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Cristian Badarinza and Tarun Ramadorai

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**Cristian Badarinza, University of Oxford**
**Tarun Ramadorai, University of Oxford and CEPR**

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Centre for Economic Policy Research
77 Bastwick Street, London EC1V 3PZ, UK
Tel: (44 20) 7183 8801, Fax: (44 20) 7183 8820
Email: cepr@cepr.org, Website: www.cepr.org

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# ABSTRACT

## Home Away From Home? Safe Haven Effects and London House Prices*

Historical time-series data is short relative to the frequency of political and economic crises. This makes it difficult to use pure time-series methods to identify the impacts of safe haven demand on asset prices, in the face of confounding effects from a wide range of alternative drivers. We present a new method to identify safe-haven effects which relies on combining the cross-section of asset prices with time-series measures of economic and political risk. We employ this strategy on large databases of historical housing transactions in London, and show that economic and political risk in Southern Europe, China, the Middle East, Russia, and South Asia is an important factor in explaining the dynamics of London house prices over the past two decades.

JEL Classification: C53, D80, E47, F21, G12 and G15
Keywords: economic risk, hedonic pricing, house prices, London, political risk and safe-haven

Cristian Badarinza
Saïd Business School and Oxford-Man Institute of Quantitative Finance
University of Oxford
Park End Street
Oxford OX1 1HP

Email: cristian.badarinza@sbs.ox.ac.uk

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Tarun Ramadorai
Saïd Business School and Oxford-Man Institute of Quantitative Finance
University of Oxford
Park End Street
Oxford OX1 1HP

Email: tarun.ramadorai@sbs.ox.ac.uk

For further Discussion Papers by this author see: www.cepr.org/pubs/new-dps/dplist.asp?authorid=159984

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# 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 long been used by press commentators and academics alike to explain unusual variation during crises in the prices of a long list of assets, including gold, international real estate, currencies, sovereign debt, and international equities.[1] A better 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 stems from a long-acknowledged issue, which is that safe-haven demand effects generally appear during both country-specific and global crises. Historical time-series data is short relative to the frequency of crises, and in the time-series, distinguishing shocks to risk-aversion from a range of other changes can be difficult. As a result, it can be problematic to convincingly attribute asset price movements to price pressure from safe-haven demand rather than to a host of other candidate explanations, including movements in price-relevant information, or the impact of institutional frictions. We help to resolve this issue by employing a new strategy that is grounded in agents' microeconomic motivations, and focuses on identifying safe-haven effects using the cross-section of asset prices rather than the time-series. We view this methodology as being more generally applicable in identifying safe-haven effects in a range of asset prices.

[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).

Our empirical contribution is to employ this strategy on large databases of historical housing transactions in London, a city which has been the focus of substantial recent attention for its unusual house price appreciation during the crisis. To explain this phenomenon, numerous media commentators have invoked the possibility of price pressure from safe-haven demand effects from overseas purchasers.[2] We provide evidence that safe-haven demand effects from Southern Europe, China, the Middle East, Russia, and South Asia are indeed important factors in explaining the dynamics of London house prices. These results help to shed light on a current phenomenon of wide public interest, and suggest a partial explanation for the famous Lucas (1990) puzzle about the reason for the limited flow of capital from relatively rich to relatively poor countries.[3] Our results suggest that there is an "uphill" flow of capital from poor to rich countries that is driven by the desire of a subset of poor-country residents to insure themselves against economic and political uncertainty in their home countries.

Our identification strategy is relatively simple to explain. We sub-divide London into smaller geographical areas, and enumerate the strength of the links of each such London area with specific foreign countries. When there is an increase in political or economic uncertainty 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 cross-sectional identification strategy may be useful here. 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

[2]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.

[3]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.

London, we make the identifying assumption 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 identifying assumption is correct, following heightened uncertainty in Greece, we would expect to see relatively higher price increases in these *specific* areas of London, over and above increases in *all* London house prices. In our specifications, the rates of price appreciation of 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 increases in political and economic uncertainty in Southern Europe, China, and East Asia on the prices of houses in London areas with high shares of people born in these regions.

We believe this identification strategy is sensible for at least two reasons. First, at least some portion of safe-haven housing demand is likely to be associated in the minds of purchasers with 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 a natural factor affecting demand for real estate. Note that this line of reasoning does not necessarily require actual future immigration to materialize, just that it is a factor influencing property selection in the minds of the purchasers. That said, we do find suggestive evidence that the patterns of growth in ward-level immigration shares in our data appear consistent with this line of reasoning.

Second, even if such an implicit or explicit future immigration motive is not a factor at the time of purchase, there may be social network effects that help to lower informational asymmetries, which arise from the presence of foreign-origin settlement in particular London wards. Here we have in mind both direct communication between foreign-origin local London residents and overseas safe haven purchasers, as well as the presence 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 wards.[4]

Our identification strategy is not restricted to the use of any single variable representing the strength of the links between London areas and foreign countries. Indeed, we view our approach as generalizable to any situation in which investor demands can be matched to specific preferred habitat assets. 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 moving capital 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, and using this additional identification, we find that London areas with high average income levels experience unusual price appreciation following increases in political uncertainty in China, the Middle East, East Asia, and Russia. Conversely, London areas with relatively lower average income levels experience price appreciation following turbulence in Southern Europe and South Asia. These findings are over and above the foreign-born share effects described earlier.

The economic magnitudes that we estimate are large. In London areas with a one standard deviation higher share of people originating from a particular country, we find that prices are 40 basis points higher in months following a one standard deviation increase in that country's political uncertainty in the previous year. Moreover, London areas with one standard deviation higher average incomes appreciate by 30 basis points on average in response to political uncertainty shocks. These impacts are likely underestimates, as they do not include safe-haven demand effects that have

[4]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.

homogenous impacts across London areas. Moreover, 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 pioneered by Rosen (1974).[5]

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 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), and Jotikasthira et al. (2012).

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 US). On the connection between foreign capital flows and local house prices, Aizenman and Jinjarak (2009) find substantial co-movement between current account deficits and real house prices using a panel of OECD countries, and Jinjarak and Sheffrin (2011) analyze the relationship between capital account surpluses and the real estate market in a time-series VAR setup, finding statistically significant UK house price appreciation following capital inflow shocks. Our findings are consistent with these results, but

[5] 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)).

are distinguished by a focus on the identification of safe-haven effects using a new cross-sectional approach, and implementation using transactions-level data.

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.

# 2 Methodology

## 2.1 Identifying Safe Haven Effects in London House Prices

Consider a standard hedonic pricing model for properties in London:

$$\ln P_{l,w,t} = \alpha + \beta \mathbb{X}_l + \Pi_{w,t} + u_{l,w,t}, \tag{1}$$

Here, $P_{l,w,t}$ is the price of property $l$, physically located in ward $w$, measured at time $t$. $\mathbb{X}_l$ is a vector of time-invariant hedonic characteristics for property $l$, and the second component on the right-hand side of equation (1), denoted $\Pi_{w,t}$, represents (unrestricted) ward-time fixed effects. The heart of our empirical approach is to restrict these ward-time fixed effects in an economically meaningful fashion. These restrictions allow us to identify the impacts of foreign political and economic uncertainty (varying across regions of the world and time) on the cross-ward time-variation in London house prices – by associating these 624 wards with specific world regions through the share of the population of each London ward originating from each region.

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 $t-1$.[6] We model $\Pi_{w,t}$ in the following fashion:

$$\Pi_{w,t} = \delta_t + \theta \mathbb{C}_w + \sum_{k \in K} \gamma^k f_w^k z_{t-1}^k. \tag{2}$$

[6] 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.

In equation (2), $\delta_t$ are time fixed effects, $\mathbb{C}_w$ is a vector containing demographic and socioeconomic characteristics of wards, and $f_w^k$ is the fraction of people in ward $w$ who were born in country or world region $k$.[7]

The inclusion of the time fixed effects in equation (2) eliminates common time-variation in London house prices, meaning that the remaining coefficients are purely estimated using cross-ward variation, and importantly, cross-ward variation interacted with time- and cross-country (or cross-world-region) variation in the indicator $z_{t-1}^k$. This strategy means that while safe haven flows may influence London house price appreciation across the board, we do not allow our identification of safe haven effects to be influenced by this source of aggregate variation. Consequently, our estimates are likely an underestimate of the total impact of safe haven flows.

Our identifying assumption is that London wards with relatively high pre-existing shares of population from a specific country $k$ are preferred locations for safe-haven demand emanating from country $k$. If this identifying assumption is correct, $\gamma^k$ will be estimated positive, as we expect to see relatively higher price increases in wards (relative to aggregate increases in all London house prices) with relatively high shares of country $k$-origin population following increases in risk in country $k$.

We also control, in our empirical specification, for the possibility that London wards with relatively high pre-existing shares of country $k$-origin population may have higher or lower prices *on average* than other wards, i.e., we include $f_w^k$ in the set of ward-level characteristics $\mathbb{C}_w$. This is to ensure that $f_w^k$ is not simply capturing other unobservable hedonic characteristics of properties located in particular wards.

[7] 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 consider multiple world regions $k$ simultaneously. The latter specification additionally allows for the possibility that there may be spillovers from political and economic uncertainty in country $j$, even to wards which have high shares of population from country $k \neq j$.

In addition to the identification approach described above, we also consider the possibility that safe-haven demand is concentrated in desirable London regions, driven by high-net worth foreigners moving capital into London real estate in response to uncertainty in their home countries. In order to control for this possibility, we add a ward-level indicator of desirability $y_w$ as a conditioning factor:

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

In our empirical implementation, we use net average income in the ward as our choice of $y_w$, which in our view effectively captures this notion of desirability. Again, we allow for the possibility that wards with higher net average income levels have higher or lower prices on average, i.e. $y_w$ is also included in the set of ward characteristics $\mathbb{C}_w$.

