i$ are property-level characteristics. In this specification, z_{t-1}^k is a lagged monthly 1-year moving average of each of the ICRG risk indicators. $\overline{f}_w$ is the total ward-level share of people born outside the UK. The coefficients in the first column correspond to our benchmark empirical specification, where we impose the constraint $\gamma_2^k = 0$. All right-hand side variables are normalized by subtracting the in-sample mean and dividing by the standard deviation. *, **, *** denote statistical significance at the 10%, 5%, and 1% level respectively. The standard errors are clustered at the ward-time level.

	Registry dataset		Loans dataset	
	Baseline	Control for total foreign-born people share	Baseline	Control for total foreign-born people share
Spain	2.49***	2.45***	1.63***	1.48***
Greece	1.48***	0.69***	1.00***	0.27**
Egypt	1.44***	0.48***	0.83***	-0.24
Tunisia	0.95***	0.28***	0.32***	-0.26*
Pakistan	0.91***	0.47***	0.77***	-0.05
China	0.90***	0.34***	0.88***	0.07
Portugal	0.74***	0.12**	0.30***	-0.07
Italy	0.73***	0.36***	0.85***	0.40***
Singapore	0.68***	0.48***	0.17	-0.06
Japan	0.54***	-0.02	0.35***	-0.16**
Malaysia	0.34***	-0.03	0.51***	-0.36***
Qatar	0.24***	0.37***	-0.37***	0.28**
Russia	0.19***	0.06	-0.34***	-0.36***
Sri Lanka	0.17***	0.06*	0.17**	-0.26***
Libya	0.16***	0.44***	0.33***	0.89***
UAE	-0.10**	0.54***	-0.52***	0.93***
India	-0.14***	-0.43***	0.15*	-0.51***
Algeria	-0.35***	0.16***	-0.32***	0.86***
Saudi Arabia	-0.40***	0.22***	-0.25***	0.83***

Table 4
Safe haven effects across world regions

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

$$\ln P_{i,t} = \alpha + \beta \mathbb{X}_i + \phi_w + \delta_t + \tau_{b,\nu} + \sum_{k \in K} \left(\gamma_0^k f_w^k + \gamma_1^k y_w \right) z_{t-1}^k + u_{i,t},$$

where f_k^w are the shares of people in ward w born in world region k, y_w is average net income and $\tau_{b,\nu}$ are borough-year fixed effects, for borough b in year ν. In this specification, z_{t-1}^k is a lagged monthly 1-year moving average of each of the ICRG risk indicators corresponding to the six world regions listed in table 1. All right-hand side variables are normalized by subtracting the in-sample mean and dividing by the standard deviation. In column (1) we report estimation results from our benchmark specification, where we restrict $\tau_{b,\nu} = 0$. In column (2) we include borough-year fixed effects. *, **, *** denote statistical significance at the 10%, 5%, and 1% level respectively. The standard errors are clustered at the ward-time level.

	Registry dataset		Loans dataset	
Residents linked to foreign countries	(1)	(2)	(1)	(2)
China	0.80***	0.21***	0.89***	0.27***
East Asia	0.13***	-0.07*	0.13*	0.00
South Asia	-0.12***	-0.16***	0.08	0.14*
Russia	-0.35***	0.00	-0.83***	-0.28***
Middle East	0.05	0.04	0.00	0.22
Southern Europe	1.72***	0.67***	1.34***	0.76***
Net average income	(1)	(2)	(1)	(2)
China	2.61***	1.93***	1.57***	1.58***
East Asia	0.68***	0.48***	0.27***	0.29***
South Asia	-1.70***	-1.18***	-2.09***	-1.92***
Russia	0.44***	0.32***	1.23***	0.85***
Middle East	0.83***	0.40***	1.52***	1.15***
Southern Europe	-1.35***	-0.68***	-1.29***	-0.71***
Borough-year fixed effects	No	Yes	No	Yes

Table 5

Cross-country panel analysis

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

$$s_{k,t} = \mu_k + \rho s_{k,t-1} + \zeta \overline{z}_{k,t-1} + u_{k,t},$$

where a bar denotes the 1-year moving average. The specification labeled *Year-on-year changes* is given by:

$$s_{k,t} - s_{k,t-12} = \mu_k + \zeta(z_{k,t-12} - z_{k,t-24}) + u_{k,t}.$$

Since we use the same f_w^k for our entire group of seven Arabic countries, the resulting price spreads are identical for all of them. We pool this group of countries into a single entity, weighting right-hand side variables by the level of US Dollar-denominated GDP. All panel specifications are thus estimated for a remaining set of 13 distinct entities. In the case labeled *Non-overlapping growth* rates, only observations from December of each year are included in the sample. When considering bond yields and economic policy uncertainty indices, the samples start in 2000 and 2004 respectively, due to the limited data availability. *, **, *** denote statistical significance at the 10%, 5%, and 1% level respectively. In parentheses, we report Driscoll-Kraay (1998) standard errors with a lag length of 12 months.

	Political risk				**Financial risk**		**Economic risk**	
	(1)	(2)	(3)	(4)	(1)	(2)	(1)	(2)
Registry dataset	0.57***	0.22***	1.06***	2.00**	0.78***	0.93*	0.24*	0.42
	(0.15)	(0.07)	(0.34)	(0.89)	(0.22)	(0.47)	(0.14)	(0.34)
Number of observations	2,496	2,496	2,340	195	1,560	1,543	432	420
Loans dataset	1.39***	0.85***	0.17	-0.06	1.73***	0.63	0.86***	-0.14
	(0.28)	(0.23)	(0.58)	(0.92)	(0.30)	(0.68)	(0.17)	(0.62)
Number of observations	2,496	2,496	2,340	195	1,560	1,543	432	420
Country fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Time fixed effects	No	Yes	No	No	No	No	No	No
Year-on-year changes	No	No	Yes	Yes	No	Yes	No	Yes
Non-overlapping growth rates	No	No	No	Yes	No	No	No	No

Table 6

Foreign risk and transaction volumes

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

$$\ln V_{w,t} = \mu + \vartheta_w + \varsigma_t + \sum_{k \in K} \left(\chi_0^k f_w^k + \chi_1^k y_w \right) z_{t-1}^k + \upsilon_{w,t},$$

where $V_{w,t}$ is the number of transactions in ward w in period t.In Panel B we report the estimated coefficient θ from the following hedonic specification:

$$\ln P_{i,t} = \alpha + \beta \mathbb{X}_i + \phi_w + \delta_t + \theta \ln V_{w,t} + \sum_{k \in K} \left(\gamma_0^k f_w^k + \gamma_1^k y_w \right) z_{t-1}^k + u_{i,t},$$

where z_{t-1}^k is a lagged monthly 1-year moving average of each of the ICRG risk indicators for the six world regions listed in table 1. In the first column, we restrict $\gamma_0^k = \gamma_1^k = 0$. The estimation is carried out in the Registry sample of housing transactions. *, **, *** denote statistical significance at the 10%, 5%, and 1% level respectively. In parentheses, we report standard errors, clustered at the borough-time level in Panel A and the ward-time level in Panel B.

Panel A

Explaining the evolution of transaction volumes

	Residents linked to foreign countries	Net average income
China	0.84***	0.77***
East Asia	-0.09	1.24***
South Asia	0.19*	-7.29***
Russia	-0.07	2.17***
Middle East	2.10***	4.94***
Southern Europe	1.08***	1.70***

Panel B

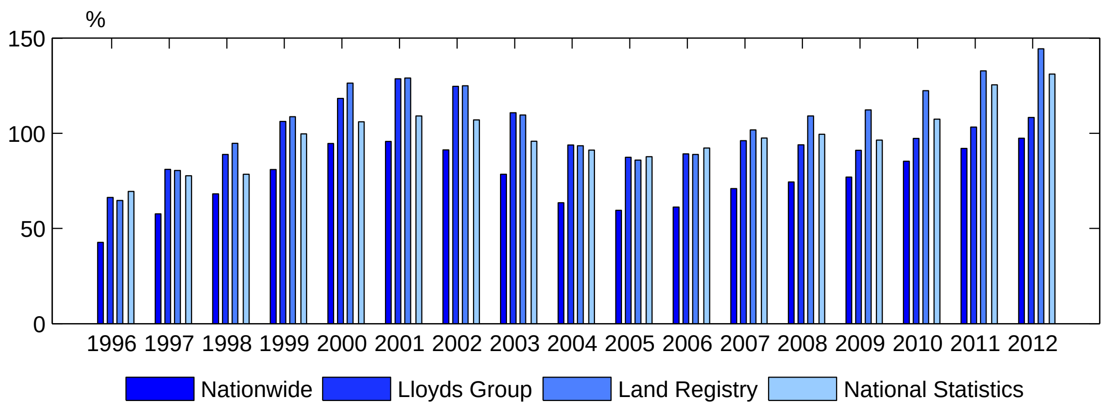
Hedonic house price regression

		(1)	(2)
Transaction volume	θ	1.34***	1.05***
		(0.17)	(0.16)
Save haven effects		No	Yes

Figure 1
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. The list summarizing our country coverage is given in table 1. We generate aggregate values by weighting the observations according to annual GDP in US Dollars, reported by the World Bank. 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

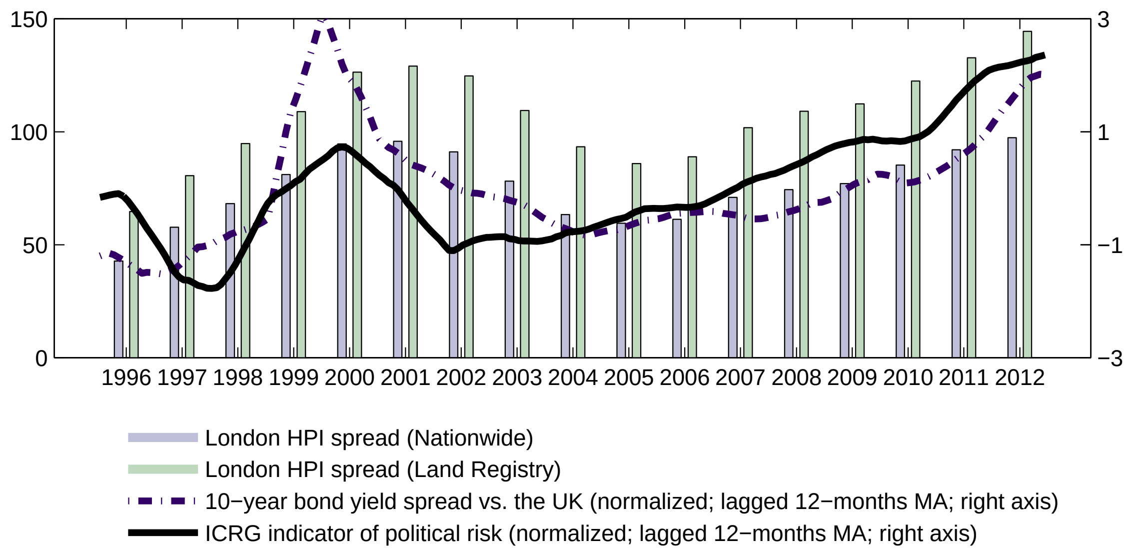


Figure 2
Shares of foreign-born people
- correlation across wards -

This figure shows the pairwise correlation coefficients between the shares of people born in foreign countries, across the 624 London wards. Blocks in darker shading indicate a higher tendency of the population from the respective countries to cluster around similar areas of the city. We calculate the fractions of foreign-born people relative to the total ward population by using 2001 census data, indicating the country of birth. For Russia and the Middle East, we use the number of people speaking Russian and Arabic, respectively, as registered in the 2011 census.

