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# Capital Flows, Real Estate, and Local Cycles: Evidence from German Cities, Banks, and Firms

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We study how capital flows affects German cities' GDP growth depending on the state of their real estate markets. Identification exploits a policy framework assigning refugees to cities on a quasi-random basis and variation in nondevelopable area for the construction of an exposure measure to real estate market tightness. We estimate that the most exposed cities to real estate market tightness grew at least 1.9 percentage points more than the least exposed ones, cumulatively, from 2009 to 2014. Capital inflows shift credit to firms with more collateral, which leads firms to hire and invest more in response to these shocks. (*JEL* F3, R3, E3)

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It is well known that capital flows are procyclical at business cycle frequency and comove positively with asset prices. As an asset class, real estate is also procyclical and has a large weight in economies' income and wealth (Davis and Nieuwerburgh 2015). Not surprisingly, a large macroeconomic

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literature shows that capital inflows are correlated with outcomes in housing and good markets. However, establishing causality based on country-level data is challenging because of the general equilibrium feedback effects at work.

In this paper, we study how a capital flow shock affects German cities' gross domestic product (GDP) growth depending on the state of their local real estate markets using city-, bank-, and firm-level data. Identification exploits cities' differential exposure to tightness in local real estate markets. By doing so, we connect two large strands of literature: one studies the relationship between capital flows, house prices and output by relying on country variation in the data or the calibration of DSGE models on individual economies (e.g., Aizenman and Jinjarak 2009; Cesa-Bianchi, Ferrero, and Rebucci 2018; Jordà, Schularick, and Taylor 2017; Favilukis, Ludvigson, and Nieuwerburgh 2017); the other exploits variation in the data at the local, firm, or household level to establish causal effects of credit supply and house price shocks on real outcomes (e.g., Adelino, Schoar, and Severino 2015; Chaney, Sraer, and Thesmar 2012; Favara and Imbs 2015; Gan 2007; Mian, Sufi, and Verner 2017, among others).

Our main finding is that the impact on annual GDP growth of a capital flow shock is more significant in cities in which real estate markets are tighter. Moreover, we show that the bulk of this differential impact can be accounted for by *commercial* property price changes triggered by the capital inflow shock that we identify, with no role for *residential* property prices. We estimate that, in cities most exposed to real estate market tightness, real per capita GDP grew at least 1.9 percentage points more, cumulatively, during the 2009–2014 period, than in cities least exposed. We then unpack the transmission mechanism focusing on the commercial real estate sector. The main findings of this part of the empirical analysis are consistent with the working of a collateral channel on the firm side: in response to a capital flow shock, firms with more real estate collateral receive more credit, invest and hire more, without evidence of capital being misallocated across firms, thereby contributing to higher output growth.

We focus on the repatriation of foreign assets by German banks from Southern Europe after the global financial crisis (GFC) during the 2009–2014 period, a particular component of aggregate capital flows during a specific time period. Before the GFC, German and other Northern European banks built up claims along the periphery that were far in excess of their respective countries' bilateral surpluses (Hale and Obstfeld 2016). After the GFC, they reduced cross-border holdings of sovereign debt and increased their holdings of locally issued debt (Brutti and Sauré 2016). Figure 1 shows that Germany's real GDP growth strongly outperformed that of Southern Europe (panel A), as Portugal, Italy, Spain, and Greece were engulfed in a deep and persistent sovereign debt and banking crisis (panel B). Banks rebalanced the composition of their loan portfolios toward domestic households and firms by reducing exposure