A few notes on our identification strategy and our empirical approach: first, our approach means that the ward-time fixed effects $\Pi_{w,t}$ in equation (1) are restricted in a number of ways. *Pure* time-variation in these ward-time fixed effects is never used to identify our coefficients of interest, as a result of the inclusion of time fixed effects $\delta_t$ in our specifications. This means that we depart from the usual strategy of identifying safe-haven effects in the time-series – where time-samples are short, and the infrequent arrival of crises makes it difficult to pinpoint precise impacts on demand and price movements. Our identification of safe-haven effects comes from the covariance between ward-level price appreciation (in which wards are appropriately grouped, either by the nationality of the occupants or the average income level of the ward) and the country-level political and economic uncertainty indicators. Second, all standard errors reported in our tables and figures are clustered at the ward-time level (using the robust clustering procedure described in Cameron and Trivedi (2005)) to account for any unexplained commonality driving property prices for all properties in the same ward during the same year.

The next sub-section describes our approach to connect our first identification assumption (that wards with a high pre-existing share of foreign-born people are more likely safe-haven property investment destinations) to data on ward-level immigration patterns in London. This subsection also describes how we attempt to close our analysis by asking whether price increases in wards with higher shares of foreign-born people are a signal of increased future immigration into those wards.

## 2.2 Explaining Ward-level Immigration in London

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.

Our first test involves the regression equation:

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

Here, 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 $f^k_{w,2001}$. This allows us to check whether there is persistence in shares of foreign-born people at the ward level, and acts as a "reality check" on one of our identifying assumptions in our safe-haven price specifications above. The sign and significance of the coefficient $\rho^k$ indicates the degree to which immigrants from country $k$ move into wards with a pre-existing high share of people originating from their home country. It would be consistent with our identification strategy in our hedonic pricing regressions in the previous subsection, if $\rho^k$ is estimated to be statistically significant and positive.

Second, we estimate the following regressions:

$$\Delta f^k_{w,2011} = \alpha + \rho^k f^k_{w,2001} + \pi^k_1 \Delta \ln P_{w,2001} + e_{w,2011}, \quad (5)$$

$$\Delta f^k_{w,2011} = \alpha + \rho^k f^k_{w,2001} + \pi^k_2 \Delta \ln \bar{P}^k_{w,2001} + \pi^k_3 \Delta u_{w,2001} + e_{w,2011}. \quad (6)$$

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 $\pi^k_3$ 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 $\pi_3^k$ towards zero.

Together, specifications (5) and (6) 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.

# 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 HM Land Registry

We obtain data on the complete set of house purchases in London from HM Land Registry. All purchasers of UK houses are required to report transactions to the Land Registry, and the data cover 2,154,590 transactions over the period from 1996 to 2011, of which we utilize a 10% random sample in our empirical analysis. 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 key piece of information for our analysis is that the

Land Registry provides a postcode for each property. In the UK, postcodes provide for a very granular geographical location of a property, often covering just a segment of a street. This allows us to link each property to electoral-ward-level information, described 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 our identification strategy.

## 3.2 Nationwide

In addition to the Land Registry data, we use proprietary loan-level mortgage data covering the period 1996 to 2011, 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 144,149 observations of house purchases widely spread across London electoral wards (we occasionally refer to these simply as "wards") over the sample period.

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.

## 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 socioeconomic 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 subdivisions, 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 on the population density, the shares of different types of houses, the number of cars per household, the age distribution, the education levels and occupations of the adult population, the net income levels and the fraction of properties owned through a mortgage contract.

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.

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 two separate variables - first, we obtain a list of indicators from the International Country Risk Guide (ICRG), which are annual data on the political situation around the world. These indicators are available country-by-country each year, and rank 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.[8] 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

[8] 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.

are interested. They also emphasize that the ICRG indicators capture subjective assessments of risk, which is also relevant, as we are most interested in capturing the sentiment of prospective safe-haven investors' sentiment with respect to 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.

As an alternative measure, we also use country-specific bond yield data as a measure of sovereign risk for each of the countries, taking the spread of the country-level 10-year bond yield over the equivalent 10-year UK government bond yield.[9] These data are retrieved through Reuters-Datastream. Bond yield data are available for different periods for different countries in the sample, the broad date ranges that we use are 1995 to 2011. For some countries long-term bond data are unavailable; we describe the specific data restrictions for these variables in Appendix A.

In our empirical implementation, we subtract the time series mean for each of the above economic and political risk variables, and divide through by the time-series standard deviation across the entire sample period. We then take moving averages over 12 months of each of these variables, making sure to lag them by at least one month relative to the left-hand side house price variables.

[9]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.

# 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 source country 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 understand the economic consequences for both house prices and migratory patterns in a richer and more satisfying fashion than a pure time-series analysis.

## 4.2 Preliminaries and 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 are careful to control in our empirical analysis for the number of transactions occurring in each ward in each month to control for any possible impacts of ward-time variation in

housing market illiquidity.

In our empirical analysis, as described earlier, we present both country-by-country results ($z_{t-1}^k$ and $f_w^k$ 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.

Table 2 shows the mean values of selected variables reported in the column headers, in the top 25% of wards ranked by the share of the population originating from each of the world regions shown in rows. At the bottom of the table, we also provide the statistics for the top 25% of the population ranked by the share of the indigenous UK population, as well as the unconditional cross-ward averages. It appears that wards with high immigration shares are different from ones with predominantly UK-born populations in a number of ways: they are more heavily populated, have far higher shares of flats rather than independent houses, and fewer cars per household. It also appears that wards with higher shares of foreign-born people have a greater fraction of people working in managerial positions, and higher average net incomes.

There does appear to be substantial cross-ward variation in all of these variables, conditional on the share of the population originating overseas, i.e., it is important in our empirical specifications that we control for the direct effect of this variation in ward characteristics on house prices. It is also 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.

Since part of our analysis depends on the cross-sectional spread of variables across wards, it is important to have filtered out any variation in house prices that arises from cross-ward and cross-property variation in these characteristics. In all of the tables and figures in the paper, we report results from both of the datasets for which we have transaction-level data, namely the Land Registry and Nationwide. For the most part, results are strongly consistent across the two datasets, despite the fact that there are two major differences between these datasets. First, 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, and second, Nationwide has a far more comprehensive set of hedonic characteristics than the Land Registry. We view the consistency between the two sets of results as suggestive of the presence of intra-ward spillovers of the effects of safe-haven demand on house prices.

Panel A of Table 3 shows the coefficients $\theta$ from our benchmark specification, which represent the effect of ward characteristics on London house prices. The most relevant ward characteristics for house prices appear to be the types of houses located in the ward, the fraction of people in higher professional occupations, and the average net income level. The bottom part of Panel A shows the estimated *unconditional* effects of the shares of foreign-born people on house prices - here it is evident that wards with high shares of Middle Eastern and Southern European residents have higher prices controlling for all other determinants, and residents with high shares of UK-born and South Asian residents have relatively lower prices.

Panel B of Table 3 shows the estimated time fixed effects, averaged across all months in each year. An alternative approach to our identification strategy would be

to interact these time-fixed effects with $f_w^k$ and $y_w$ or functions of these variables, thus constructing ward-time averages of prices which are cleaned of variation in hedonic characteristics. In this draft, we choose to adopt a restricted version of the ward-time fixed effects which are modelled as functions of the political and economic uncertainty indicators. Under these assumptions, the estimated time-fixed effects will also absorb safe-haven effects which have the *same* impacts on London house price appreciation across all wards. 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.

Panel C shows the estimated coefficients on loan and property-level hedonic characteristics. The estimated contributions to marginal hedonic utility of bedrooms, bathrooms, parking spaces, and floor area (the omitted category is the smallest possible unit in all cases) appear very reasonable. Older properties, all else equal, are valued higher than newer ones (with the exception of properties built in the 2000s which are not new builds), and detached houses are worth more than any other category of houses, once floor-space area is controlled for. Currently, our empirical specifications focus on interactions between ward-level characteristics and foreign political and economic uncertainty, so 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.[10]

[10] 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.

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.

## 4.3 Cross-Country Results, Single Country

In our specification (3), the coefficients $\gamma_0^k$ and $\gamma_1^k$ constitute our identification of safe-haven demand effects on London house prices. Figure 3 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. Panel A of the figure shows the results estimated from Land Registry data, and the top and bottom figures in the panel show the coefficients $\gamma_0^k$ (interaction with the ward-share of foreign-born people) and $\gamma_1^k$ (coefficient with the ward-level average income level) respectively. The figure plots estimated positive coefficients in a dark shade, and estimated negative coefficients in a lighter shade.

The figure clearly 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 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 $k$ (very likely the case).

In Panel A of Figure 3, the negative coefficients are also interesting in the net average income interaction, i.e., $\gamma_1^k$ plotted across countries. For example, this shows that any price appreciation associated with political risk in Pakistan, Portugal, Singapore, and Sri Lanka is concentrated in wards with relatively *low* average income levels, whereas the highest price appreciation in high-income wards is associated with political risk in China, Russia, and the Middle East.

Panel B of Figure 3 shows the results estimated using the Nationwide data, in which the hedonic controls are more extensive, though this is a more restricted sample. The results appear broadly consistent with the patterns detected in the Land Registry data. What is more clearly evident comparing the two pictures of $\gamma_0^k$ and $\gamma_1^k$ here is that there almost appears to be an inverse relationship (controlling for the other determinant) of income and foreign origin. For example, when Greece and countries in South Asia experience high political risk, this is associated with higher future prices in London wards with high shares of people originating from these countries, but controlling for this tendency, to be 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 4, we substitute the country-specific bond yield spread over UK treasuries for the ICRG measure. The figure shows that similar broad patterns hold when we

use the yield spread instead of the risk measure, but the results seem far stronger, especially for the net average income interaction ($\gamma_1^k$), where the coefficients are almost uniformly positive. Of course, given the evidence reported by Longstaff et al. (2011) that sovereign yields are highly correlated across countries, the country-by-country estimation using this measure is likely to suffer from the omitted variable bias referred to earlier. We therefore turn to a more reliable estimation of safe-haven effects in which we include all of the countries in our sample simultaneously, grouped into world regions as described earlier.