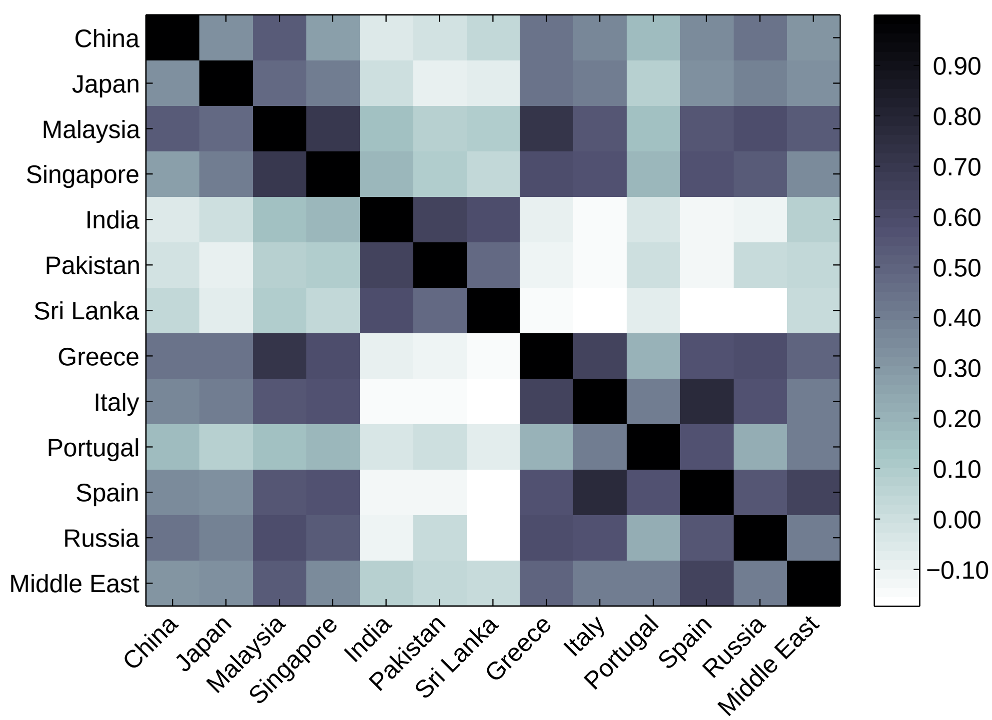


Figure 3

The evolution of foreign-born people shares through time

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

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

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. Statistical significance is reported through error bars, indicating 95% confidence intervals. The estimated standard errors are White heteroskedasticity-robust.

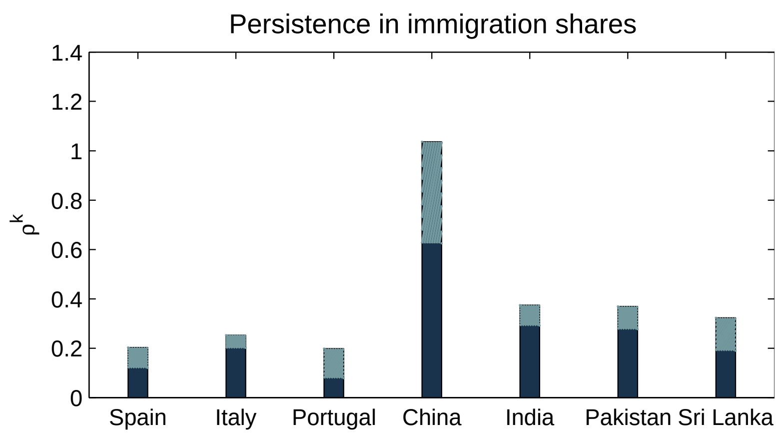


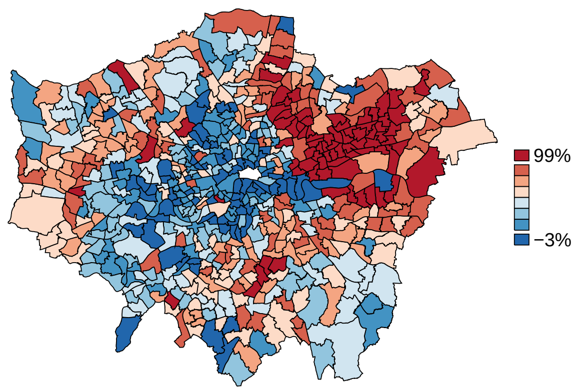
Figure 4

House prices, demographic structure and economic conditions at ward level

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

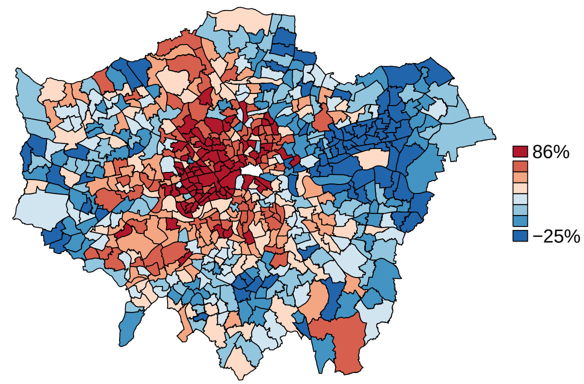
A: House price appreciation

between 2001 and 2006



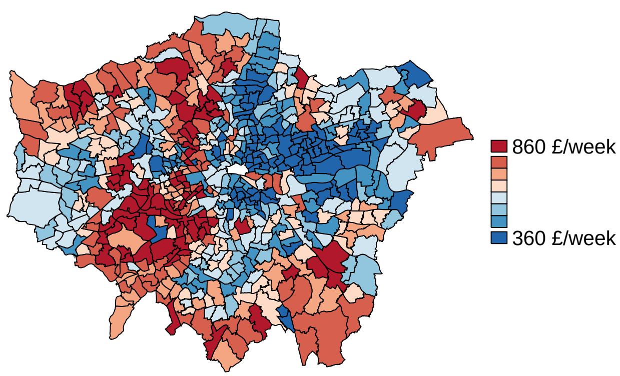
B: House price appreciation

between 2007 and 2012



C: Net average income

ONS (2001)



D: Foreign-born people share

2001 census

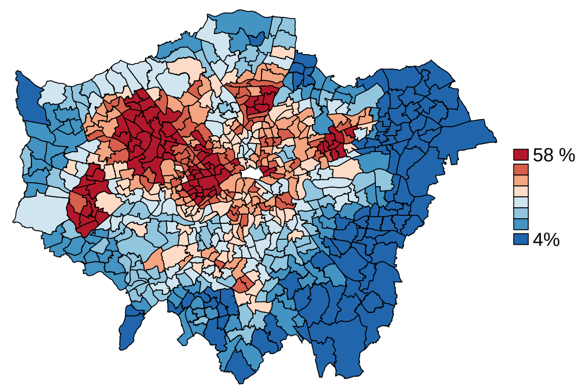


Figure 5
International political risk and London house prices

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

$$\ln P_{i,t} = \alpha + \beta \mathbb{X}_i + \phi_w + \delta_t + \left(\gamma_0^k f_w^k + \gamma_1^k y_w\right) z_{t-1}^k + u_{i,t},$$

where f_w^k are the shares of people in ward w born in country k and y_w is average net income. z_{t-1}^k is a lagged monthly 1-year moving average of each of our risk indicators. All right-hand side variables are normalized by subtracting the in-sample mean and dividing by the standard deviation. We plot absolute values of all estimated coefficients and indicate negative values by using light shading. Statistical significance is reported through error bars, indicating 95% confidence intervals. The standard errors are clustered at the ward-time level.