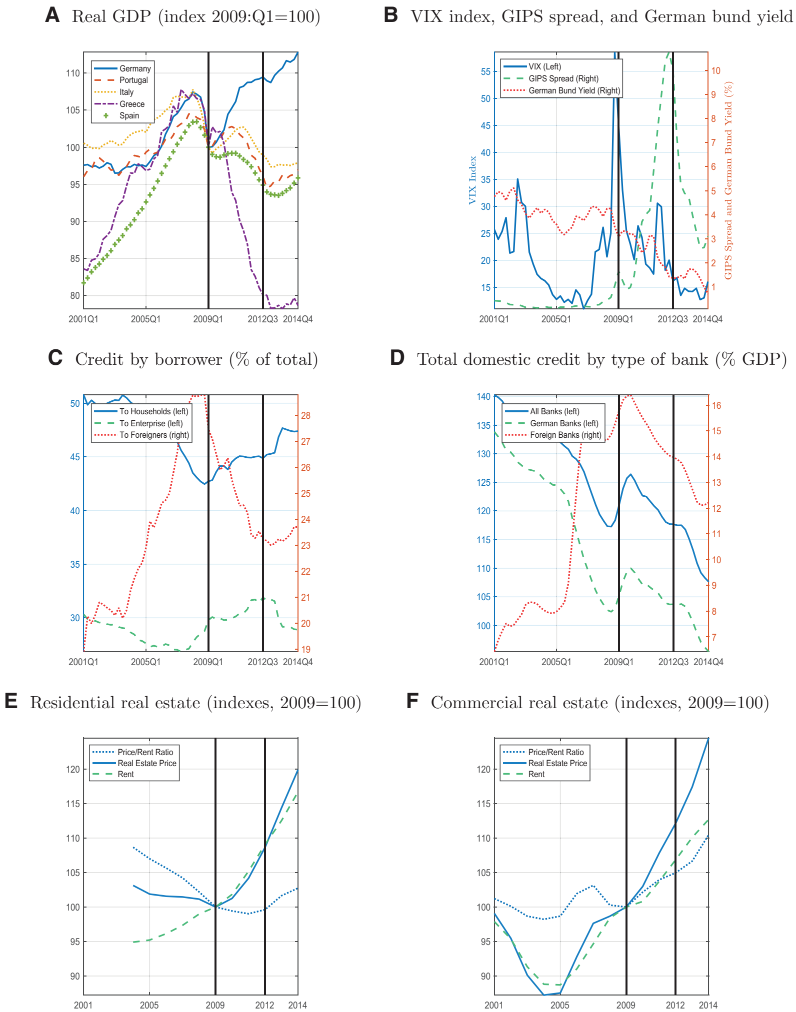


**Figure 1**
**Macroeconomic background**
Panel A plots real GDP for Germany, Portugal, Italy, Greece, and Spain. Panel B plots the U.S. VIX index, together with the GIPS spread and the 10-year German Bund yield. Panel C plots the share of total lending by German banks to different borrowers. Panel D plots total credit as a share of GDP extended by different type of banks. Panels E and F plot, respectively, national residential and commercial real estate prices, rent and price-to-rent indexes. Residential data are not available from 2001–2003. The vertical lines represent the beginning of the German recovery in 2009:Q1 and the “Whatever It Takes” speech by ECB Governor Draghi in 2012:Q3, respectively. See the appendix for variable definitions and data sources.

to foreigners (panel C). Meanwhile, long-term interest rates fell dramatically (panel B), the stock market soared (not reported), and Germany experienced the first property price boom in 20 years, with a cumulative increase exceeding

20% in both the residential and the commercial sector during the 2009–2014 period (panels E and F).

To investigate the transmission of capital flow shocks to city output growth via real estate markets, we assemble a new database that includes city-level and bank-firm-level data described in detail in Section 1. At the aggregate level, we focus on bank flow data, based on the BIS Locational Statistics, which is an important component of total capital flows (Bruno and Shin 2014). Next, we construct a new matched city-level data set that, in addition to publicly available variables, includes a proprietary database on residential and commercial property price indexes from Bulwiengesa AG (a reputable German real estate data provider) and detailed city land cover and use data. Last, to unpack the transmission mechanism through the commercial real estate sector, we construct a second novel bank-firm relationship-level data set based on the German credit register, the Bundesbank supervisory database, and Bureau van Dijk's Amadeus.

To establish causation, we rely on identification by geographic variation. We first establish that the sovereign bond spread of Southern European countries over Germany—the so-called "GIPS spread"—is closely associated with alternative measures of bank inflows from the rest of the euro area during our sample period.[1] This association is particularly strong with German banks' repatriation of gross foreign assets. The link is tight both at the aggregate level and the level of individual bank flows. A higher GIPS spread is also associated with lower domestic lending-deposit spreads at the aggregate level and with lower firm borrowing costs at the firm level, consistent with the notion that, when banks repatriate foreign assets, they expand domestic credit supply. We then interact the GIPS spread, as a proxy for bank inflows from Southern Europe, with an ex ante measure of real estate market tightness that varies across cities quasi-randomly.

Our exposure measure is the product of two variables: the city share of refugees in total refugees and a city measure of nondevelopable area. We evaluate both variables at their presample values in 2008 to control for time variation in the aggregate refugee inflow and any change in land-use regulation. The share of refugees in total refugees (henceforth the "share of refugees") is a novel instrument in the real estate literature. It is a good candidate instrumental variable because it exploits features of the German policy framework that assigns refugee immigrants within non-city-states on a quasi-random basis with respect to the business cycle. In the paper, in fact, we document that the rules and regulations that govern the refugee allocation across cities, together with a language barrier, strictly limit their mobility decisions and the possibility to enter the labor market.

[1] This evidence provides a rationale for why, throughout the paper, we use the "GIPS spread," "bank flows," and "capital flows" interchangeably.

At the same time, refugees need shelter. So they can affect local real estate markets, especially in the commercial sector. In Germany, as in other countries, refugees are entitled to housing benefits and are initially housed in accommodation centers, which compete for the use of built-up structures and developable land for other commercial uses. Unable to integrate in the labor market, in Germany, many refugees remain in these public facilities well past obtaining a work permit and reaching status. As we illustrate in a simple model of local real estate markets reported in the appendix, refugees can push up *commercial* property prices via their claim on the existing supply of commercial real estate services for other productive uses. Indeed, in the paper we find that our exposure measure is a very good predictor of commercial property prices.

Refugees in principle can also bid up residential property prices, as they represent an additional source of demand for residential services. Once they reach status, in fact, they can rent in the open market in Germany. Rent is paid either through earned income or through vouchers, and municipalities pay reimbursements.[2] Our simple model, however, shows that the sign of the impact of a higher city share of refugees on residential property prices is ambiguous. This is because a higher city share of refugees lowers the supply of commercial real estate services for production uses. Hence, total output and consumption of the residents are also lower, weakening a much larger component of the local demand for residential real estate services. Higher taxes needed to pay for refugee benefits further reduce the housing demand of the residents. The net effect of a higher city share of refugees thus depends on parameter values in our model, and it is negative for a model parametrization consistent with the institutional details of the German policy framework and some of the stylized facts of our data. To be able to assess the relative importance of the residential and commercial sectors in the transmission of the capital flow shocks, in the paper, we also propose an alternative instrument for residential property prices, based on demographics, which we will discuss in more detail below.