## 4.4 Cross-Country Results, World Regions

Figure 5 shows the results of estimating our hedonic regressions including $K$ world regions 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. As before, the top panel reports the estimated coefficients $\gamma_0^k$ and the bottom panel reports $\gamma_1^k$ for each of the six world regions in our estimation. In these specifications, we simply use the ICRG measure of risk as our $z_{t-1}^k$ variables, GDP-weighted within each region.

The figure clearly 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 an underestimate of the total effect that might arise from including homogenous impacts of foreign

uncertainty shocks across London wards. Simply 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-born 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.

In Figure 6, we extend the analysis to allow for the possibility that the impacts of foreign political risk differ by the price of the property. This is to investigate the frequent statements in the popular press that foreign-origin safe-haven demand is most prevalent at the very top of the market.[11] In order to do this, we simply sort both Land Registry and Nationwide samples within each of 46 postcode groups each year on the basis of their transaction price.[12] A transaction is then classified as either belonging to the bottom (1st to 70th percentile), middle (70th to 90th percentile), or top ($>$ 90th percentile) group on the basis of its rank in the postcode group-year distribution.

Figure 6 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,

[11] 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.

[12] We simply group properties based on the first two (or first) letter of their four-character postcodes for the purposes of this sort, to ensure that we have sufficient numbers of properties in each year in each of our price percentile categories.

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 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 in the most expensive properties.

The fact that the price responses to uncertainty in different world regions are not uniform is interesting. One possibility is that some countries' aggregate safe-haven demands are driven purely by a desire to preserve capital, without any associated physical movement of purchasers. Other countries' aggregate safe-haven demands may be primarily driven by prospective immigrants with a desire to reside in London at some future date. We explore the links between immigration, safe-haven demand, and house price increases in London in the next sub-section.

## 4.5 Immigration and House Prices

The final part of our analysis brings additional evidence on the connections between immigration and house price increases. Figure 7 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.

Figure 8 shows estimates of equations (5) and (6). 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 $\pi_3^k$ that we detect for Sri Lanka.

# 5 Conclusions

In this paper, we propose a novel empirical method to identify the impact of safe-haven demand on asset prices. Our empirical results use micro-data from the UK to estimate the effects of foreign economic and political risk on the cross-section of London house

prices. We find economically large, statistically significant, and robust effects along two separate cross-sectional dimensions, and find that the measured impacts of foreign shocks appear to be greatest in the highest value properties. We also attempt to connect safe-haven housing demand to subsequent immigration patterns and provide suggestive evidence along this dimension.

We view our results as providing a more general methodological contribution to the analysis of macro-variation in prices, by using an identification strategy that is grounded in microeconomics. 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. Moreover, our empirical results document that increases in political uncertainty in Southern Europe, China, Russia, and the Middle East are associated with well-identified increases in London house prices in specific wards. This provides 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 iterations of this paper, we hope to better understand the link between house prices and immigration. We also hope to provide a clearer understanding of the specific mechanisms involved in the transmission of price spillovers in housing markets, using London as our laboratory.

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## Appendix A

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. bonds |
|---|---|---|
| Algeria | 1995 - 2011 | |
| China | 1995 - 2011 | 1995 - 2011 |
| Egypt | 1995 - 2011 | |
| Greece | 1995 - 2011 | 1998 - 2011 |
| India | 1995 - 2011 | 1996 - 2011 |
| Italy | 1995 - 2011 | 1995 - 2011 |
| Japan | 1995 - 2011 | 1995 - 2011 |
| Libya | 1995 - 2011 | |
| Malaysia | 1995 - 2011 | 1996 - 2011 |
| Pakistan | 1995 - 2011 | |
| Portugal | 1995 - 2011 | 1995 - 2011 |
| Qatar | 1995 - 2011 | |
| Russia | 1995 - 2011 | 1999 - 2011 |
| Saudi Arabia | 1995 - 2011 | |
| Singapore | 1995 - 2011 | 1998 - 2011 |
| Spain | 1995 - 2011 | 1995 - 2011 |
| Sri Lanka | 1995 - 2011 | |
| Tunisia | 1995 - 2011 | |
| United Arab Emirates | 1995 - 2011 | |

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 markets. We group the countries in six groups 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 loan-level Nationwide data and the transaction-level Land Registry data.

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 | Saudi Arabia | Tunisia |
| | United Arab Emirates | | |
| **6.** | **Southern Europe** | | |
| | Italy | Greece | Portugal |
| | Spain | | |

Panel B
Number of transactions

| | Nationwide | Land Registry |
|---|---|---|
| 1996 | 12,243 | 109,550 |
| 1997 | 13,108 | 132,590 |
| 1998 | 12,463 | 132,800 |
| 1999 | 19,330 | 161,840 |
| 2000 | 8,958 | 146,630 |
| 2001 | 7,522 | 161,200 |
| 2002 | 9,386 | 175,060 |
| 2003 | 7,371 | 155,770 |
| 2004 | 7,565 | 164,870 |
| 2005 | 6,649 | 137,120 |
| 2006 | 10,712 | 172,440 |
| 2007 | 9,143 | 166,850 |
| 2008 | 3,977 | 81,680 |
| 2009 | 4,394 | 75,830 |
| 2010 | 5,068 | 91,200 |
| 2011 | 6,260 | 89,160 |

Table 2
**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 | Southern Europe | 102.98 | 63.43 | 572.31 | 0.68 | 10.31 | 56.30 |
| shares of people born in: | 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 3
**Estimated coefficients in the hedonic regression framework**

This table reports estimated coefficients from the following hedonic regression:

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

where $\mathbb{X}_l$ are property-level characteristics and $\mathbb{C}_w$ ward-level control variables. In this specification, $z_{t-1}^k$ is a monthly 1-year moving average of each 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. The standard errors are clustered at the ward-time level.

Panel A

| | **Land Registry** | **Nationwide** |
|---|---|---|
| Population density | -0.003 | 0.017*** |
| Percent of detached houses | -0.040*** | -0.033*** |
| Percent of flats | 0.203*** | 0.111*** |
| Cars per household | 0.018** | 0.034*** |
| Median age | 0.021*** | -0.009*** |
| Long term unemployed | 0.004 | -0.027*** |
| Higher prof. occupations | 0.113*** | 0.102*** |
| Net average income | 0.132*** | 0.071*** |
| Mortgage ownership | -0.081*** | -0.036*** |
| Volume of transactions | -0.034*** | -0.033*** |
| Foreign-born people shares | | |
| China | -0.011*** | -0.004*** |
| East Asia | -0.026*** | -0.014*** |
| South Asia | -0.032*** | -0.037*** |
| Russia | 0.033*** | -0.016*** |
| Southern Europe | 0.037*** | 0.019*** |
| Middle East | 0.006*** | 0.024*** |
| Domestic population share | -0.049*** | -0.025*** |

Panel B

| | **Land Registry** | **Nationwide** |
|---|---|---|
| Time fixed effects | | |
| 1997 | 0.155*** | 0.209*** |
| 1998 | 0.281*** | 0.348*** |
| 1999 | 0.406*** | 0.505*** |
| 2000 | 0.606*** | 0.689*** |
| 2001 | 0.714*** | 0.800*** |
| 2002 | 0.873*** | 0.975*** |
| 2003 | 0.990*** | 1.073*** |
| 2004 | 1.072*** | 1.147*** |
| 2005 | 1.109*** | 1.172*** |
| 2006 | 1.170*** | 1.249*** |
| 2007 | 1.310*** | 1.384*** |
| 2008 | 1.297*** | 1.328*** |
| 2009 | 1.204*** | 1.252*** |
| 2010 | 1.319*** | 1.349*** |
| 2011 | 1.328*** | 1.383*** |

Table 3
**Estimated coefficients in the hedonic regression framework**
**(continued)**

Panel C

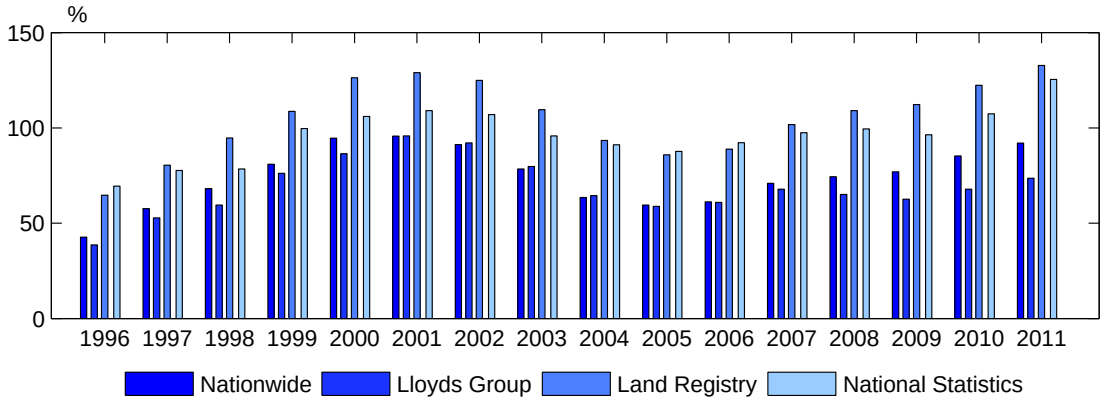
| **Land Registry** | |
|---|---|
| Semi detached house | -0.748*** |
| Terraced house | -0.433*** |
| Flat | -0.581*** |
| New property | 0.274*** |
| Leasehold indicator | -0.330*** |