Panel A

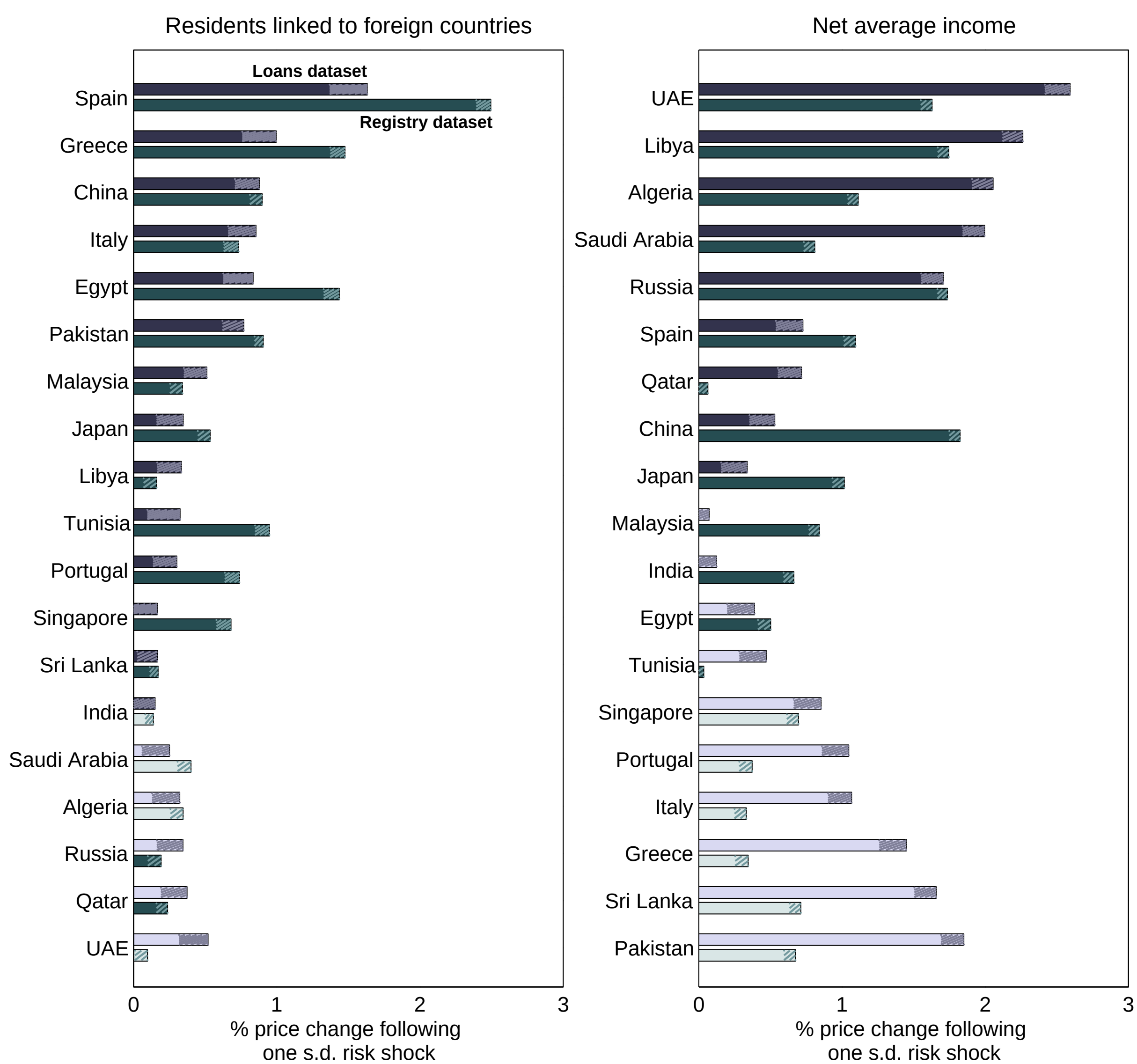


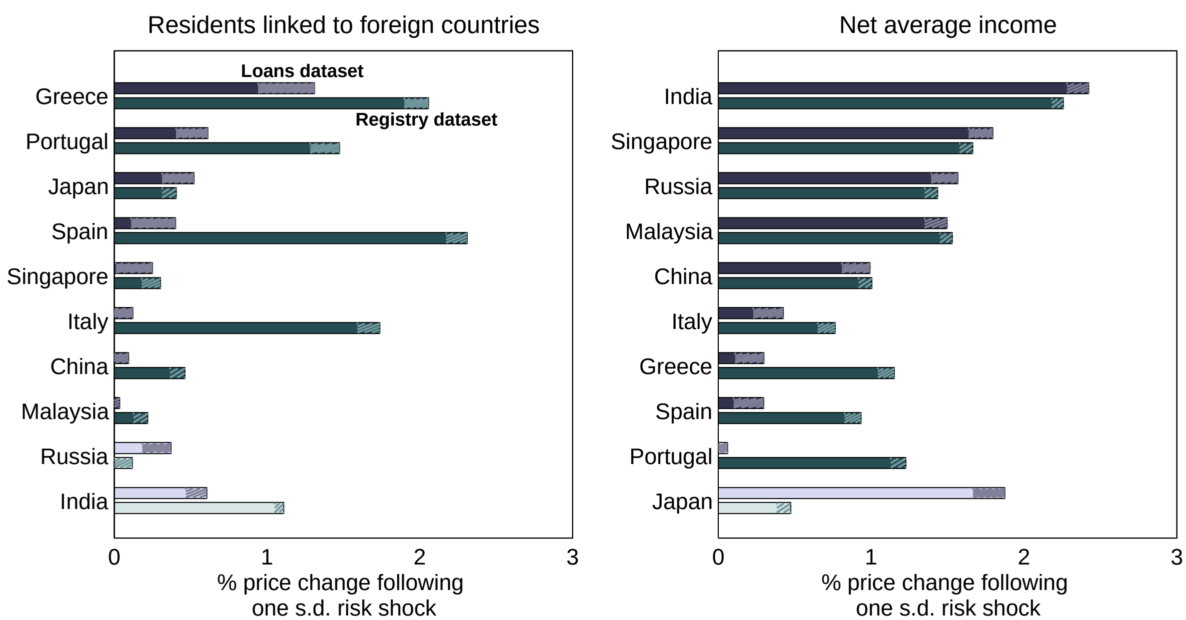
Figure 5

International political risk and London house prices

(continued)

Panel B

Bond yield spreads



Panel C

Economic policy uncertainty

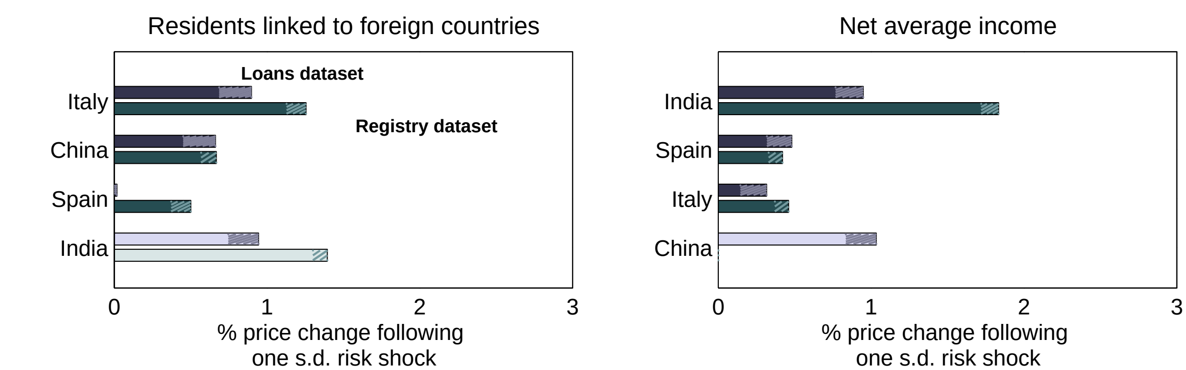


Figure 6

Cross-country overview of estimated coefficients

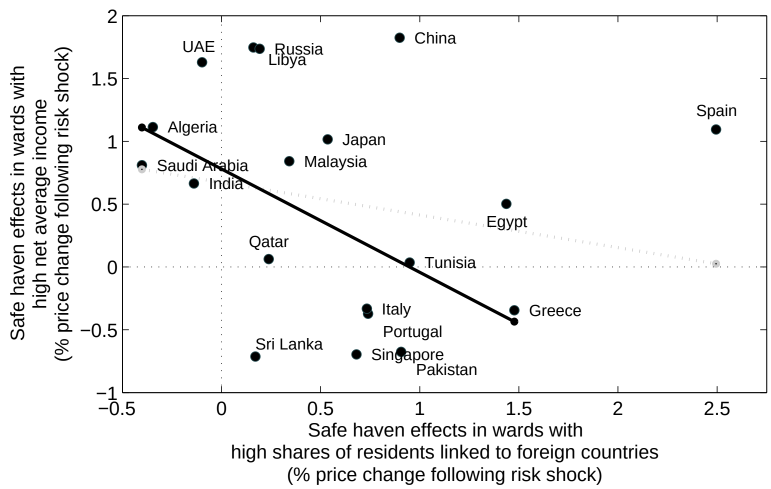
This figure documents the cross-sectional patterns of the estimated coefficients identifying safe haven effects in the London housing market. On the horizontal axes, we report the point estimates of the coefficients γ_0^k from country-by-country versions of equation (3):

$$\ln P_{i,t} = \alpha + \beta \mathbb{X}_i + \phi_w + \delta_t + \left(\gamma_0^k f_w^k + \gamma_1^k y_w\right) z_{t-1}^k + u_{i,t},$$

where f_w^k are the shares of people in ward w born in country k and y_w is average net income. z_t is a 1-year moving average of the ICRG indexes of political risk. On the vertical axes, we report the corresponding point estimates of the coefficients γ_1^k, for each country k. The lines indicate univariate cross-country fitted values. The black line is generated by excluding Spain from the sample.

Panel A

Registry dataset



Panel B

Loans dataset

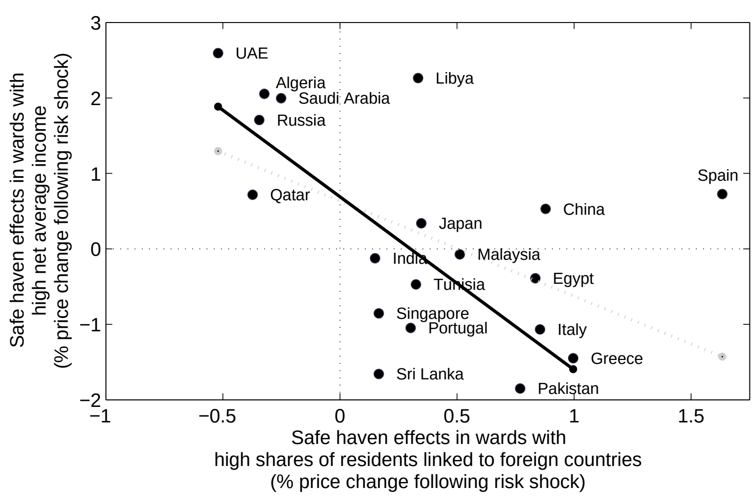


Figure 7
Safe haven effects across price categories

This figure reports the estimated coefficients $\gamma_0^{k,\eta}$ and $\gamma_1^{k,\eta}$ from the following hedonic regression:

$$\ln P_{i,t} = \alpha + \beta \mathbb{X}_i + \phi_w + \delta_t + \sum_{\eta=1}^{3} \sum_{k \in K} \left(\gamma_0^{k,\eta} f_w^k + \gamma_1^{k,\eta} y_w \right) z_t^k + u_{i,t},$$

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-1}^k and price category dummies. In this specification, z_{t-1}^k is a lagged monthly 1-year moving average of the ICRG risk indicator. All right-hand side variables are normalized by subtracting the in-sample mean and dividing by the standard deviation. We report absolute values of all estimated coefficients and indicate negative values by using light shading. Statistical significance is reported through error bars, indicating 95% confidence intervals. The standard errors are clustered at the ward-time level.