To control for city-level real estate supply elasticity, we interact the share of refugees with a measure of land scarcity due to geography and land-use regulation in the spirit of Saiz (2010). The indicator that we employ is the 2008 value of the fraction of a city's reference area covered by water, agriculture, forest and other open areas, adding up to open space. Unlike in the United States, in Germany, city variation in the incidence of steep-slope terrains and water bodies has a more limited impact on the distribution of nondevelopable area across cities (OECD 2017b). Moreover, in Germany, land-use regulations are distributed more uniformly than in the United States (Schmidt and Buehler 2007). Indeed, in the paper, we argue that both components of our exposure measure are plausibly distributed quasi-randomly across cities, but neither of

[2] Municipalities house refugees in independent accommodations even if they are unable to self-sustain depending on personal and family circumstances.

them predicts property prices as well as the interaction of the two in both segments of the local real estate markets.

Equipped with this city measure of exposure to real estate market tightness, we estimate the differential impact of a capital flow shock on German cities' output in both reduced-form and instrumental variable setups. In the reduced-form specification, we find that changes in the GIPS spread have a stronger impact on output growth in cities with tighter real estate markets. Moreover, when we horse race the residential and the commercial sectors against each other by introducing both property price indexes in the instrumental variable specification, we find strong evidence that the commercial sector dominates. In this setting, changes in the GIPS spread have no significant effect on output growth via residential property prices.[3] As a result, the bulk of the city output growth differential estimated in the reduced-form specification can be accounted for by the different response of commercial property prices across cities triggered by the GIPS spread increase. Specifically, we estimate that, during the 2009–2014 period, for every 100-basis point (bps) increase in the GIPS spread, cities at the 75th percentile of the exposure distribution grew 8 bps more per year than cities at the 25th percentile. Given an average increase in the GIPS spread over the sample period of 390 bps, these estimates imply that the cities most exposed to commercial real estate market tightness grew 31 bps more per year than the least exposed cities, or 1.9 percentage points more cumulatively during the 2009–2014 period.

In light of the finding that the residential sector does not appear to play an important mediating role during the episode of bank retrenchment that we study, in the second part of the paper, we focus on the role of the commercial sector. When we unpack the transmission mechanism through the commercial sector, we find that firms with more real estate collateral, as measured by tangible fixed assets, receive more credit when banks repatriate foreign assets and retrench from Southern Europe. Firms with more collateral also invest and hire more, thereby contributing to higher output growth. During this retrenchment episode, however, we find no evidence that better credit access and higher investment by firms with more real estate collateral leads to capital misallocation.[4] Overall, the bank-firm-level evidence on the transmission mechanisms that we report in the second part of the paper is consistent with the working of a collateral channel on the firm side (e.g., Liu, Wang, and Zha 2013; Chaney, Sraer, and Thesmar 2012; Gan 2007; Schmalz, Sraer, and Thesmar 2017; Adelino, Schoar, and Severino 2015, among others).

[3] We are agnostic on the interpretation of this surprising finding. We note, however, that, unlike the case of the United States and other countries in which household debt, together with house prices, rose sharply, the German housing boom is not associated with a credit boom, at least through the end of our sample period (see, e.g., panel D of Figure 1).

[4] Again, while we are agnostic about this finding, we note it can be related to the fact that the post-GFC German real estate boom is not associated with a credit boom, with retrenchment possibly reflecting a "flight-to-safety."

Our paper relates to the literature along multiple dimensions. First, our paper connects to the literature on the relationship between capital flows, the business cycle and house prices based on cross-country evidence or DSGE models of individual economies (for a survey of both theory and evidence, see Favilukis et al. 2013). Our main contribution here is to identify the causal effect of a capital flow shock on short-term output growth *via* real estate markets by exploiting city variation in the data. As far as we are aware, this is the first paper that documents empirically the mediating role of property prices in the transmission of capital flow shocks in a causal manner. For example, Aizenman and Jinjarak (2009) document a strong positive association between the current account (i.e., net capital flows) and house prices, holding constant certain characteristics in a large panel of countries. We document a similarly close association between bank flows and property prices, but we establish causation. Cesa-Bianchi, Cespedes, and Rebucci (2015) and Cesa-Bianchi, Ferrero, and Rebucci (2018) show that residential house prices comove strongly with consumption growth conditional on a bank flow shock identified in the time-series dimension of the country panel and relate consumption sensitivity to the shock with different country characteristics. We exploit the quasi-random variation of our real estate market exposure to assess causally the differential impact of a bank flow shock across cities in one advanced open economy. Moreover, unlike most of the empirical literature on capital flows and the business cycle, with Forbes and Warnock (2012) being one of a few exceptions, we investigate an episode of capital "retrenchment." Favilukis, Ludvigson, and Nieuwerburgh (2017) study theoretically the impact of capital flows into the United States and show that lower bond yields associated with inflows of foreign capital cannot explain the U.S. residential house price boom. We distinguish between the commercial and the residential real estate sectors and provide disaggregated evidence that firms' real estate collateral introduces additional channels of transmission of capital flow shocks. Moreover, we find evidence consistent with Favilukis, Ludvigson, and Nieuwerburgh (2017) that residential property prices are not part of the transmission mechanism.