| **Nationwide** | |
|---|---|
| Semi detached house | -0.146*** |
| Terraced house | -0.197*** |
| Detached bungalow | 0.016 |
| Semi detached bungalow | -0.063*** |
| Purpose built flat | -0.242*** |
| Purpose built maisonette | -0.275*** |
| Flat conversion | -0.186*** |
| Maisonette conversion | -0.217*** |
| New property | 0.139*** |
| Leasehold indicator | -0.116*** |
| First purchase | -0.066*** |
| Two bedrooms | 0.166*** |
| Three bedrooms | 0.220*** |
| Four or five bedrooms | 0.265*** |
| More than five bedrooms | 0.358*** |
| Two bathrooms | 0.064*** |
| Three bathrooms | 0.017*** |
| More bathrooms | 0.001 |
| Parking space | 0.029*** |
| Single garage | 0.066*** |
| Double garage | 0.091*** |
| 1900 to 1920 | -0.004 |
| 1920 to 1940 | -0.049*** |
| 1940 to 1960 | -0.119*** |
| 1960 to 1980 | -0.137*** |
| 1980 to 2000 | -0.016*** |
| after 2000 | 0.011* |
| 50 to 70 $m^2$ | 0.111*** |
| 70 to 90 $m^2$ | 0.187*** |
| 90 to 110 $m^2$ | 0.280*** |
| 110 to 130 $m^2$ | 0.381*** |
| 130 to 150 $m^2$ | 0.472*** |
| 150 to 170 $m^2$ | 0.568*** |
| above 170 $m^2$ | 0.735*** |

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

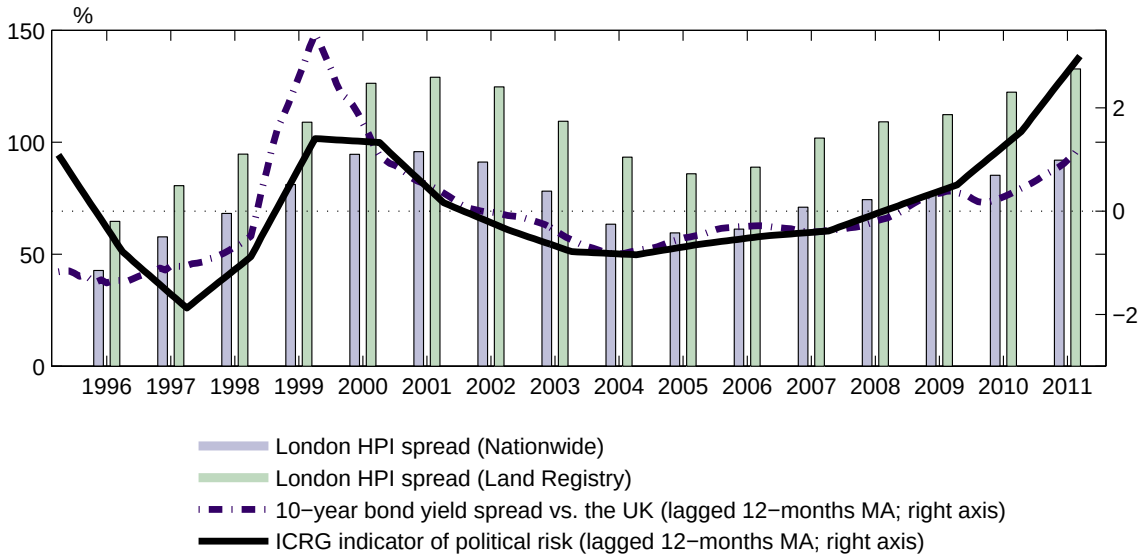


Figure 2
**Shares of foreign-born people**
**- correlation across wards -**

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.

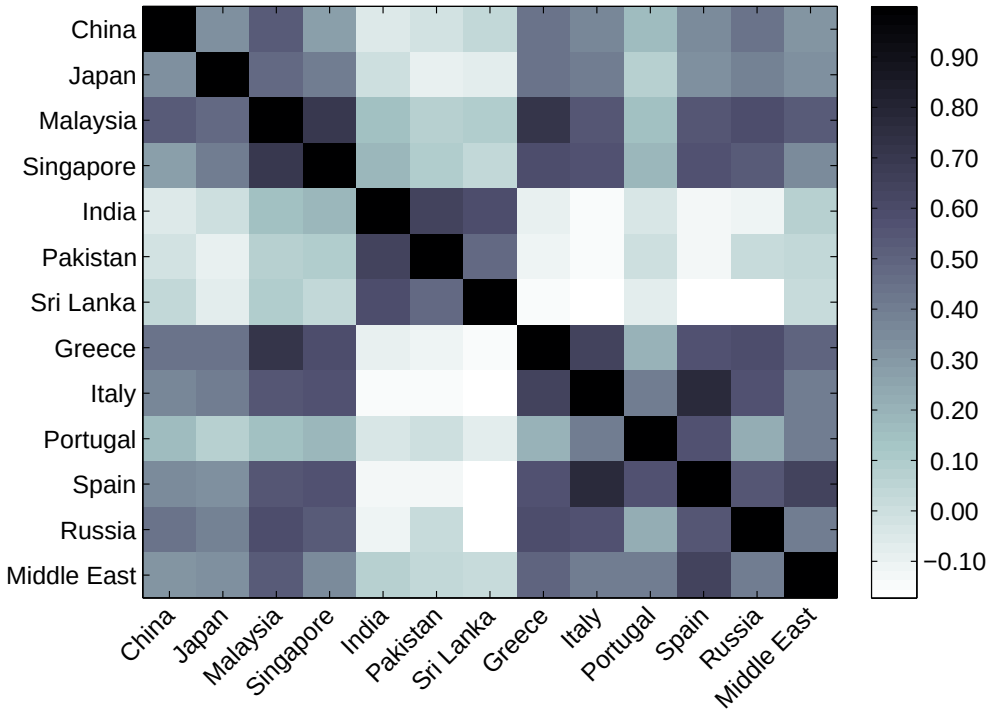


Figure 3
**International political risk and London house prices**

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

$$\ln P_{l,w,t} = \alpha + \beta \mathbb{X}_l + \theta \mathbb{C}_w + \delta_t + \left(\gamma_0^k f_w^k + \gamma_1^k y_w\right) z_{t-1}^k + u_{l,w,t},$$

where $f_w^k$ are the shares of people in ward $w$ born in country $k$ and $y_w$ is average net income. In this specification, $z_{t-1}^k$ is a lagged monthly 1-year moving average of the ICRG risk indicator. All 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
Land Registry

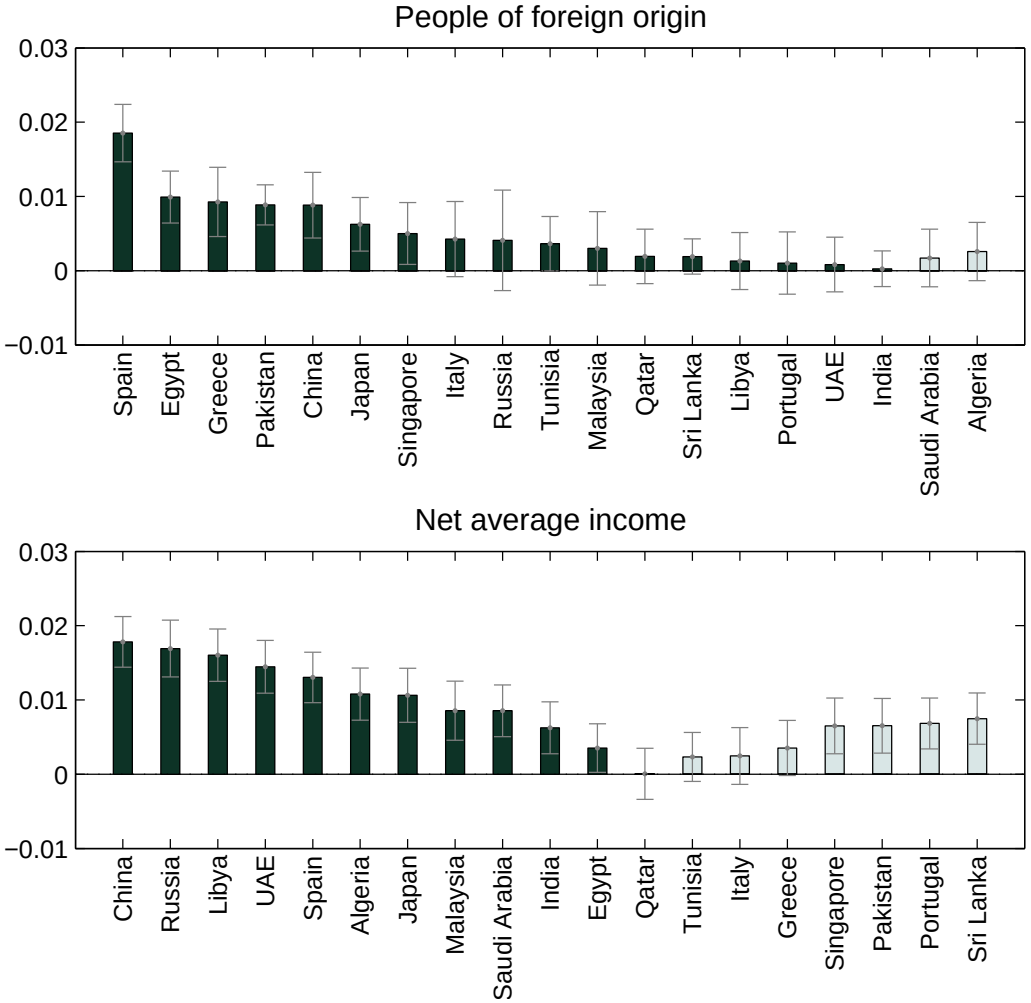


Figure 3
**International political risk and London house prices**
**(continued)**