Panel A
Registry dataset

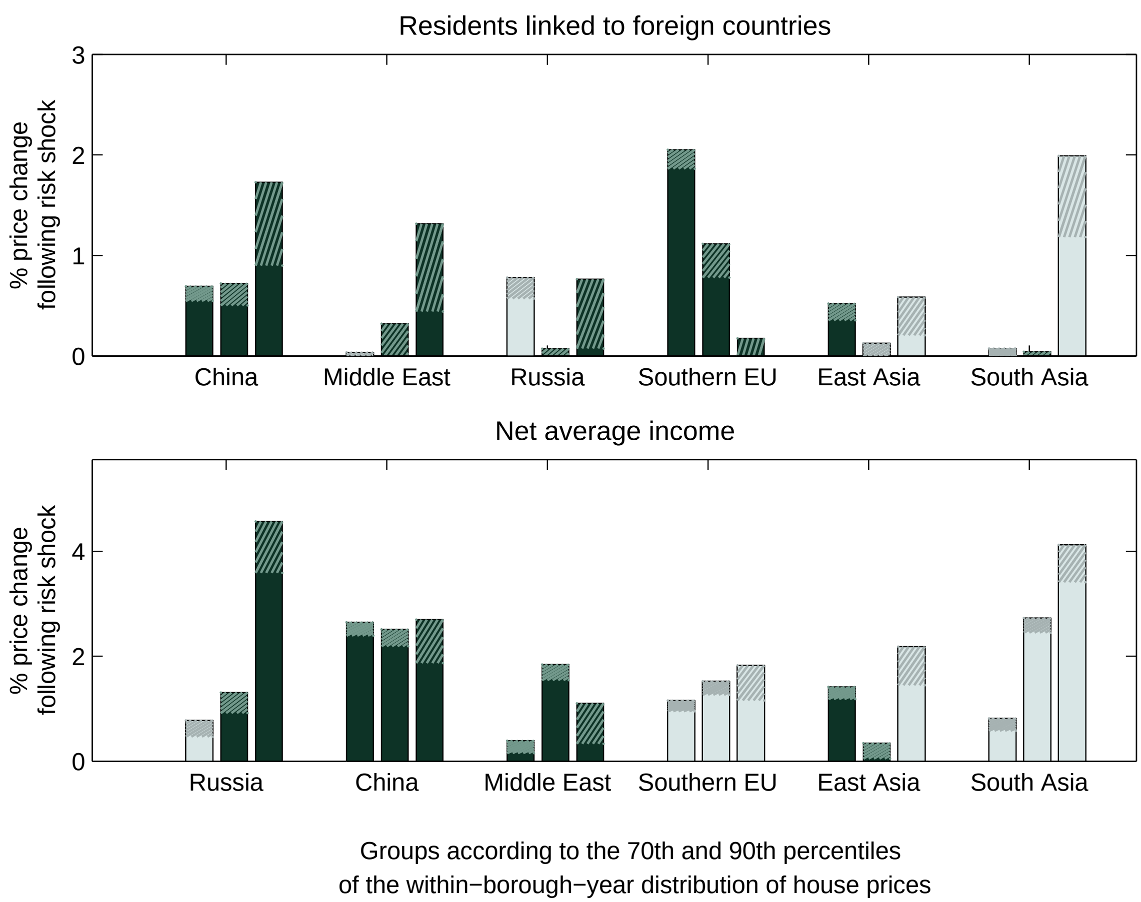


Figure 7
Safe haven effects across price categories
(continued)

Panel B
Loans dataset

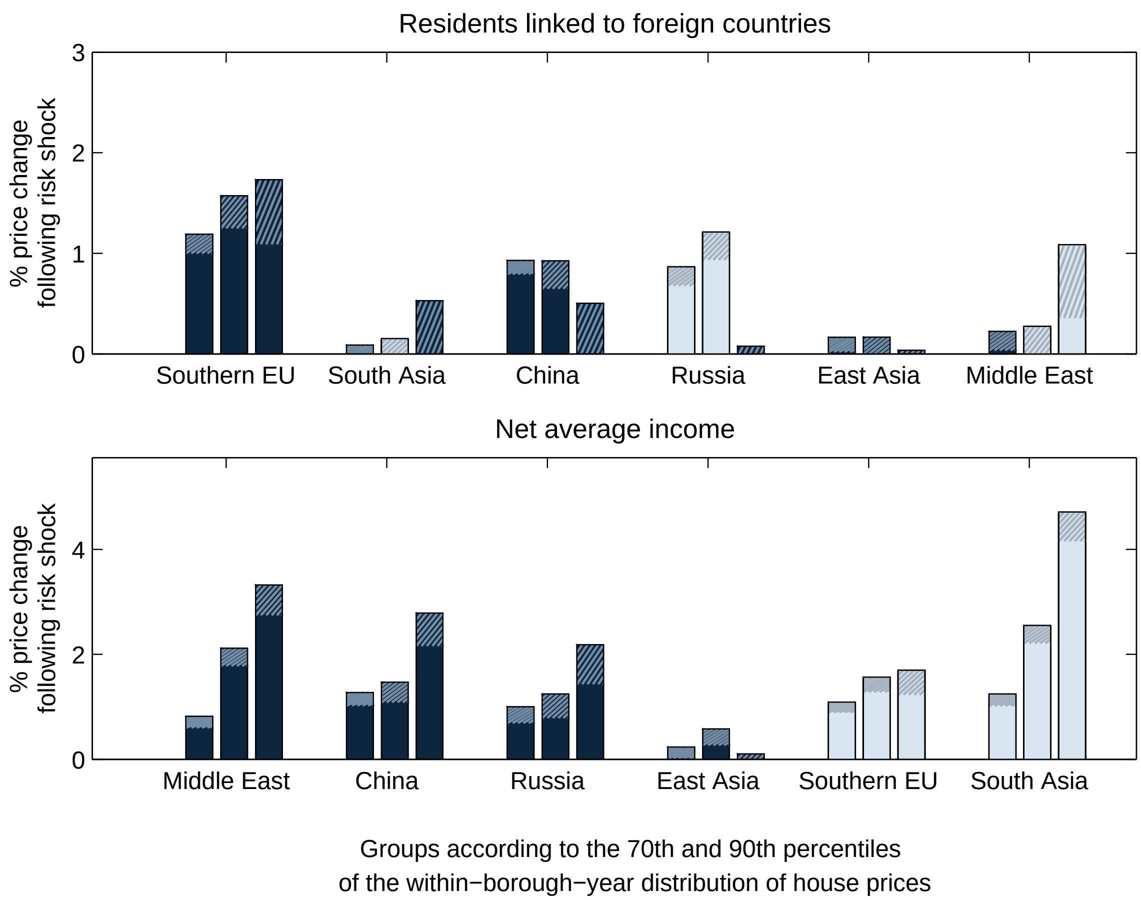


Figure 8

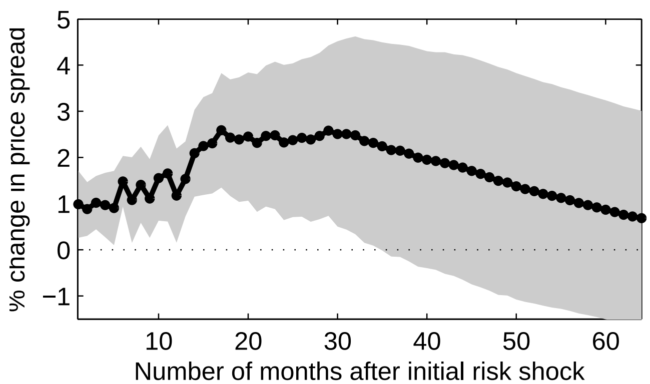
The price impact of safe haven flows: time series dynamics

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

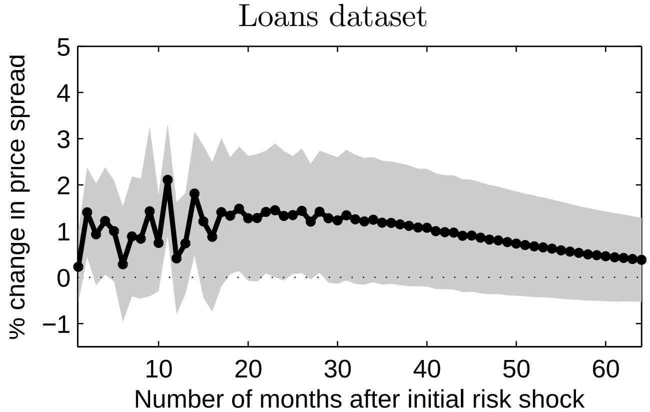
$$s_{k,t} = \mu_k + \sum_{q=1}^{Q} \rho_q s_{k,t-q} + \sum_{q=1}^{Q} \zeta_q z_{k,t-q} + u_{k,t}, \quad \text{and} \quad z_{k,t} = \theta_k + \sum_{q=1}^{Q} \pi_q z_{k,t-q} + \varepsilon_{k,t},$$

where $s_{k,t}$ is the price spread in period t between the top and bottom 20% of wards with respect to the share of people born in country k and $z_{k,t}$ is the level of the ICRG index of political risk. Here, we consider the case $Q = 18$ months. The shaded areas indicate 95% confidence intervals, based on Driscoll-Kraay (1998) standard errors with a maximum lag length of 12 months.

Registry dataset



Loans dataset



Online Appendix for

Preferred Habitats and Safe-Haven Effects: Evidence from the London Housing Market

Table A.1

Data availability, risk indicators

This 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 - 2012	.	.
China	1995 - 2012	1995 - 2012	1995 - 2012
Egypt	1995 - 2012	.	.
Greece	1995 - 2012	1998 - 2012	.
India	1995 - 2012	1996 - 2012	2003 - 2012
Italy	1995 - 2012	1995 - 2012	1997 - 2012
Japan	1995 - 2012	1995 - 2012	.
Libya	1995 - 2012	.	.
Malaysia	1995 - 2012	1996 - 2012	.
Pakistan	1995 - 2012	.	.
Portugal	1995 - 2012	1995 - 2012	.
Qatar	1995 - 2012	.	.
Russia	1995 - 2012	1999 - 2012	.
Saudi Arabia	1995 - 2012	.	.
Singapore	1995 - 2012	1998 - 2012	.
Spain	1995 - 2012	1996 - 2012	2001 - 2012
Sri Lanka	1995 - 2012	.	.
Tunisia	1995 - 2012	.	.
United Arab Emirates	1995 - 2012	.	.

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
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)
Shares of foreign-born people	Number of people (Count) - born in the UK, China, Japan, Malaysia, Singapore, India, Pakistan, Sri Lanka, Greece, Italy, Portugal or Spain	Country of Birth, 2001 (UV08)
	Number of people (Count) - speaking Arabic or Russian	Main Language (detailed), 2011 (QS204EW)

Table A.3

Summary statistics across wards

The table reports mean values for selected variables, calculated for the wards in the top quarter 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 25% of wards with highest shares of people born in:	Southern Europe	102.98	63.43	572.31	0.68	10.31	56.30
	China	81.06	51.15	539.68	0.76	8.75	59.53
	East Asia	81.01	52.89	632.37	0.85	11.72	54.72
	South Asia	64.44	28.52	533.72	0.95	6.83	59.14
	Russia	87.55	58.18	568.53	0.71	10.13	57.86
	Middle East	87.60	51.80	556.99	0.79	8.78	55.89
	UK	39.02	16.87	551.54	1.14	5.15	59.09
Full sample of wards		70.68	39.41	546.14	0.88	7.62	59.88

Table A.4

Sources of bond yield data

The bond yield data is retrieved through Datastream. The second column reports the codes assigned to each of the variables we use, the third column reports the title descriptions of the series and the fourth column indicates the sources of the data.