Second, the paper relates to the literature on the link between capital flows, credit, the real economy and house prices that exploits regional variation in the data. Employing bank-firm-level data from the Turkish credit registry, di Giovanni et al. (2018) show that capital inflows increase the volume and reduce the price of domestic credit. We provide similar evidence using credit register data for a major advanced economy and also evaluate the transmission mechanism to house prices and firm outcomes, including misallocation. Mian, Sufi, and Verner (2017) show that an aggregate credit supply shock boosts local demand and amplifies the expansion phase of the business cycle in the United States, with higher GDP, employment, residential investment, and house prices. We document comparable dynamics for Germany, but explore the transmission mechanism at the bank-firm level. Giroud and Mueller (2018) show that leverage buildups by large U.S. publicly listed firms lead to boom-bust cycles in employment, with a short-run expansion and a medium-term contraction. We find consistent evidence on the city response of German employment to the capital flow shock. Cetorelli and Goldberg (2012) show that global banks contracted their direct and indirect cross-border lending during the GFC, leading to a reduction in credit supply in regions from which capital was pulled. We study the complementary case of a country whose banks repatriated foreign assets during and after the GFC and establish that bank retrenchment led to an increase in domestic credit supply, benefiting especially firms that are richer in real estate collateral.

Our paper also relates to the large literature on the collateral channel and real estate prices. The underlying mechanism is that agents use pledgeable assets as collateral, typically land and buildings, to finance productive projects, residential housing, and durable consumption. In this setup, fluctuations in property prices can have sizable effects on macroeconomic aggregates. Iacoviello (2005) and Liu, Wang, and Zha (2013) develop closed-economy DSGE models of the collateral channel on the household and the firm side, respectively, estimated with U.S. data. Liu, Wang, and Zha (2013), in particular, introduce land in the firm borrowing constraint and show that the model can explain the comovement between land prices and business investments; a correlation that the collateral channel from the household side cannot match. We show that commercial property price changes triggered by bank flow shocks can account for all the differential impact of these shocks on city output growth, thus providing more granular evidence consistent with the working of a collateral channel on the firm side. Chaney, Sraer, and Thesmar (2012) use U.S. firm-level data showing that an exogenous variation in property prices triggered by aggregate mortgage rate changes has a sizable impact on corporate investment. Using comparable data and methodology, we find that these effects are quantitatively sizable in the transmission of bank flow shocks. Moreover, we horse race residential and commercial property prices and show that, unlike in the U.S. case, the commercial sector dominates in the transmission. Other studies with micro data show that fluctuations in property prices can affect firm employment, exit and entry decisions, and capital structure (e.g., Schmalz, Sraer, and Thesmar (2017); Cvijanović (2014)). We provide micro evidence on the transmission mechanism of bank flow shocks through similar effects on firm hiring and investment decisions and total factor productivity.

The paper also speaks to the new literature on the role of foreign purchases in global cities like London, New York and Vancouver. Favilukis and Nieuwerburgh (2017) develop a heterogeneous spatial model of cities and show that an increase in out-of-town home buyers can drive up local real estate prices. Consistent with their findings, we show that higher refugee shares predict property price changes in the commercial sector. However, refugees can lead to a fall in residential property prices if commercial real estate is a productive asset. Badarinza and Ramadorai (2018) use a "preferred habitat" framework to document that foreign risk can affect real estate valuations in global cities. We show that instability in Southern Europe was associated with bank retrenchment and affected city-level real estate valuations in Germany.

Finally, other papers have used the government allocation of refugees for identification purposes. Dustmann, Vasiljeva, and Damm (2019) and Eckert, Walsh, and Hejlesen (2018) exploit the quasi-random nature of the refugee allocation in Denmark to study the impact of immigration on voting outcomes and the urban wage premium, respectively. As far as we are aware, this is the first paper that uses the spatial distribution of refugees as an instrument for property prices.

## 1. Data

To conduct the empirical analysis, we assembled a new and unique data set at the annual and quarterly frequency, from 2009:Q1 to 2014:Q4.[5] As a source of aggregate capital flow shocks, we focus on *cross-border bank flows* from the BIS Locational Statistics, or “bank flows” for brevity, which is an important share of total flows (Bruno and Shin 2014). In particular, we will focus on the component of bank flows predicted by the GIPS spread. In addition to city-level output from official statistics, the data set includes an annual proprietary panel data set on residential and commercial property price indexes at the city level from Bulwiengesa AG. To study the details of the transmission mechanism, we then merge information on bank and firm characteristics from Bundesbank supervisory data and Bureau van Dijk’s Amadeus with individual bank-firm relationship data from the German credit register.

### 1.1 City-level data

Data on residential and commercial nominal property price indexes at the city level are proprietary from the research consultancy Bulwiengesa AG, accessed through the Bundesbank.[6] Both residential and commercial indexes are at the annual frequency. To construct nominal property price indexes by city and type of property, Bulwiengesa AG uses both valuation and transaction data from building and loan associations, research institutions, realtor associations, as well as the chambers of industry and commerce. Residential indexes include the price of town houses, owner-occupied apartments, and single-family detached homes. Commercial indexes include information on two segments of the market, retail and office buildings. The indexes are calculated at the city level as simple averages of the individual unit prices. As city-level CPI indexes are not available, we deflate nominal property price indexes by using *state-level* official consumer price indexes.

[5] Table A1 in the appendix defines all city and bank-firm-level variables that we use and describes their sources. Table A2 in the appendix reports summary statistics for all variables used in the empirical analysis.

[6] The Bundesbank relies on this provider for the publication of national indexes, also shared with the European Central Bank.