Panel B
Nationwide

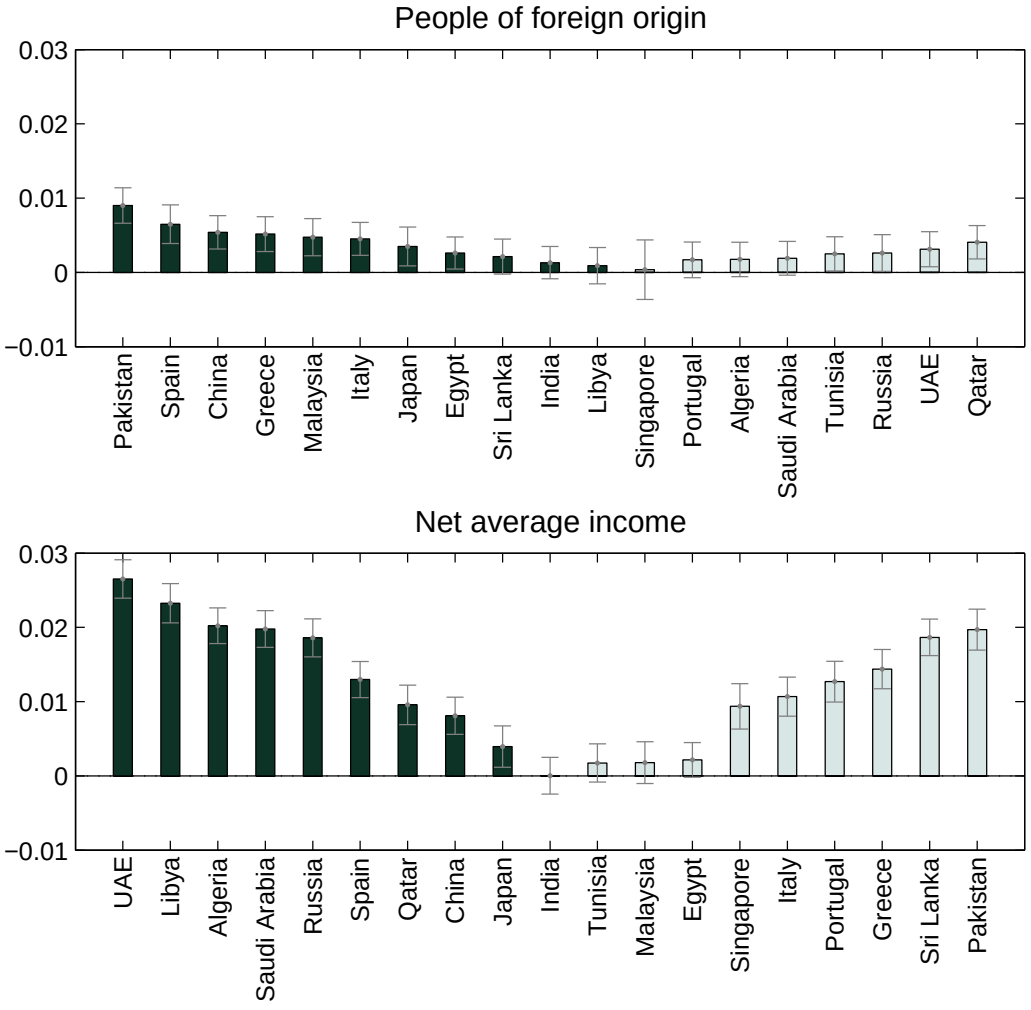


Figure 4
**International bond yield spreads and London house prices**

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

$$\ln P_{l,w,t} = \alpha + \beta \mathbb{X}_l + \theta \mathbb{C}_w + \delta_t + \left(\gamma_0^k f_w^k + \gamma_1^k y_w\right) z_{t-1}^k + u_{l,w,t},$$

where $f_w^k$ are the shares of people in ward $w$ born in country $k$ and $y_w$ is average net income. In this specification, $z_{t-1}^k$ is a lagged monthly 1-year moving average of the 10-year yield spread versus the comparable UK government bond. All 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.

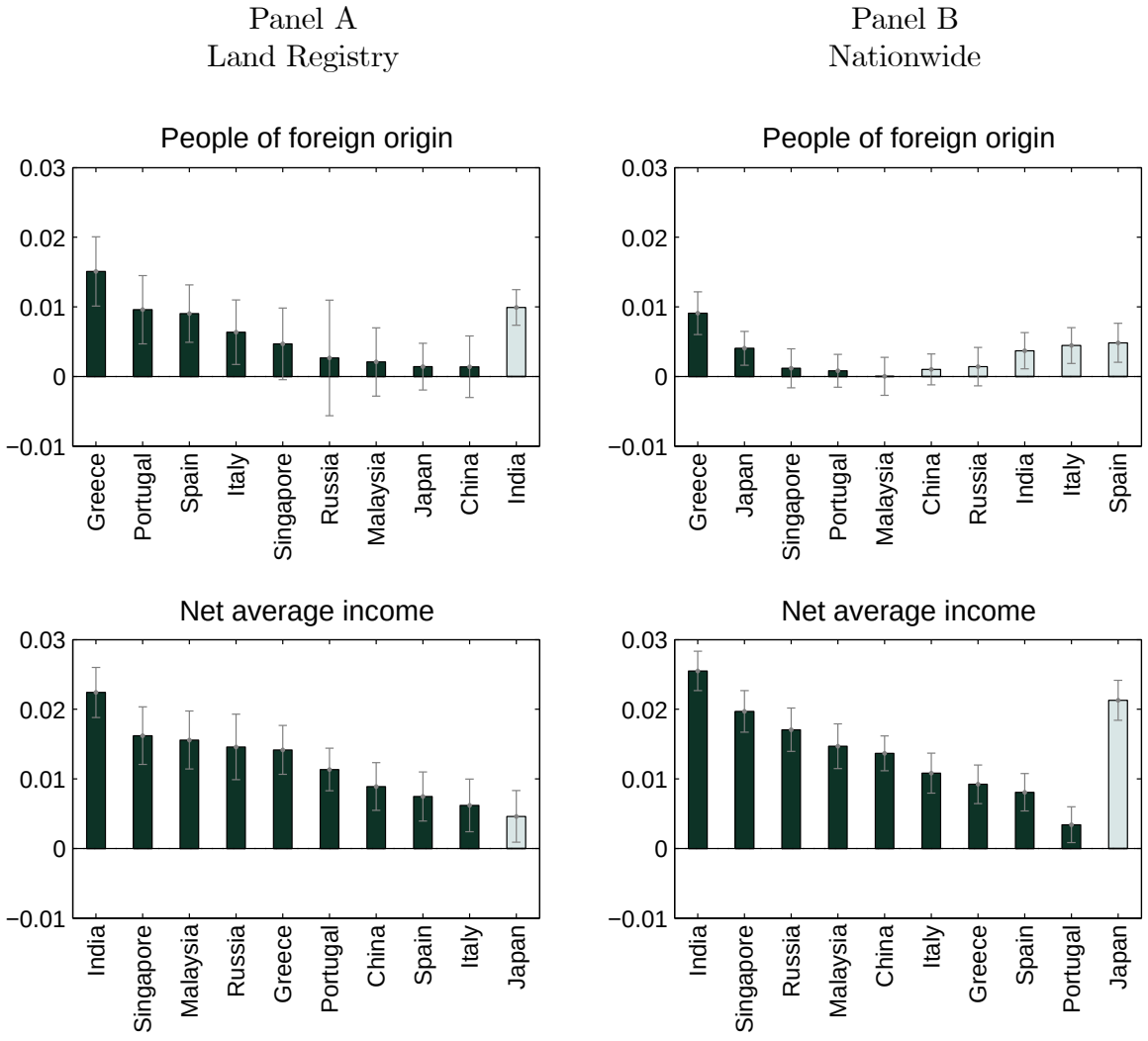


Figure 5
**Analysis at the level of world regions**

This figure reports the coefficients $\gamma_0^k$ and $\gamma_1^k$ from the following hedonic regression:

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

where $f_k^w$ are the shares of people in ward $w$ born in world region $k$ and $y_w$ is average net income. 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 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.