Country	Variable code	Description	Source
China	CHXRLG..R	Interest Rate Government Securities: 10 Year Nominal Par Yield	Oxford Economics
Greece	GROIR080R	Yield 10 Year Government Bonds	OECD Main Economic Indicators
India	INGBOND	Treasury Bond Yield, 10 Year	Reserve Bank of India
Italy	ITOIR080R	Yield 10 Year Government Benchmark Securities	OECD Main Economic Indicators
Japan	JPGBOND	Interest-Bearing Government Bonds - 10 Year	OECD Main Economic Indicators
Malaysia	MYGBOND	Government Bond Yield - 10 Year	Central Bank of Malaysia
Portugal	PTGBOND	Benchmark Bond Redemption Yield - 10 Year	Banco de Portugal
Russia	RSOIR080R	Long-Term Government Bond Yields / 10 Years	OECD Main Economic Indicators
Singapore	SPGBOND	10 Year Government Bond Yield	Monetary Authority, Singapore
Spain	ESGBOND	Central Government Bond - 10 Year Yield	Bank of Spain
UK	UKAMNZC	British Government Securities, 10 Year	Bank of England

Table A.5

Country-by-country estimation: Isolating the local component of political risk

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

$$\ln P_{i,t} = \alpha + \beta \mathbb{X}_i + \phi_w + \delta_t + (\gamma_0^k f_w^k + \gamma_1^k y_w + \gamma_2^k \overline{f}_w)(z_{t-1}^k - \overline{z}_{t-1}^k) + u_{i,t},$$

where $\mathbb{X}_i$ are property-level characteristics. In this specification, z_t^k is a monthly 1-year moving average of each of the ICRG risk indicators and $\overline{f}_w$ is the total ward-level share of people born outside the UK. $\overline{z}_t^k$ is a 1-year moving average of the fitted value from a regression of the country-specific ICRG index on the first principal component of our full set of political risk indexes from 19 countries. The coefficients in column (1) correspond to the case where we do not control for ward-level foreign population shares (i.e. $\gamma_2^k = 0$). All variables are normalized by subtracting the in-sample mean and dividing by the standard deviation. *, **, *** denote statistical significance at the 10%, 5%, and 1% level respectively. The standard errors are clustered at the ward-time level.

	Registry dataset		Loans dataset	
	(1)	(2)	(1)	(2)
Spain	2.02***	2.16***	1.14***	1.38***
Greece	0.84***	0.27**	0.78***	0.17
Egypt	1.07***	0.38***	0.66***	-0.12
Tunisia	-0.06	0.02	-0.28***	0.02
Pakistan	1.18***	1.09***	0.84***	0.19**
China	0.77***	0.32***	0.79***	0.07
Portugal	-0.05	-0.37***	-0.14*	-0.29***
Italy	-0.20***	-0.34***	0.29***	0.07
Singapore	-0.10	0.09	-0.41***	-0.06
Japan	0.46***	-0.01	0.30***	-0.13
Malaysia	0.17***	-0.09	0.39***	-0.35***
Qatar	-0.45***	0.20**	-0.76***	0.47***
Russia	-0.22***	-0.14	-0.75***	-0.46***
Sri Lanka	0.71***	0.99***	0.39***	0.45***
Libya	-0.01	0.39***	0.21***	0.91***
UAE	-0.59***	0.38***	-0.71***	0.96***
India	0.31***	0.76***	0.46***	0.66***
Algeria	-0.33***	0.17*	-0.31***	0.85***
Saudi Arabia	-0.76***	0.14	-0.47***	0.93***
Local component	Yes	Yes	Yes	Yes
Control for total foreign-born people share	No	Yes	No	Yes

Table A.6

Safe haven effects: global risk and total foreign population shares

This table reports estimated coefficients γ_0 and γ_1 from the following hedonic regression:

$$\ln P_{i,t} = \alpha + \beta \mathbb{X}_i + \phi_w + \delta_t + (\gamma_0 \overline{f}_w + \gamma_1 y_w)\overline{z}_{t-1} + u_{i,t},$$

where $\mathbb{X}_i$ are property-level characteristics. In this specification, $\overline{f}_w$ is the total ward-level share of people born outside the UK and $\overline{z}_t$ is a monthly 1-year GDP-weighted moving average of the ICRG risk indicators. All variables are normalized by subtracting the in-sample mean and dividing by the standard deviation. *, **, *** denote statistical significance at the 10%, 5%, and 1% level respectively. In parentheses, we report standard errors, clustered at the ward-time level.

	(1)	(2)	(3)	(4)
Registry dataset				
Safe-haven effect				
- in wards with strong links	2.00***	2.23***	2.31***	0.54**
to foreign countries	(0.43)	(0.21)	(0.21)	(0.26)
- in wards with high net			1.10***	0.77***
average income			(0.19)	(0.23)
Loans dataset				
Safe-haven effect				
- in wards with strong links	1.71***	2.40***	2.33***	1.04***
to foreign countries	(0.29)	(0.13)	(0.13)	(0.18)
- in wards with high net			-0.38***	-0.10
average income			(0.14)	(0.14)
Monthly time fixed effects	Yes	Yes	Yes	Yes
Ward-level fixed effects	No	Yes	Yes	Yes
Borough-year fixed effects	No	No	No	Yes

Table A.7

Effects of political risk on ward-level prices

This table reports the coefficients λ^j from the following panel regression specifications:

$$(1) \quad : \quad \ln P^j_{k,t} = \alpha^j + \mu^j_k + \lambda^j \overline{z}_{k,t-1} + u^j_{k,t}, \quad \text{for } j \in \{high, low\}$$

$$(2) \quad : \quad \ln P^j_{k,t} = \alpha^j + \mu^j_k + \delta^j_t + \lambda^j \overline{z}_{k,t-1} + u^j_{k,t}, \quad \text{for } j \in \{high, low\}$$

$$(3) \quad : \quad \ln P^j_{k,t} = \alpha^j + \mu^j_k + \rho^j \ln P^j_{k,t-1} + \lambda^j \overline{z}_{k,t-1} + u^j_{k,t}, \quad \text{for } j \in \{high, low\}$$

where $P^{high}_{k,t}$ is the average price in period t in the top 20% wards with the highest shares of people born in country k and $P^{low}_{k,t}$ is the average price in the respective bottom 20% of wards. $z_{k,t}$ is the ICRG indicator of political risk in country k and $\overline{z}_{k,t}$ is a 1-year moving average thereof. *, **, *** denote statistical significance at the 10%, 5%, and 1% level respectively. In parentheses, we report Driscoll-Kraay (1998) standard errors with a lag length of 12 months.

	Registry dataset			Loans dataset		
	(1)	(2)	(3)	(1)	(2)	(3)
Top 20% of wards	10.89*	0.82***	0.16**	10.10	0.88***	0.11
	(6.35)	(0.25)	(0.08)	(6.56)	(0.24)	(0.15)
Bottom 20% of wards	7.74	-0.33	0.01	7.67	-0.62***	-0.05
	(6.11)	(0.23)	(0.06)	(6.69)	(0.21)	(0.10)
Time fixed effects	No	Yes	No	No	Yes	No
Autoregressive component	No	No	Yes	No	No	Yes

Figure A.1

Cross-ward distribution of the shares of foreign-born people

This figure shows the distribution of the shares of people born in respective countries, across the set of 624 London wards. We report the shares in percent of the total ward population. For Russia and the Middle East, we consider the number of people speaking Russian and Arabic, respectively.

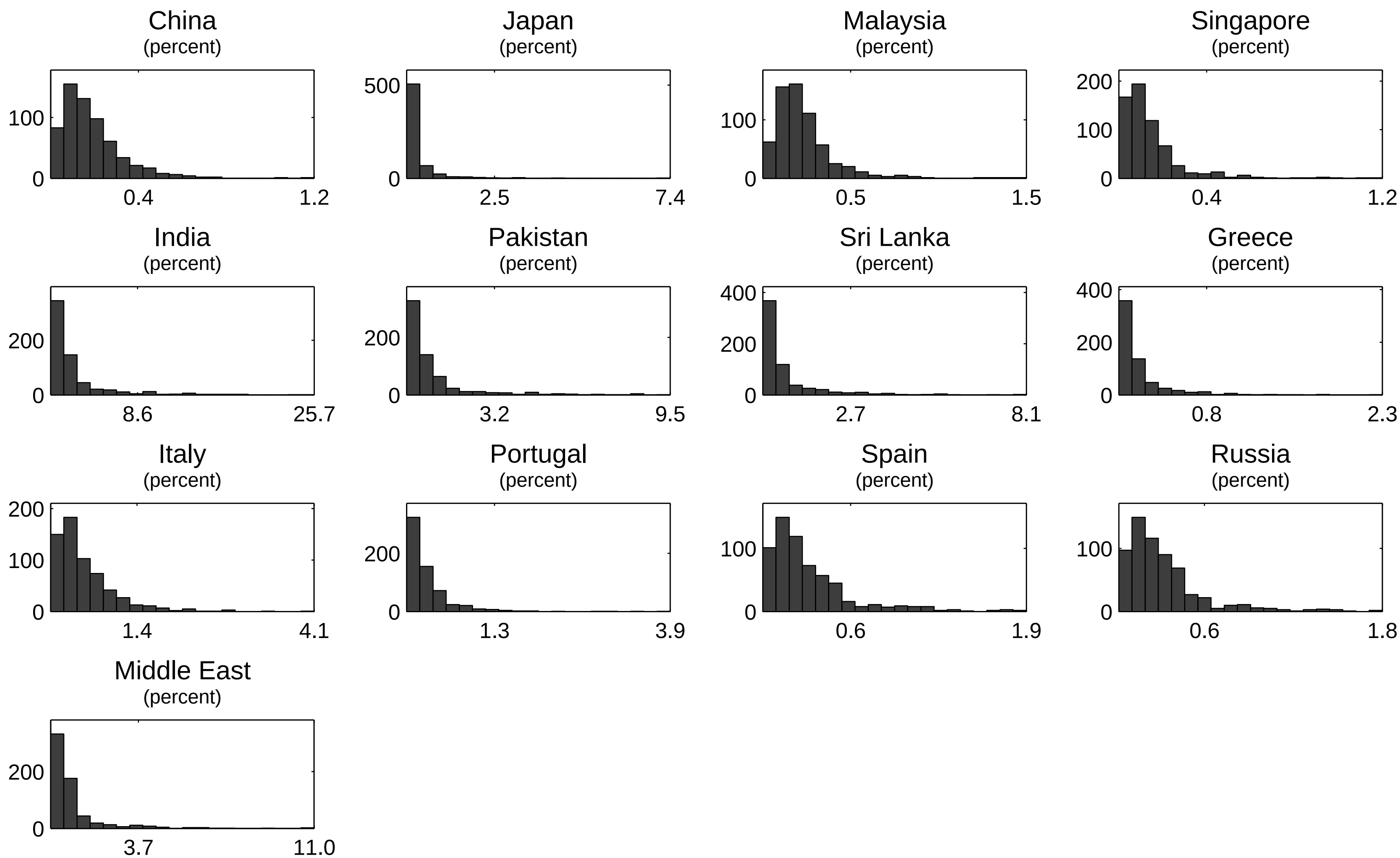


Figure A.2
Cross-ward distribution of demographic and socioeconomic variables

This 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.