The dependent variable in our main city-level regressions is real per capita GDP growth. We deflate nominal GDP by using the same official state-level consumer price indexes used to deflate property price indexes. To construct our instrumental variables described in more detail in Section 2.2 and to conduct various robustness exercises, we match data on refugees, population levels, and population density from official German statistics with land use data from the German Monitor of Settlement and Open Space Development (IOER Monitor), which is a detailed database combining information from satellite imaging with geo expert data and other statistical sources, capturing both man-made and geographical limits on real estate supply.[7]

To match GDP, real estate price indexes, and other city-level variables, we use the common city identifiers in the official German statistics. The Federal Republic of Germany has 16 states, 3 of which are city-states, namely, Berlin, Hamburg, and Bremen. The 13 non-city-states are divided in administrative regions, which are the smallest geographical units for which the German national accounts report output data. There are 401 administrative regions, either urban areas (Kreisfreie Stadt, 107) or rural areas (Landkreis, 294), comprising one or more municipalities. In our empirical analysis, we focus on 85 urban areas listed in Table A3 in the appendix, which we will call cities and for which we can match all variables above, excluding the three city-states. We also drop the very small city of Coburg, as the summary statistics and a plot of real GDP growth showed clear evidence of miscalculation of the GDP index. The remaining 18 urban areas cannot be matched because Bulwiengesa does not provide commercial property price indexes.

Bulwiengesa provides residential price indexes for all 401 areas. However, it only supplies commercial property price indexes for 127 urban areas or municipalities for which it historically deemed there was enough/sufficient commercial real estate activity. These includes the 85 urban areas in our sample, the three city-states, and Coburg. The remaining 38 areas are municipalities in rural areas with commercial activity, but they are not urban areas as defined in the national accounts. Our reduced-form results, however, are robust to extending the sample to all 127 cities for which there are commercial price indexes (matching the additional 38 Bulwiengesa municipalities with rural areas in the national accounts) or the 107 urban areas in the national accounts. In contrast, when we run our reduced-form regression for the 401 cities and rural areas for which we have residential property prices, the capital flow shock has no differential impact, consistent with the evidence in the next section that the residential sector is not part of the transmission mechanism of the capital flow shock in our core sample of urban areas.

[7] In international treaties, "asylum seekers" are individuals applying for asylum, while "refugees" refers to individuals whose asylum status has been approved. In the German statistics, the total number of refugees includes (a) admitted refugees on a permanent basis, (b) admitted refugees on a temporary basis, (c) rejected asylum seekers that cannot be relocated, and (d) a small fraction of asylum seekers not processed within the year of reference.

## 1.2 Bank-firm-level data

To explore the relationship between capital flows, bank lending, firm decisions, and commercial property prices, we match data from the German credit register over the period 2009:Q1–2014:Q4 with Bundesbank bank balance sheet data and firm-level data from Bureau van Dijk's Amadeus.

The German credit register contains information on bank exposure, including loans, bonds, off-balance sheet, and derivative positions (excluding trading book positions). Financial institutions in Germany are required to report to the credit register if their exposure to an individual borrower, or the sum of the exposures to borrowers belonging to one legal entity, exceeds a threshold of 1 million euro.[8] A borrowing entity in the credit register, however, can have multiple bank relationships, a feature of the data that we exploit in our analysis. Summing all loans reported in the credit register in a given quarter amounts to approximately two-thirds of the total credit outstanding reported in the Bundesbank bank balance sheet statistics.

We match credit register data with end-quarter values on bank balance sheets from Bundesbank supervisory data. Balance sheet variables include total assets, liquid assets, the interbank-to-deposit funding ratio, the regulatory-capital ratio, nonperforming loans, the return on assets and net and gross bank foreign assets. We also match firm-level accounting variables from the Bureau van Dijk's Amadeus with the credit register data. In our analysis, we use firms' total assets (defined as the sum of current assets and noncurrent assets), tangible fixed assets (property, plant, and equipment [PPE]), total fixed assets, the equity-to-asset ratio, the return on assets, the number of employees, and capital expenditure.

Our proxy for real estate collateral at the firm level, or collateral for brevity, which plays a critical role in the second part of our empirical analysis, is the share of *tangible fixed* assets in *total* assets. Unfortunately, the German credit registry does not include information on collateral. In addition, Amadeus data do not provide separate information on buildings, land and improvements, and construction in progress; these are the three categories of tangible fixed assets that are usually considered corporate real estate assets in accounting definitions. However, for the United States, real estate is estimated to be a sizable fraction of total fixed assets, total assets, and firms' market values for publicly listed companies (Chaney, Sraer, and Thesmar 2012; Nelson, Potter, and Wilde 2000). Real estate assets are usually assumed to be even more important for private firms. Moreover, Laposa and Charlton (2002) estimate that European corporate holdings of real estate assets of public companies are even higher (as a share of total assets) than in the United States due to the underdevelopment of the property management industry. For example, recent estimates of the share of

[8] A legal borrowing entity comprises independent borrowers who are legally or economically connected to each other because of majority ownership (more than 50%) or because of profit transfer agreements. Consequently, the effective reporting threshold is usually lower than 1 million euro. The official reporting threshold was lowered from 1.5 million to 1 million euro in 2014. Because of the relatively low effective reporting threshold, however, this reduction does not affect our results.

real estate assets in total assets for German public companies, up to 2013, show substantial variation across sectors and, unlike the United States, limited decline over time during our sample period (Rochdi 2015). Hence, one clear advantage of using total fixed tangible assets from Amadeus is that this indicator is available not only for publicly listed companies but also for smaller and private firms.

The data matching at the bank-firm level is challenging because the German credit register and the Amadeus database do not share a common identifier. We proceed as follows. First, we match by the unique commercial register number when it is available. Second, for observations without this identifier, we rely on Stata's *reclink* command. At this step, we match firms either by their name and ZIP code or by their name and city, with a minimum matching reliability of 0.99. We then match the remaining firms manually.[9] Overall, we can track the records of more than 44% of German firms included in the credit register during the sample period, slightly more than in previous studies using these data (Behn, Haselmann, and Wachtel 2016).[10]

To focus on commercial banks, we exclude investment funds and special purpose vehicles that are less likely to grant traditional loans. The resultant sample after this adjustment comprises approximately 700,000 bank-firm-quarter observations.