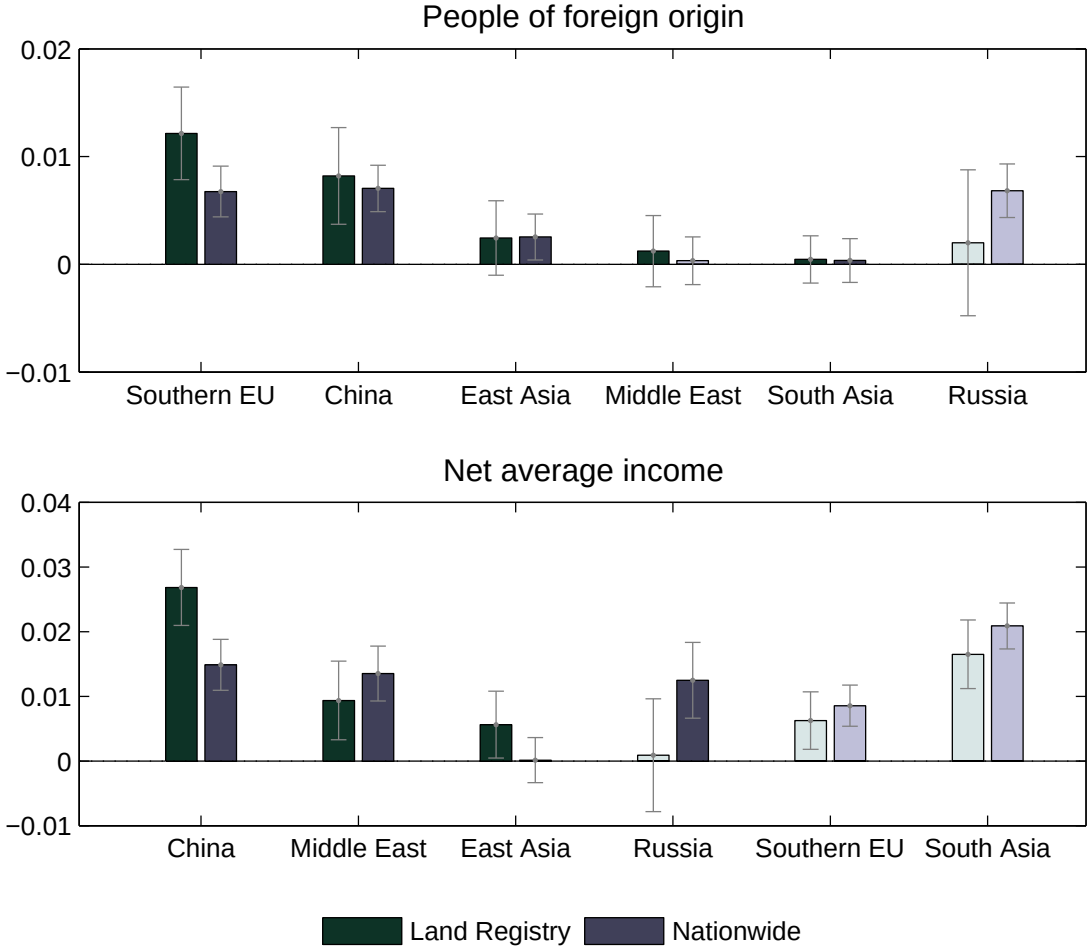


Figure 6
**Safe haven effects across price categories**

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

$$\ln P_{l,w,t} = \alpha + \beta \mathbb{X}_l + \theta \mathbb{C}_w + \delta_t + \sum_{\eta=1}^{3} \sum_{k \in K} \left( \gamma_0^{k,\eta} f_w^k + \gamma_1^{k,\eta} y_w \right) z_t^k + u_{l,w,t},$$

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 $\eta$ indicates the price category of property $l$. 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 postcode group. We generate the latter by using the first two characters of the postcode. The coefficients $\gamma$ 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 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
Land Registry

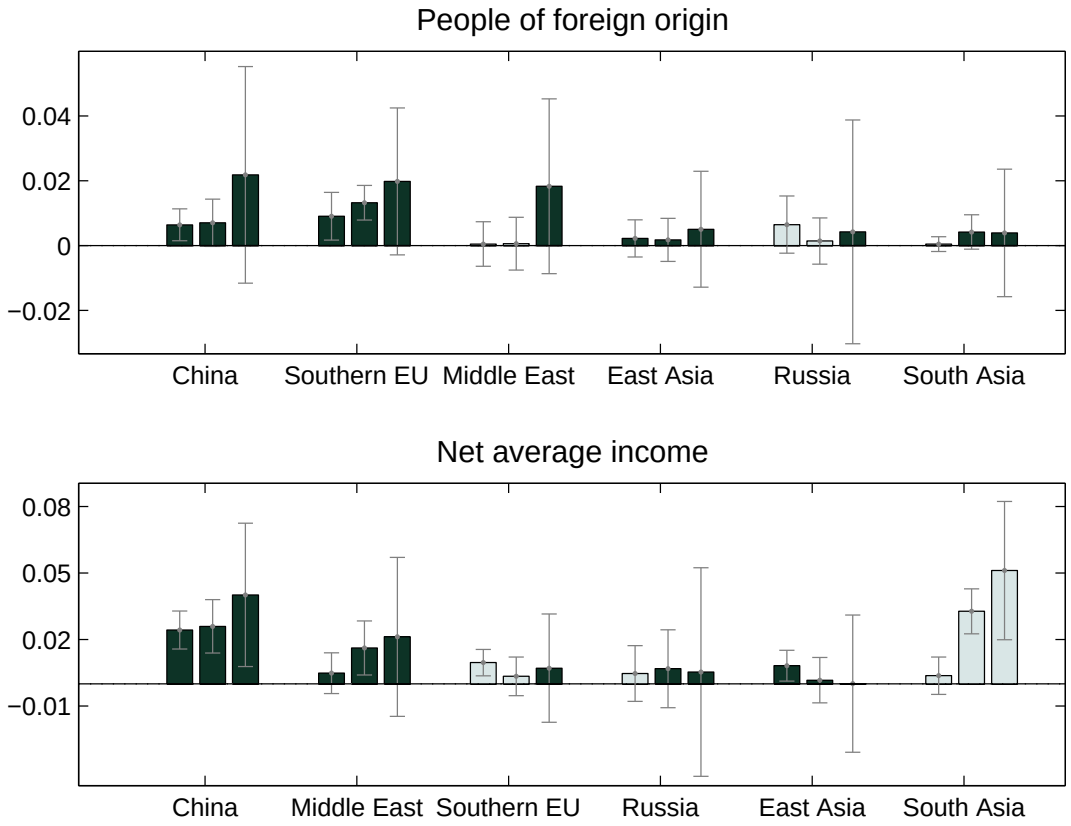


Figure 6
**Safe haven effects across price categories**
**(continued)**

Panel B
Nationwide

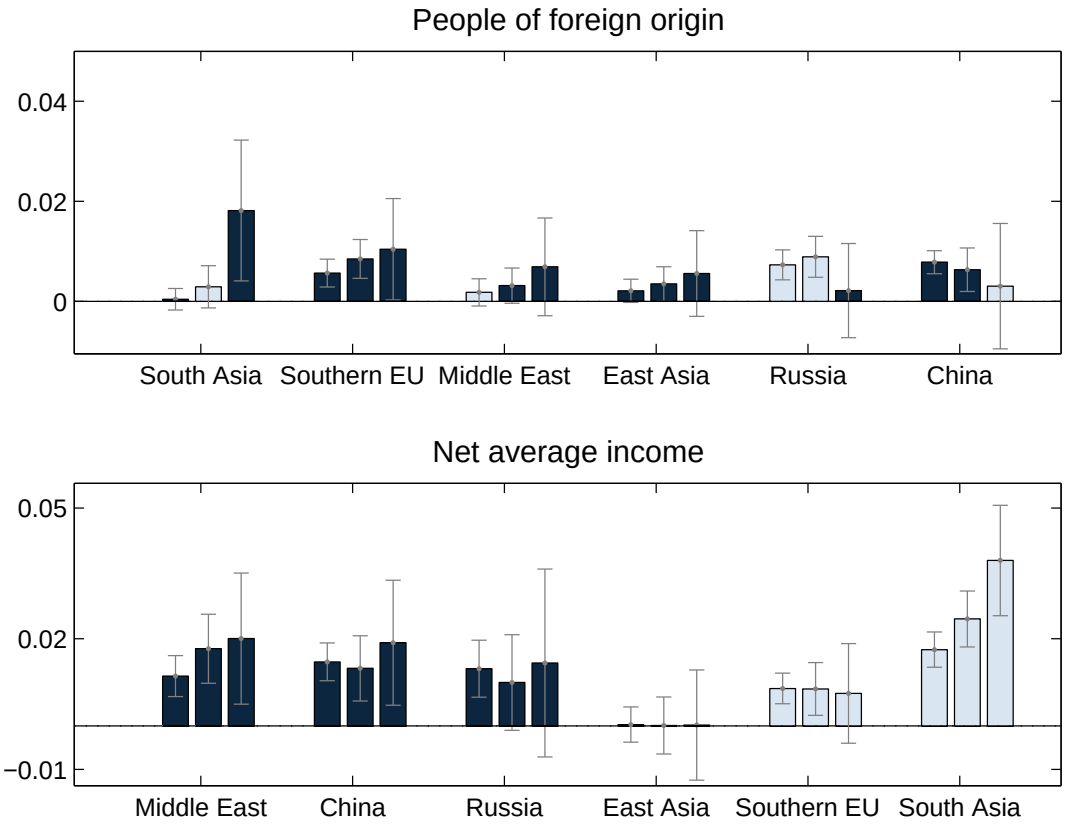


Figure 7
**The evolution of foreign-born people shares through time**

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

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

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.