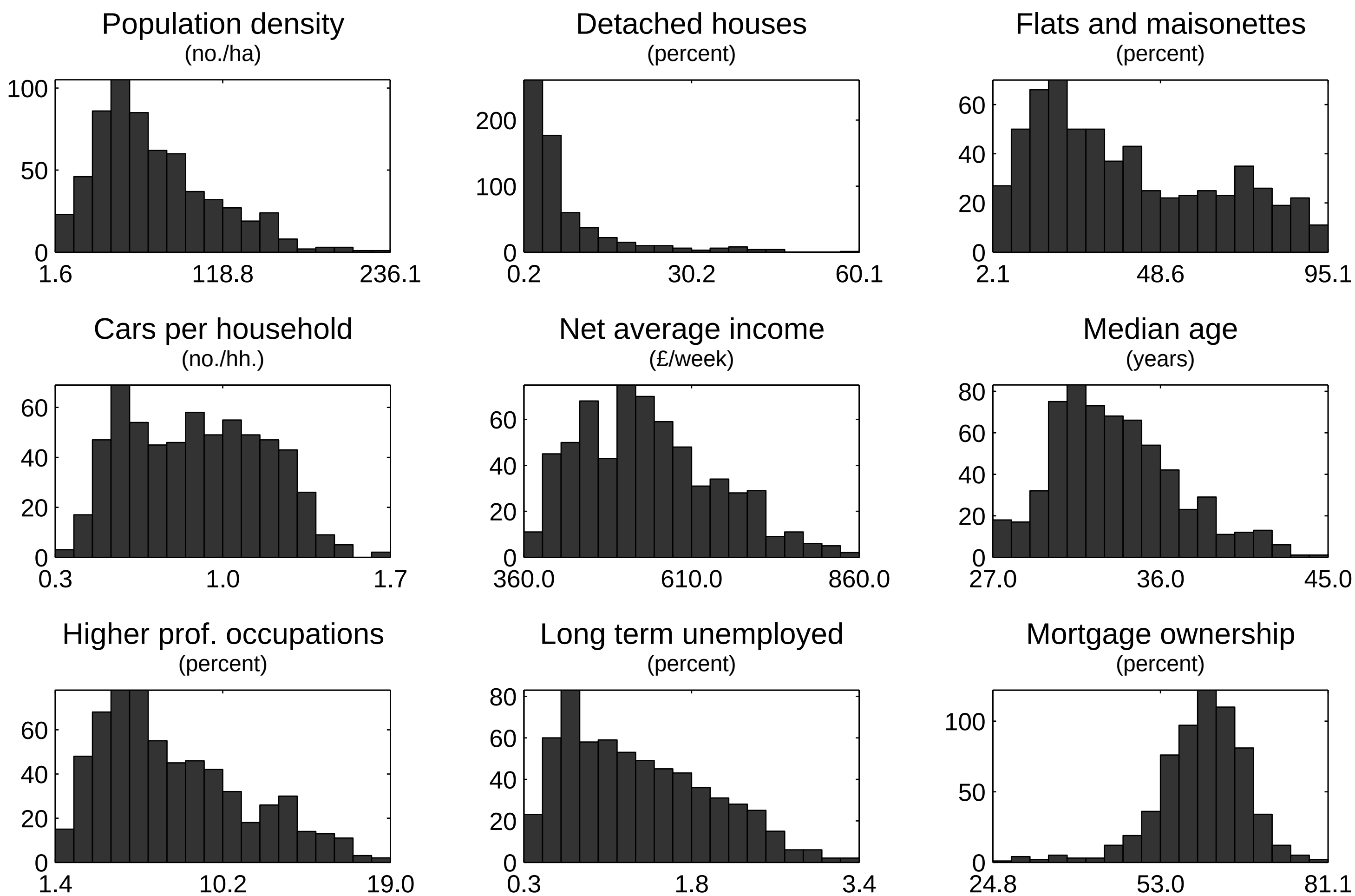


Figure A.3

Time series of 10-year government bond yields

This figure shows the evolution through time of yield rates corresponding to benchmark government bonds with maturities of 10 years. We report the numbers in percentage points.

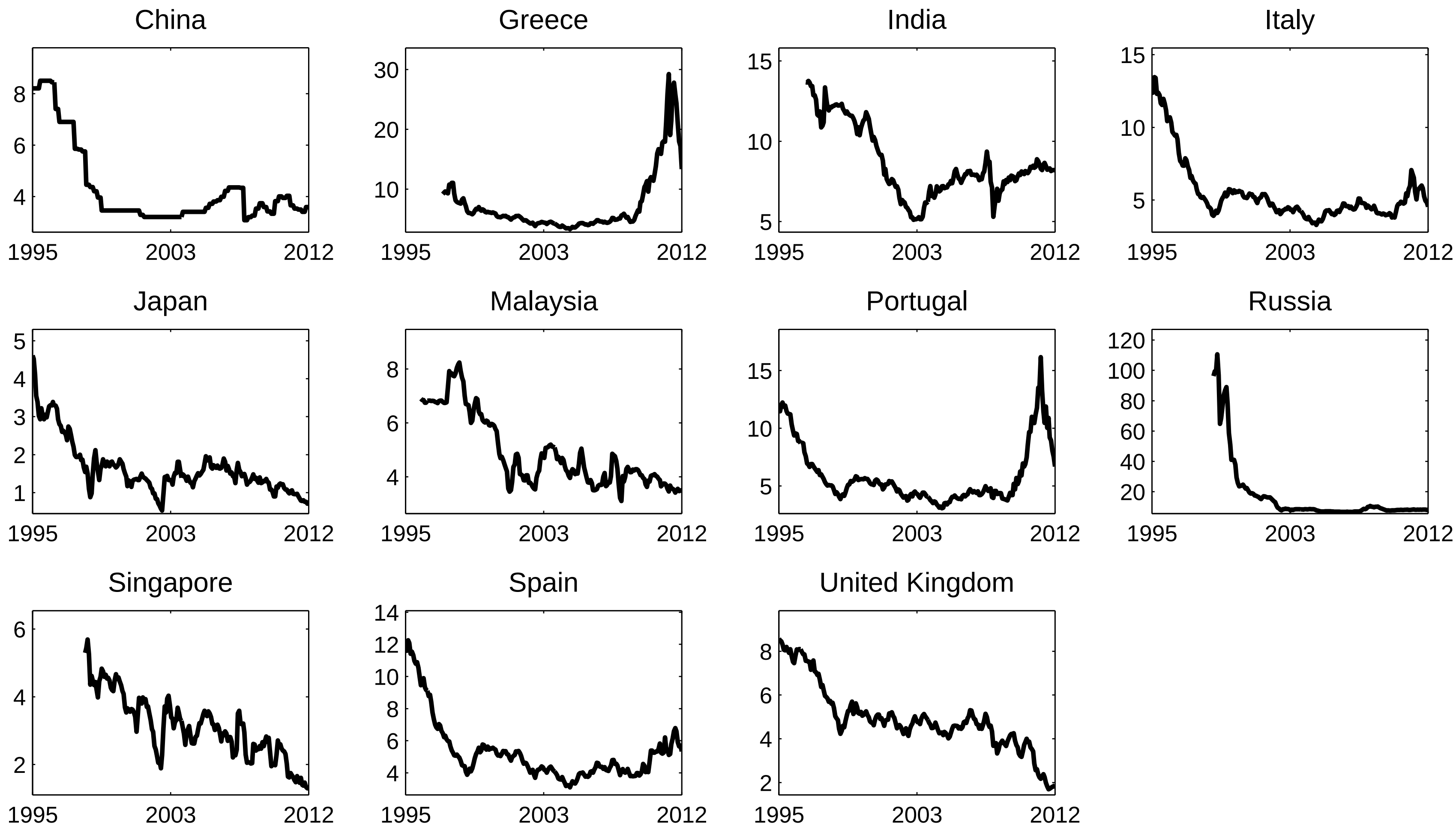


Figure A.4

ICRG indexes of political risk

This figure shows the evolution through time of the ICRG measures of political risk. In raw form, the indexes range from 0 to 100, with 0 indicating the highest possible risk. We replace them with 100 minus the original values so that high levels of the indexes indicate high levels of risk and vice versa. We normalize the indexes by subtracting the mean and dividing by the standard deviation.