## 2. Empirical Strategy

Figure 2 summarizes how capital flows affect the economy through multiple channels at the business cycle frequency. Capital flow shocks can loosen domestic financial conditions and increase credit supply. Increased credit supply stimulates real estate markets and property prices. Higher property prices can amplify the initial credit impulse through collateral channels on the household or the firm side, driving investment, employment, and other firm outcomes.[11]

The main focus of this paper is the outer loop in Figure 2. The central hypothesis is that the tighter a city's real estate markets are, the more significant is the impact of bank flow shocks on the city's output growth. In a given local real estate market, all else equal, a tighter market results in a higher sensitivity

[9] We matched 4,143 firms in the first step, 23,010 firms in the second step, and 1,038 firms by hand and hence more than 28,000 in total.

[10] Table A4 provides summary statistics comparing matched and not matched firms. Matched firms are larger, have higher shares of tangible assets, and have lower equity ratios and returns on assets, possibly indicating that firms in our sample are older and more mature.

[11] Among others, see Mian, Sufi, and Verner (2017), Hoffmann and Stewen (2020), and di Giovanni et al. (2018) on capital flows and credit supply; see Favara and Imbs (2015) and Maggio and Kermani (2017) on credit and property prices; see Iacoviello (2005) and Liu, Wang, and Zha (2013) for general equilibrium models of amplification via real estate collateral on the household or the firm side, respectively; and see Chaney, Sraer, and Thesmar (2012), Gan (2007), Cvijanović (2014), and Adelino, Schoar, and Severino (2015) on property price, collateral, and firm outcomes.

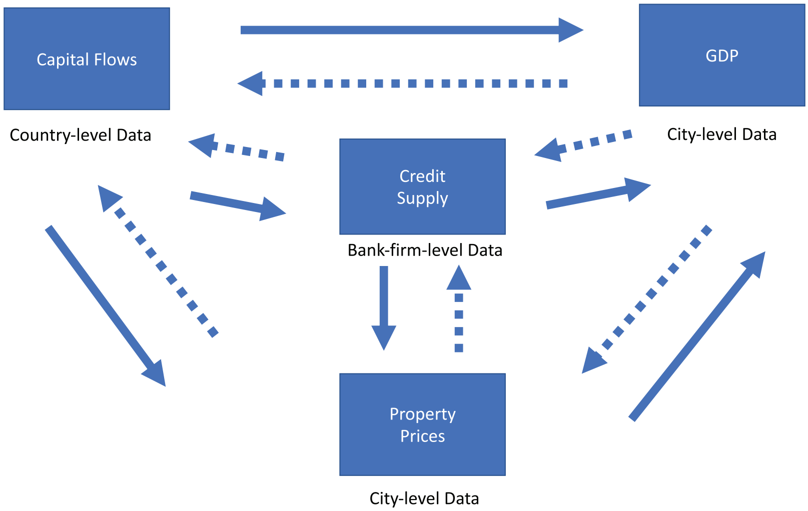


**Figure 2**
**Transmission mechanism: Roadmap**
The solid arrows represent causal linkages and the dashed arrows reverse causal effects. The top arrows represent the traditional push-pull view of the short-run association between capital flows and the business cycle. The *inner* loop emphasizes the role of credit in the transmission, and its two-way connection with property prices. The outer loop represents the possible role of real estate markets in the transmission of capital flow shocks.

of property prices to housing demand and supply shocks. Moreover, consistent with macroeconomic models with borrowing constraints as in Kiyotaki and Moore (1997), Iacoviello (2005), and Liu, Wang, and Zha (2013) in which real estate serves as collateral, as well as abundant evidence on the strong aggregate *comovement* between capital inflows, real estate prices and output, our prior is that property price increases should play an important role in the transmission of bank flow shocks to city output growth by inflating the borrowing capacity of households and firms.

The econometric challenge in investigating the outer loop of Figure 2 by exploiting city data variation is to find a measure of exposure to real estate market tightness that varies quasi-randomly across cities. Equipped with such a measure, we can then use the predicted component of property price changes triggered by an aggregate capital flow shock to estimate the *causal* impact on output. Taken together, the two steps provide an estimate of the causal effects of capital flow shocks on city output growth *through* property price changes. Thus, the strategy in the first and main part of the empirical analysis, based on city-level data, is one of identification by geographic variation. The research design, here, is grounded on (a) the availability of a well-defined aggregate or nationwide measure of capital flows and (b) the construction of an indicator of real estate market tightness (or exposure) that varies randomly across cities, as discussed below.

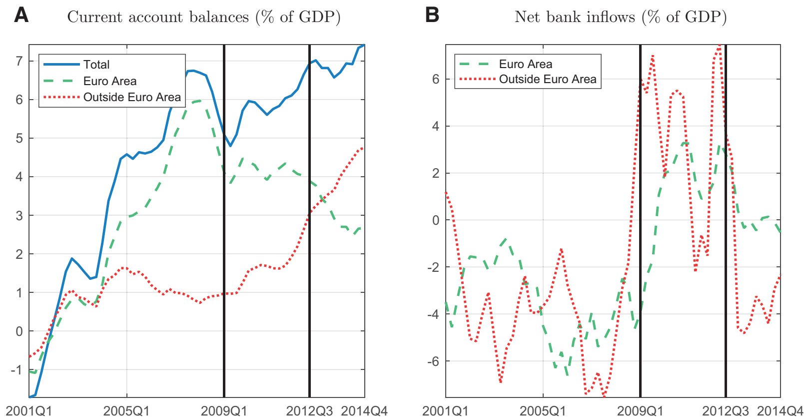


**Figure 3**
**Current account balance and net bank flows**
Panel A plots the current account balance as a share of GDP, together with its breakdown versus the rest of the euro area and outside the euro area. Panel B plots net bank flows versus the rest of the euro area and versus outside the euro based on BIS Locational Statistics. The vertical bars represent the beginning of the post-GFC recovery in 2009:Q1 and the "Whatever It Takes" speech by ECB Governor Draghi in 2012:Q3, respectively. See the appendix for variable definitions and data sources.