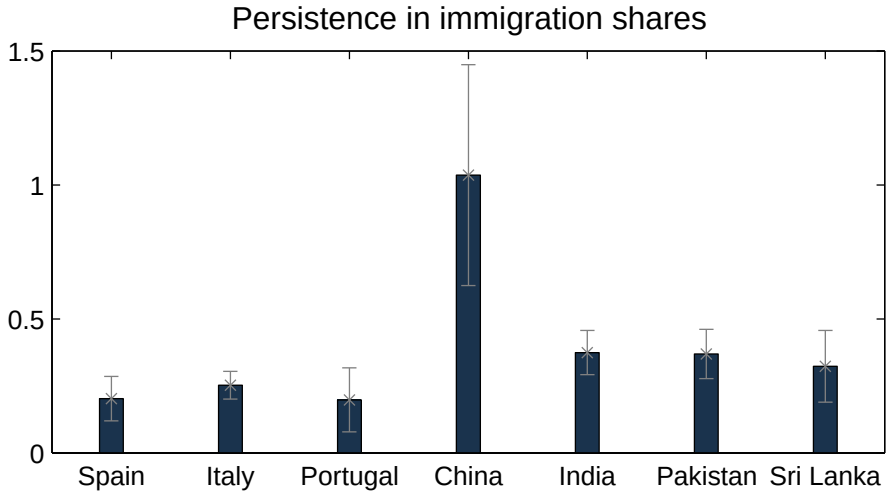


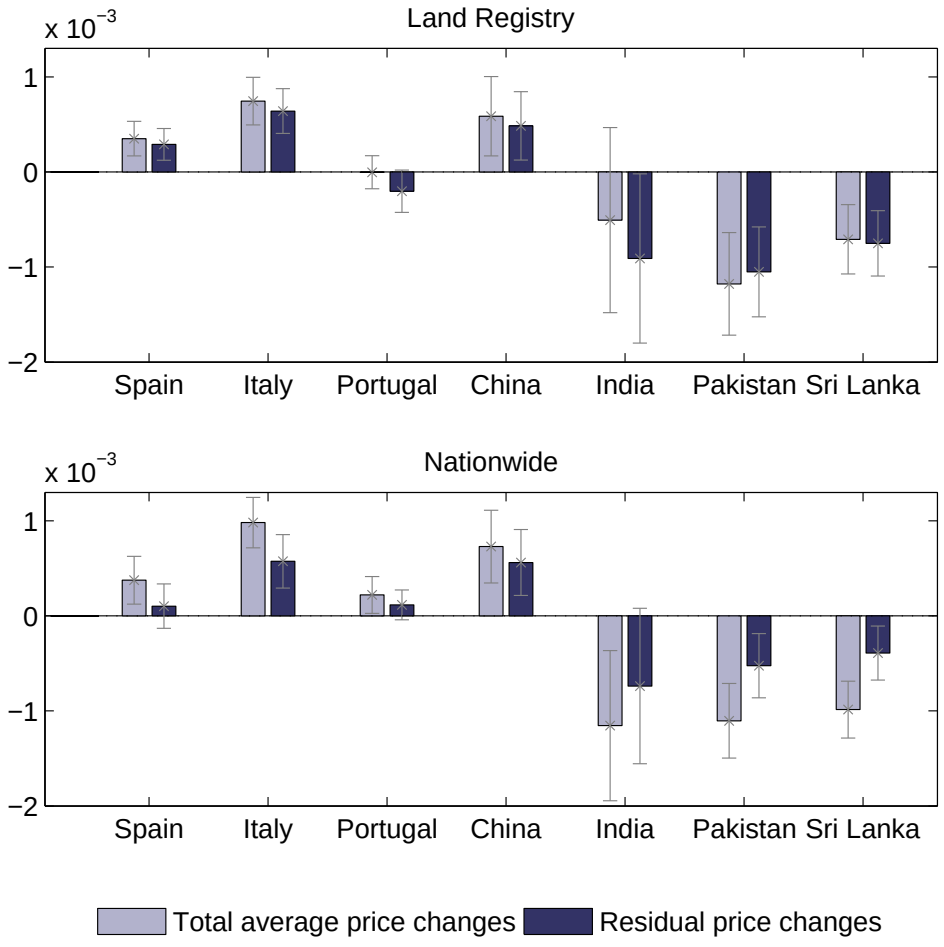
Figure 8

**Relationship between house prices and immigration shares**

This figure reports the coefficients $\pi_1^k$ and $\pi_3^k$ 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.



Internet Appendix for

# Home Away From Home?
# Safe Haven Effects and London House Prices

## Contents

Table IA.1
**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 | Number of people (Count) | Country of Birth, 2001 (UV08) |
| foreign-born people | - born in the UK, China, Japan, Malaysia, Singapore, India, Pakistan, Sri Lanka, Greece, Italy, Portugal or Spain | |
| | Number of people (Count) | Main Language (detailed), 2011 (QS204EW) |
| | - speaking Arabic or Russian | |

Table IA.2
**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 |

Figure IA.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.

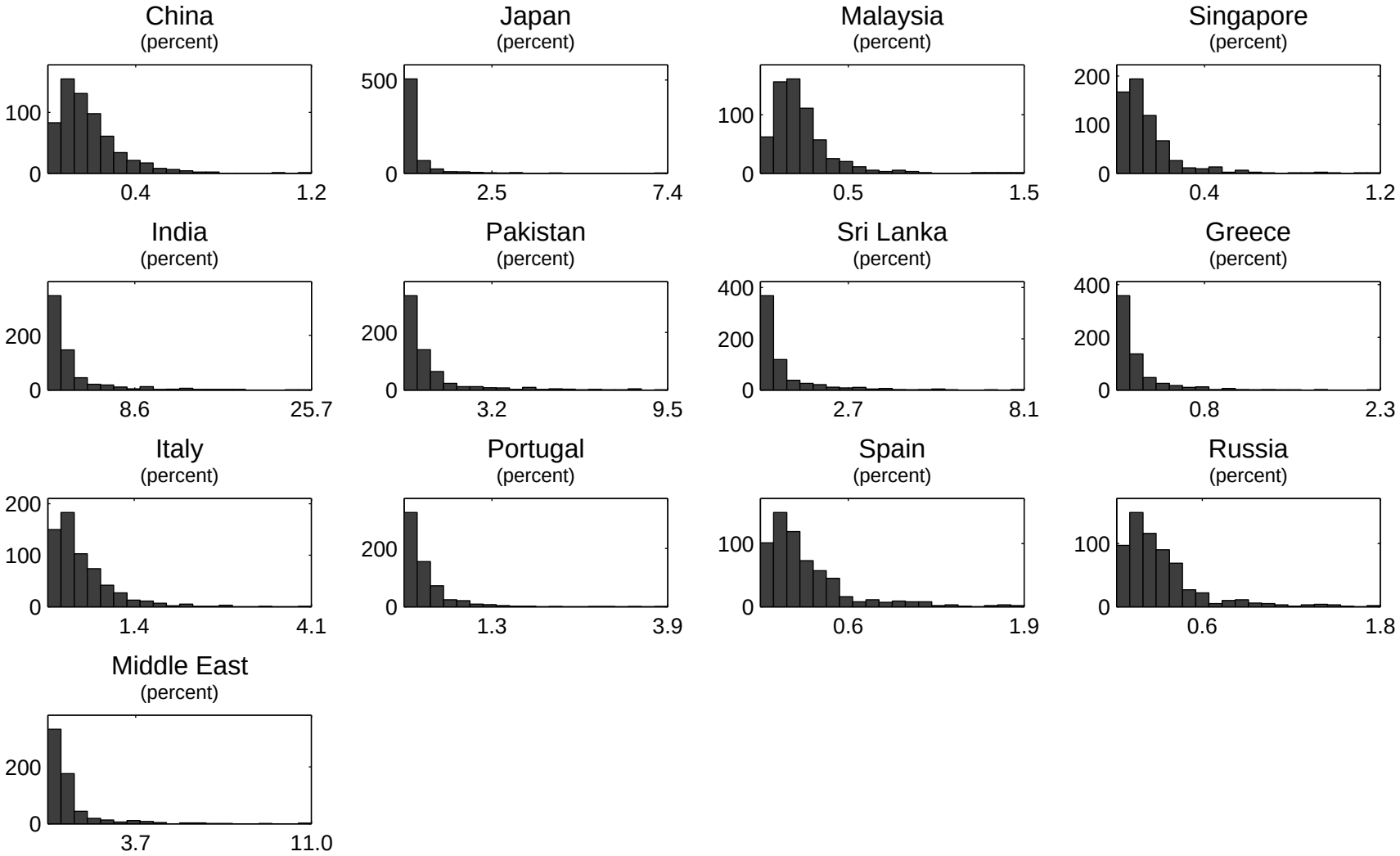


Figure IA.2
**Cross-ward distribution of demographic and socioeconomic variables**

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.