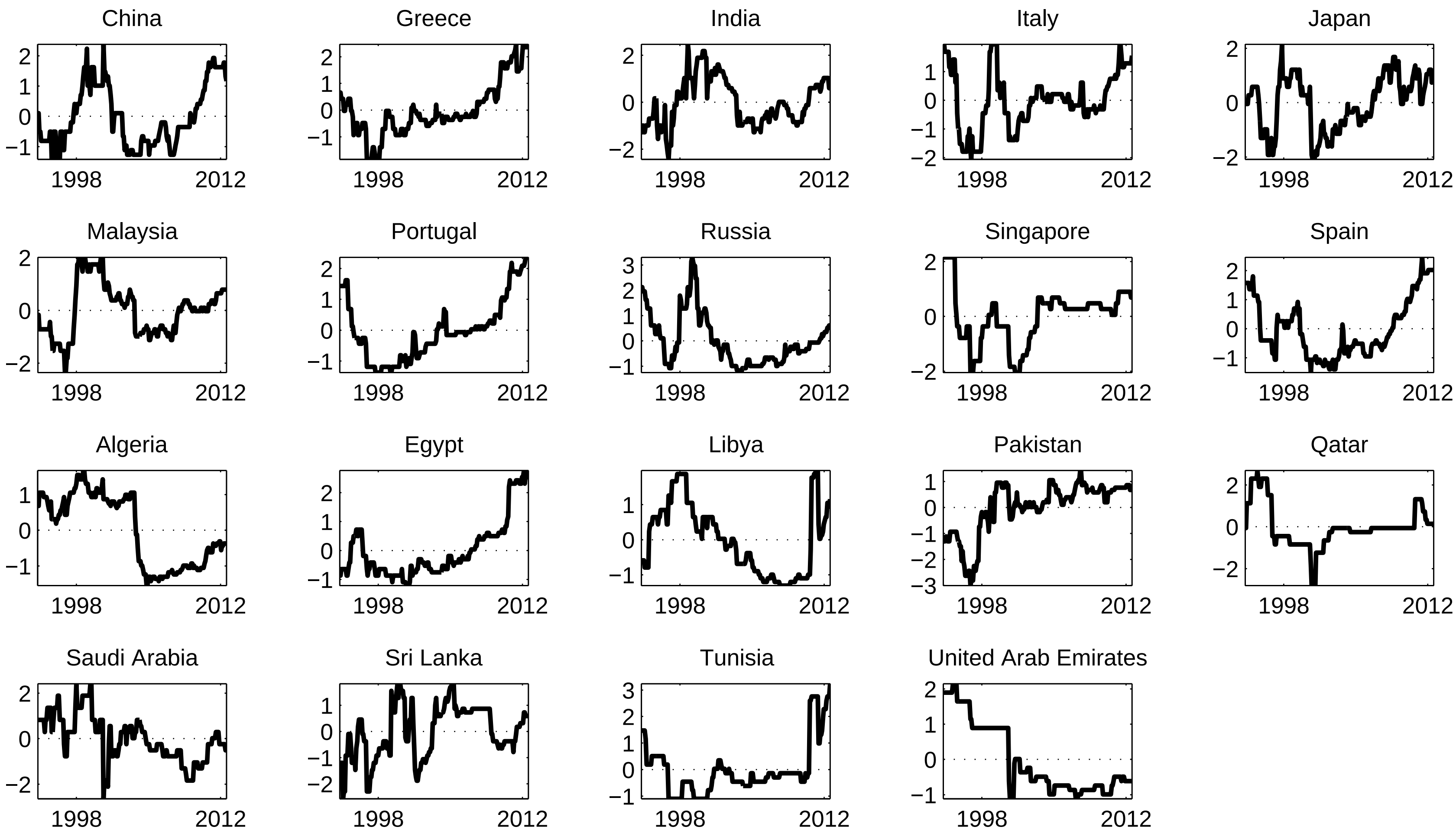


Figure A.5
Yield data: China

This 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.

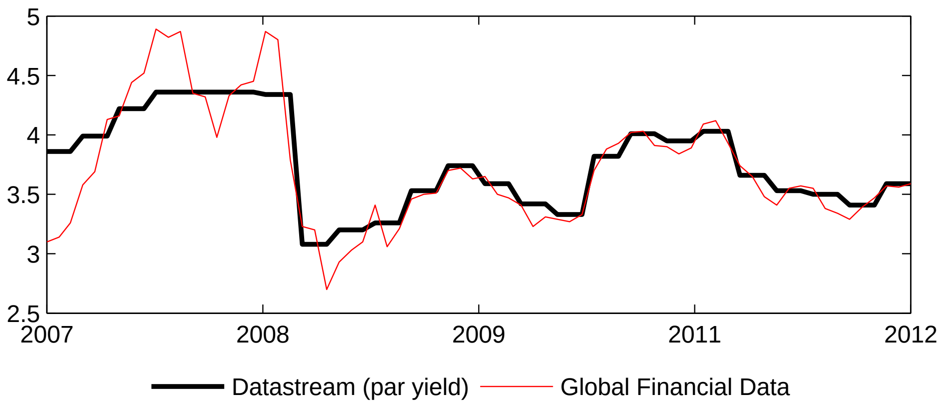


Figure A.6
Principal components analysis

This figure reports the estimated first principal component of the ICRG measure of political risk, across our set of 19 countries.

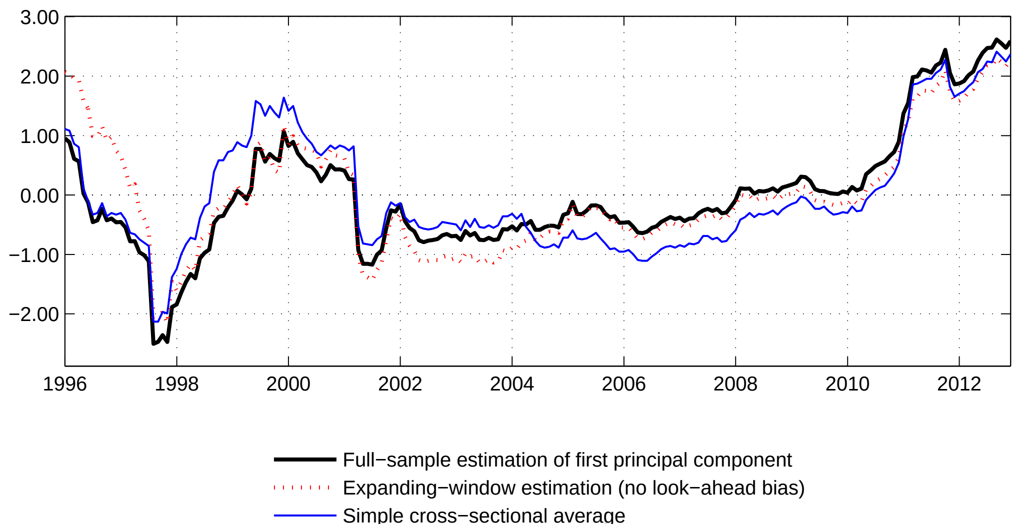


Figure A.7
Relationship between the ICRG and EPU indexes

This 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.

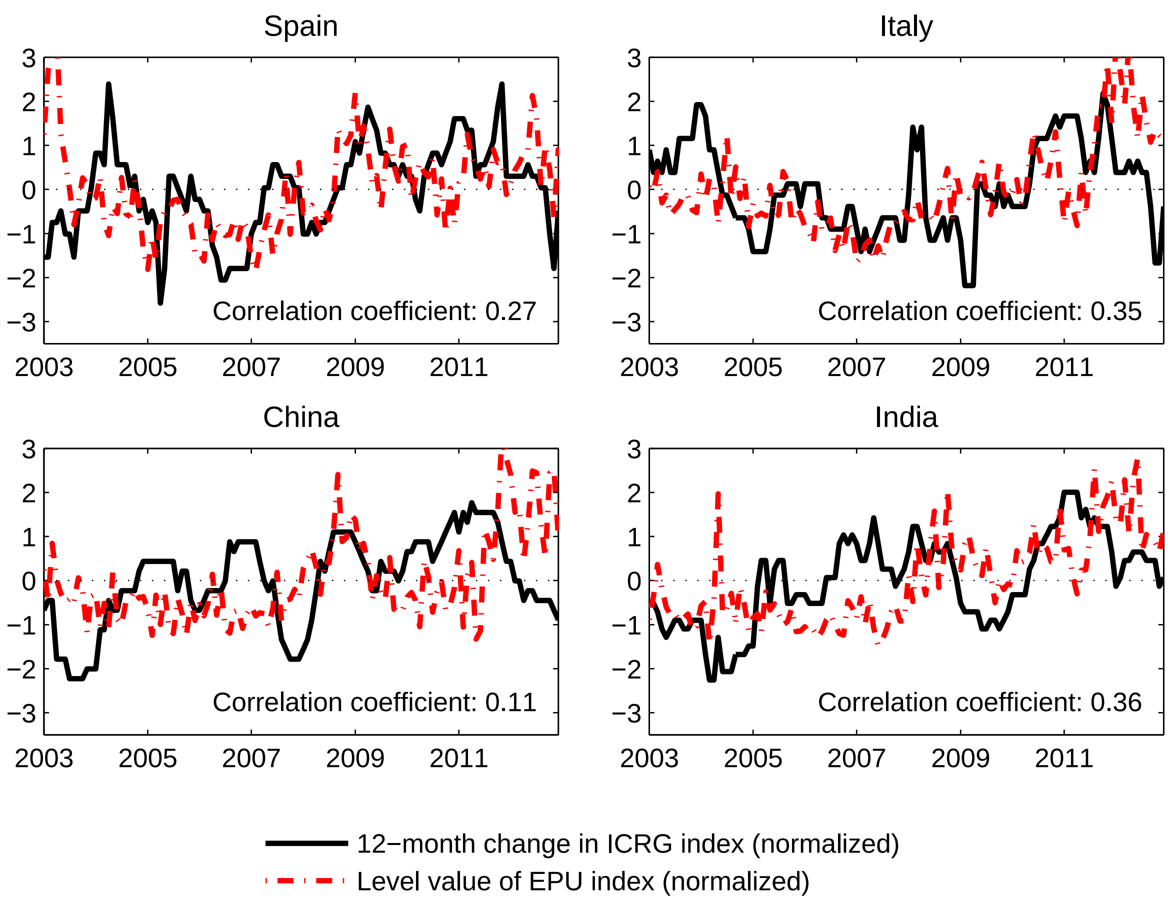


Figure A.8

Analysis at the level of world regions: effects of political risk

This figure reports the estimated coefficients $\gamma_0^{k,\eta}$ and $\gamma_1^{k,\eta}$ from the following hedonic regression:

$$\ln P_{i,t} = \alpha + \beta \mathbb{X}_i + \phi_w + \delta_t + \sum_{\eta=2}^{3} \sum_{k \in K} (\gamma_0^{\eta,k} f_w^k + \gamma_1^{\eta,k} y_w) + u_{i,t}.$$

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 category in which the ICRG index of political risk falls in period $t-12$, for each world region k. The thresholds which determine the category are given by the 33rd and 66th percentiles of the evolution of risk through time. The coefficients γ correspond thus to an interaction term between the ward-level characteristics f_w^k or y_w, and the risk category dummies. In this specification, the low risk category is taken as a reference, corresponding to the values of the ICRG risk indicator which fall below the 33rd time-series percentile. All right-hand side variables are normalized by subtracting the in-sample mean and dividing by the standard deviation. We report absolute values of all estimated coefficients and indicate negative values by using light shading. Statistical significance is reported through error bars, indicating 95% confidence intervals. The standard errors are clustered at the ward-time level.