Although the inner loop in Figure 2 is not our main focus, in the second part of the paper, we want to explore the transmission mechanism underlying our reduced-form estimates. In particular, we will provide evidence on the role of real estate collateral in the allocation of any credit supply increase triggered by the capital flow shock (see the left-hand side of the inner loop of Figure 2). Finally, we will also provide evidence on firm employment and investment decisions (see the right-hand side of the inner loop of Figure 2). The empirical strategy to address reverse causation, here, will rely on the availability of matched bank-firm-level data in regression designs typically used in the empirical literature on bank and firm behavior.

## 2.1 Measuring capital flows

As measured by the current account surplus of the balance of payments, Germany experienced sizable net capital *outflows* rather than *inflows* throughout the period we consider (Figure 3, panel A). The current account balance, therefore, is not a suitable measure for our empirical analysis. From this figure, however, we can also see that the current account surplus vis-a-vis the *rest of the euro area* started to decline during the GFC, and continued in that direction throughout the rest of the sample period. In contrast, the current account surplus vis-a-vis the *rest of the world outside the euro area* became even larger after 2009:Q1. Moreover, panel B of Figure 3 shows that the net foreign asset position of German BIS reporting banks changed dramatically during and after the GFC. So we will focus on *cross-border bank flows*, labeled "bank flows" for brevity, which are an important component of total flows.

Aggregate cross-border bank flow data pose their own challenges because subject to measurement errors and contaminated by foreign currency valuation effects that are difficult to account for. Moreover, our sample period is rather short from a time series perspective. An alternative measurement approach, often employed in the extant literature, is to use price-based indicators that comove closely with quantity-based measures of bank flows (Rey 2013). One indicator often employed to capture bank flows driven by global risk or risk aversion is the U.S. VIX index of implied equity market volatility (Forbes and Warnock 2012).

Following this approach, and also consistent with theoretical models of capital repatriation or retrenchment (Caballero and Simsek 2020), as a proxy for bank flows, we use an indicator of financial instability and risk in Southern Europe, namely, the average sovereign bond spread of Portugal, Italy, Greece, and Spain versus Germany; henceforth, this will be called the GIPS spread. The GIPS spread is plotted in Figure 4, together with German bank flows versus the rest of the euro area from panel B of Figure 3. The figure shows that these two variables correlate closely, with aligned turning points around the milestones of the sovereign debt and banking crisis in Southern Europe.

This association is also strongly corroborated by a formal econometric analysis reported in Appendix B using aggregate and bank-level data and distinguishing between net and gross bank flows. The evidence reported confirms that German banks experienced a sizable net inflow of capital from the rest of the euro area since the GFC, driven by German banks' repatriation of foreign assets, consistent with available evidence on the behavior of Northern European banks before and after the GFC discussed above. The evidence also shows that the GIPS spread is a good predictor of bank flows. Based on these findings, in the rest of the paper, we will use the GIPS spread as our proxy for bank and hence capital flows.

## 2.2 City exposure to real estate market tightness

Identification by geographic variation requires a measure of exposure to real estate market tightness that varies quasi-randomly across cities. We want to find a measure that can work in both residential and commercial sectors, and we exploit the richness of our property price data to run a horse race between the two sectors. For this purpose, we propose to use the 2008 value of the *product* of two variables: a supply scarcity indicator in the spirit of Saiz (2010) and Hilber and Vermeulen (2016) and an indicator of excess demand, consistent with the notion that foreign buyers (Badarinza and Ramadorai 2018; Favilukis, Ludvigson, and Nieuwerburgh 2017) and immigrants (Saiz 2003; Saiz and Wachter 2011) can affect property prices.[12] On the one hand, Table 1 shows that the supply side component of this measure works well as candidate instrument

[12] We use the product rather than sum of these two variables to ease the interpretation in the empirical analysis so that the unit of measure of the two components is the same.