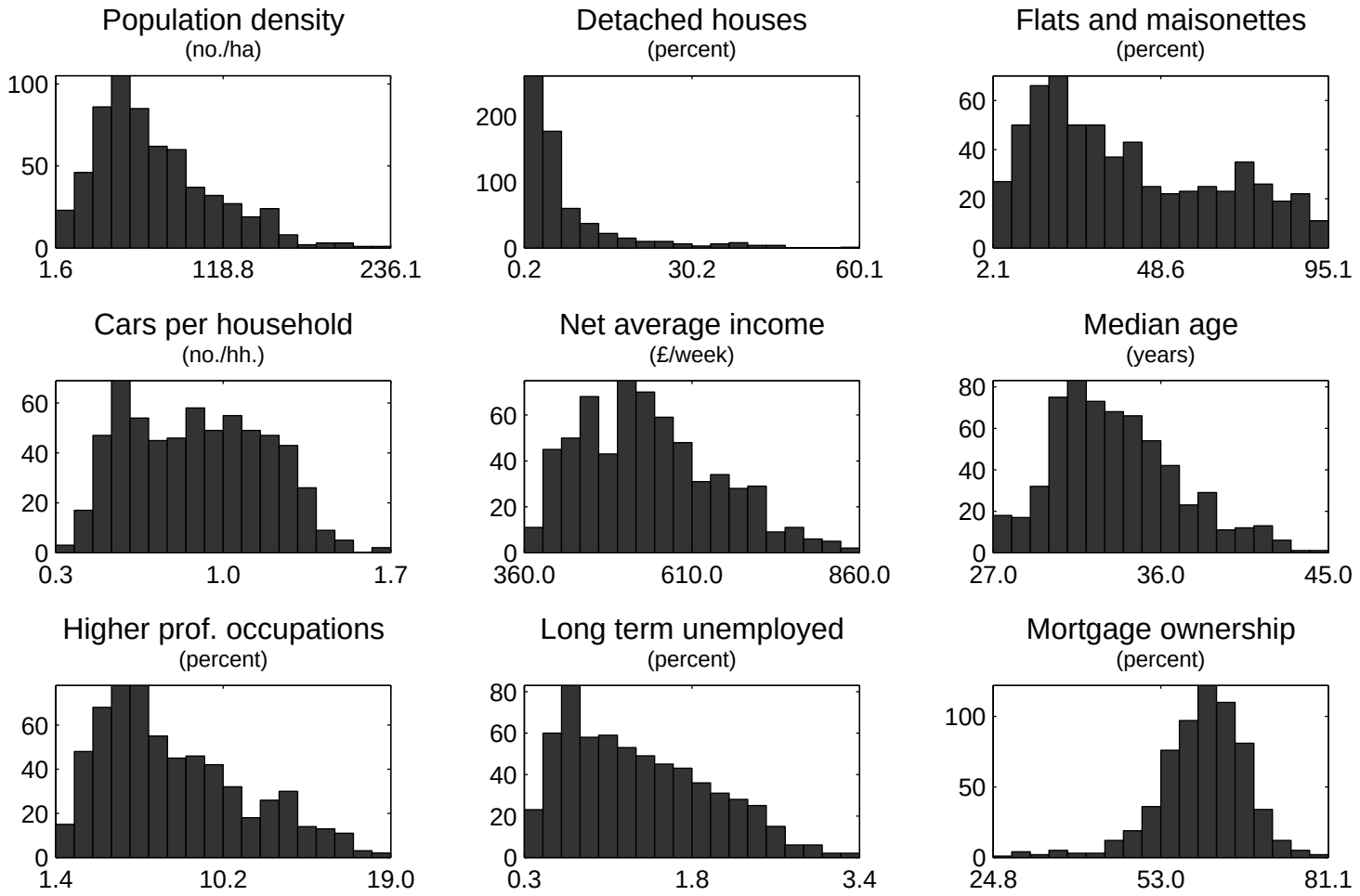


Figure IA.3
**Time series of 10-year government bond yields**

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.

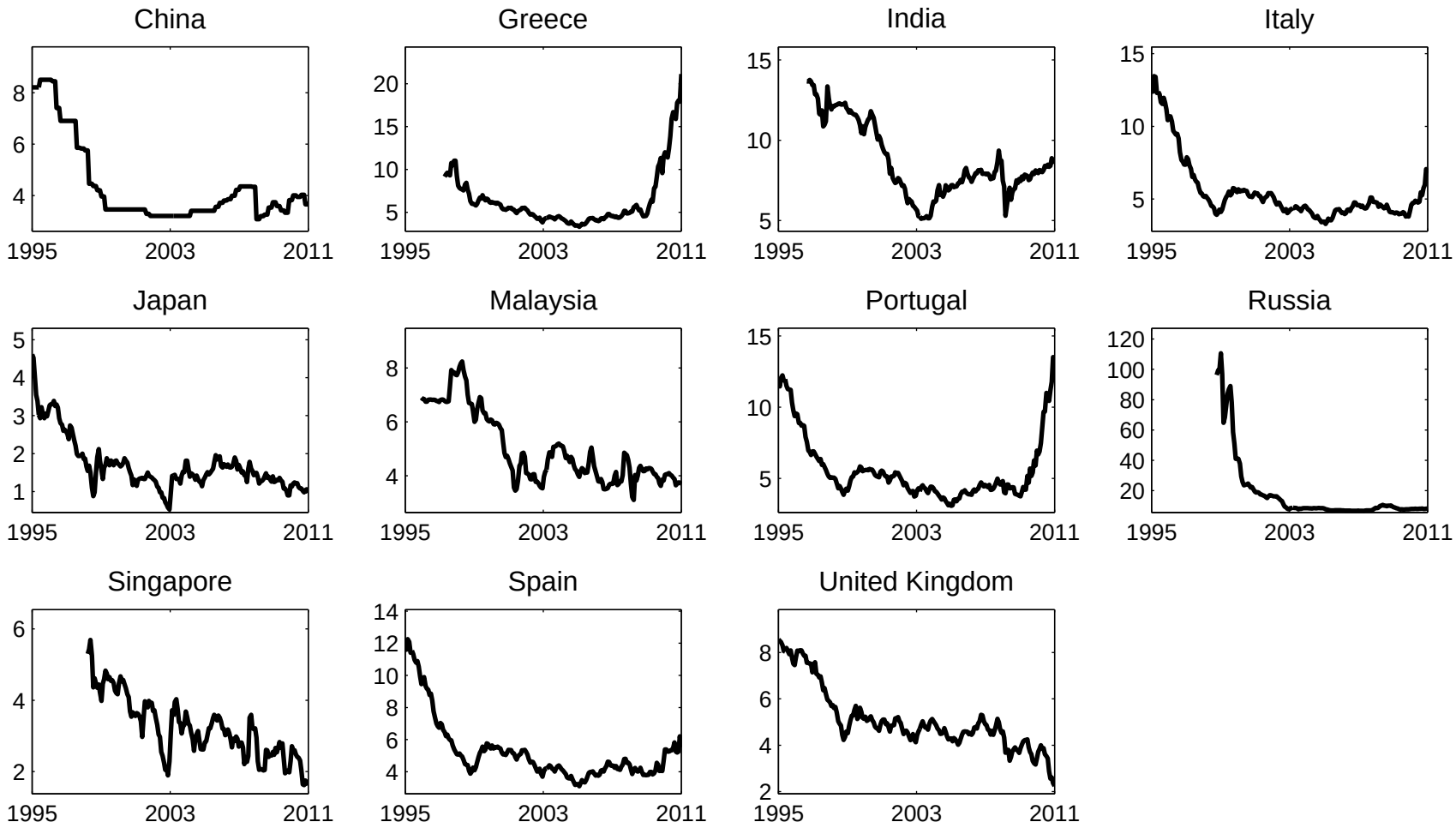


Figure IA.4
**ICRG indexes of political risk**

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.

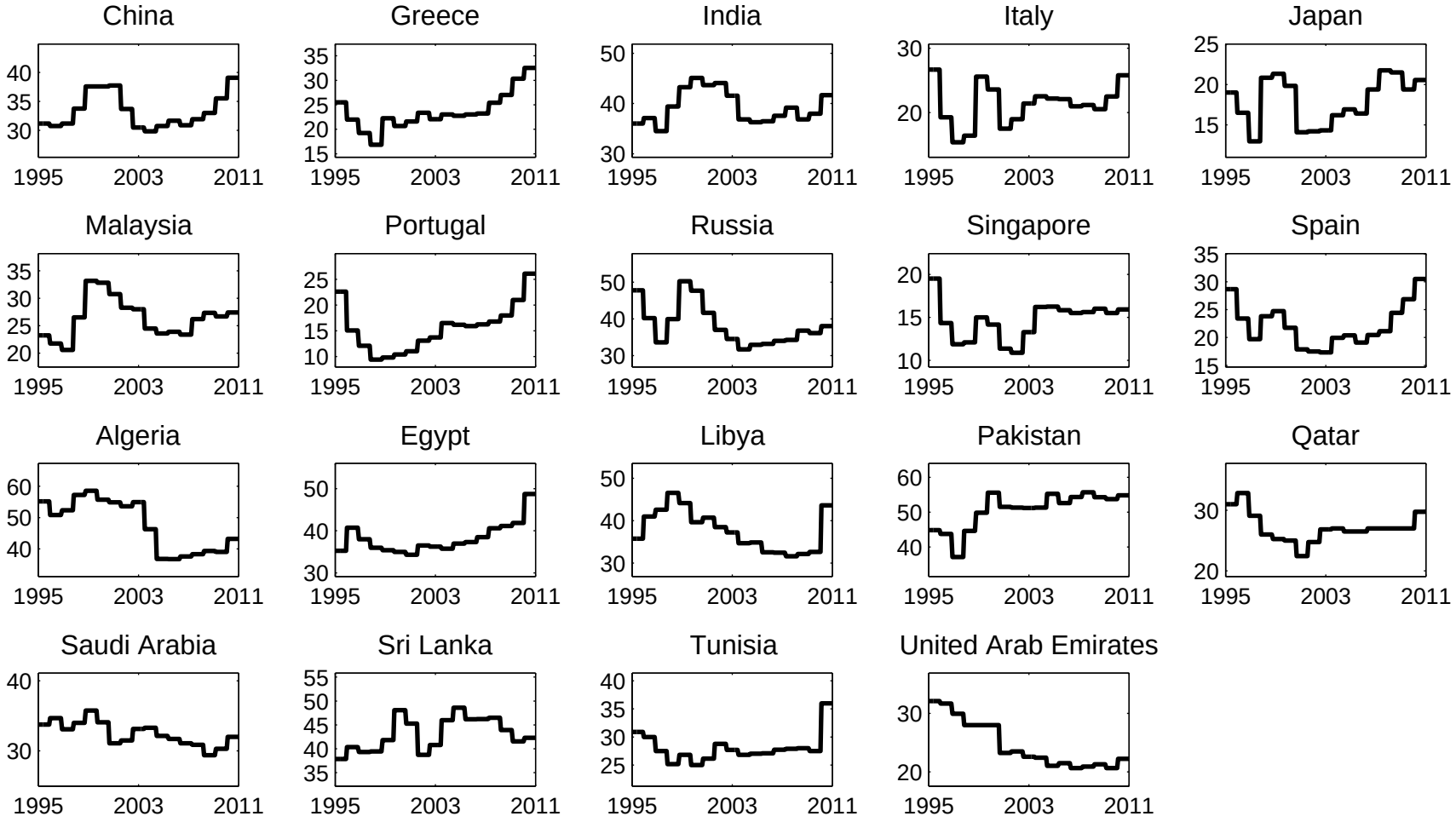


Figure IA.5
**Yield data: China**

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.