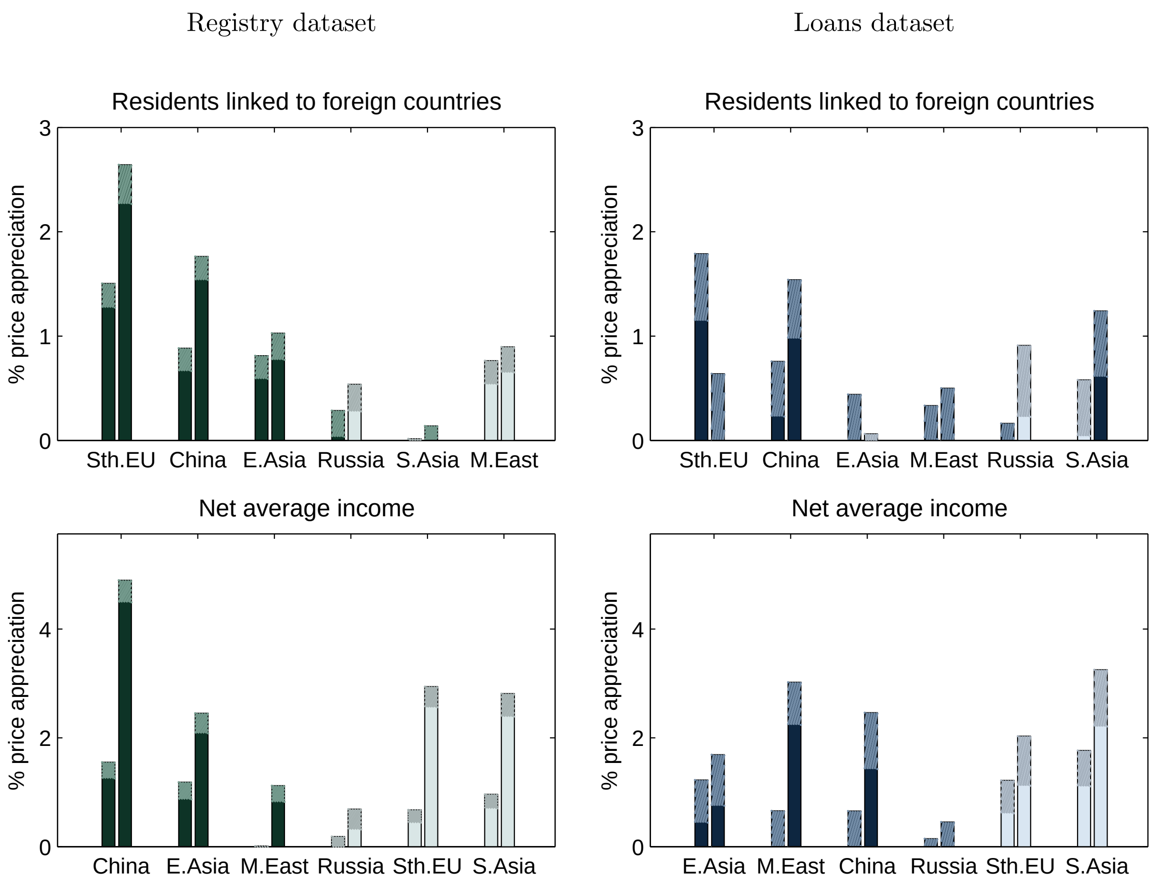


Figure A.9

Evolution of ward-level prices

This figure reports the evolution of average house prices in the Registry dataset of housing transactions, for a period starting in January 2006 and ending in December 2012. The two lines show the different evolution of prices in the top (bottom) 20% of wards which have the highest (lowest) shares of people born in the respective world regions.

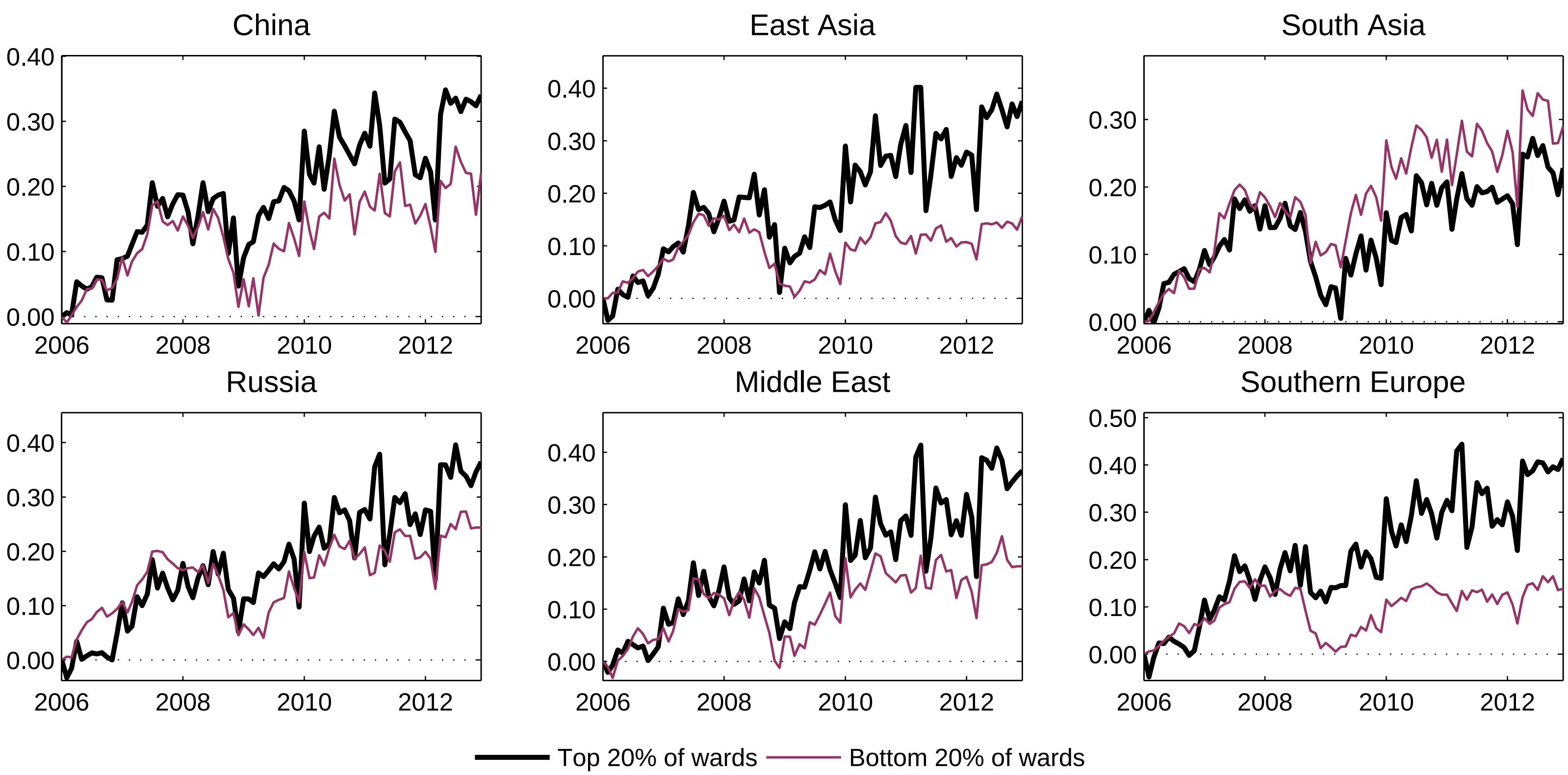
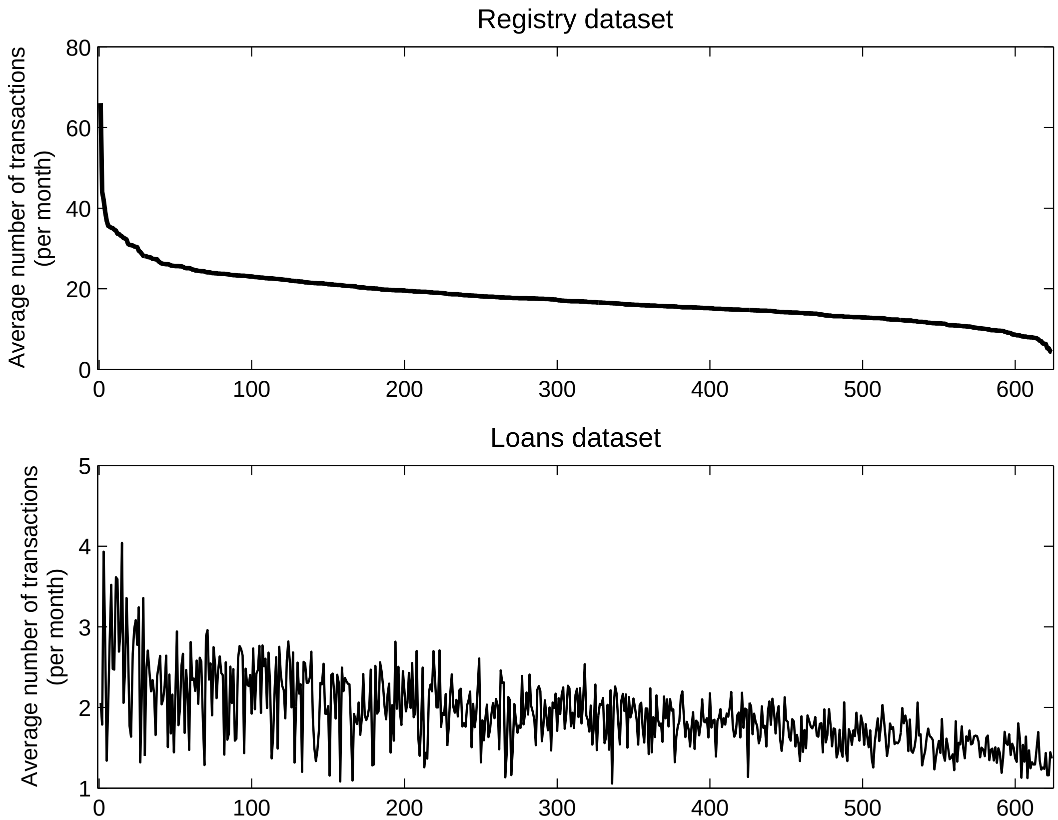


Figure A.10
Distribution of average monthly ward-level transaction volumes

This figure reports average transaction volumes per month in each of the 624 wards. In both subplots, the wards are sorted in decreasing order of transaction volumes recorded in the Registry dataset.



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:

$$\Delta f^k_{w,2011} = \alpha + \rho^k f^k_{w,2001} + \pi^k_1 \Delta \ln P_{w,2001} + e_{w,2011}, \quad (1)$$

$$\Delta f^k_{w,2011} = \alpha + \rho^k f^k_{w,2001} + \pi^k_2 \Delta \ln \bar{P}^k_{w,2001} + \pi^k_3 \Delta u_{w,2001} + e_{w,2011}. \quad (2)$$

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. $\Delta \ln \bar{P}^k_{w,2001}$ and $\Delta u_{w,2001}$ are constructed by controlling for variation in price-impacting hedonic characteristics of properties of the ward level. $\Delta \ln \bar{P}^k_{w,2001}$ 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^k_3 towards zero.

Together, specifications (1) and (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.11 shows estimates of equations (1) and (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.11

Relationship between house prices and immigration shares

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

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

In the above 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. $\Delta u_{w,2001}$ is the residual price change in ward w, constructed by controlling for variation in price-impacting hedonic characteristics of properties at the ward level. $\Delta \ln \bar{P}_{w,2001}^k$ 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. Statistical significance is reported through error bars, indicating 95% confidence intervals. The estimated standard errors are White heteroskedasticity-robust.