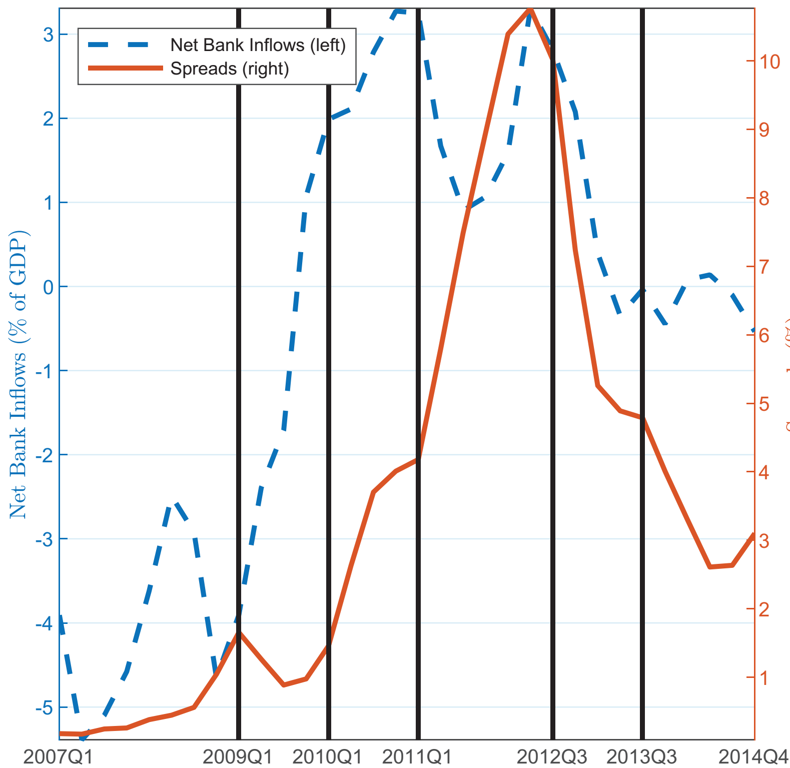


**Figure 4**
**GIPS spread, net bank flows, and the European crisis**
The figure plots the GIPS spreads and net bank inflows (% of GDP) versus the rest of the euro area. The five vertical lines represent the following events: (1) the beginning of the German recovery in 2009:Q1; (2) Greek bonds downgraded to junk status and the Troika's launch of the 2010 110 billion euro bailout; (3) 2011 downgrade and euro area leaders' disagreement on the rescue package for Greece; (4) "Whatever It Takes" speech by ECB Governor Draghi; and (5) interest rate cuts by the ECB. See the appendix for variable definitions and data sources.

only in the residential sector of West Germany in our data sample. On the other hand, in Appendix D, we show that the demand component of our exposure measure has a theoretically unambiguous impact only on property prices in the commercial sector. So, by combining these two components, we obtain an exposure measure that is a potentially suitable instrument for both sectors in our reduced-form regressions as Table 1 illustrates. We now discuss each of the two components in turn.

**2.2.1 Nondevelopable area.** To capture *total* real estate supply variation across cities due to the natural and man-made limits, such as geography and land-use regulations imposed on cities' boundaries, we employ the 2008 value of the fraction of a city's reference area covered by water, agriculture, forest and other open areas not cultivated or devoted to mining, adding up to

**Table 1**
**Property prices and alternative instrument components**

| | Full sample | | West | | East | |
|---|---|---|---|---|---|---|
| | RREP | CREP | RREP | CREP | RREP | CREP |
| **Open space** | 0.01 (0.78) | −0.04 (0.35) | 0.11 (0.02) | 0.03 (0.62) | −0.09 (0.33) | −0.64 (0.00) |
| of which: Water | 0.06 (0.18) | −0.03 (0.52) | 0.10 (0.06) | −0.08 (0.13) | 0.07 (0.47) | 0.02 (0.82) |
| of which: Agriculture | 0.11 (0.01) | 0.04 (0.35) | 0.13 (0.01) | 0.01 (0.83) | 0.23 (0.02) | 0.08 (0.44) |
| of which: Forest | −0.11 (0.01) | −0.07 (0.10) | −0.03 (0.51) | 0.04 (0.38) | −0.30 (0.00) | −0.53 (0.00) |
| of which: Other open space | −0.06 (0.20) | −0.02 (0.65) | −0.06 (0.23) | −0.16 (0.01) | 0.04 (0.69) | 0.12 (0.20) |
| **Refugees** | 0.13 (0.00) | 0.19 (0.00) | 0.11 (0.02) | 0.19 (0.00) | 0.16 (0.09) | 0.61 (0.00) |
| **Exposure** | 0.12 (0.01) | 0.25 (0.00) | 0.08 (0.10) | 0.25 (0.00) | 0.24 (0.01) | 0.58 (0.00) |

The table reports the panel correlation between each indicator at its 2008 value and the annual values of the residential and commercial property price indexes, while distinguishing between West and East Germany. $RREP$ and $CREP$ denote the residential and commercial real estate price indexes, respectively. $p$-values appear in parentheses. See the appendix for variable definitions and data sources.

open space. Thus, this variable, which we call nondevelopable area, is the complement of total settled area, which in turn includes built-up areas of different kind (transportation, residential, commercial, etc.) and urban open spaces for recreational and other social purposes.

Despite the quality of the land cover data that we use, based on combined geolocation, survey, and statistical information, our supply side indicator is positively and significantly associated with property prices only in the residential sector of West German cities (Table 1). This is due to a strong negative correlation between forest area and property prices in East Germany, which may capture urban sprawl in the sense of Ehrlich, Hilber, and Schöni (2018) rather than supply scarcity.[13] In addition, in Germany, land-use regulations are distributed more uniformly than in the United States (Schmidt and Buehler 2007). This is why using only a measure of housing supply scarcity in the spirit of Saiz (2010) would not yield a relevant instrument for our purposes. As a result, it is useful to interact this variable with a demand side indicator.

**2.2.2 The city allocation of refugees: Validity and relevance.** The demand component of our exposure measure is the share of refugees allocated by government policy to a given city relative to the total number of refugees in the country in 2008, which we call the share of refugees.[14] Unlike other migration flows, the share of refugees varies quasi-randomly across cities with respect to the business cycle because of the features of a long-standing German policy framework for their allocation.

[13] Our empirical results are robust to using the share of land covered by water and agriculture, but the estimates are more precise with the broader measure that we use. This specific proxy also has the advantage that it can be clearly interpreted as a city's boundary.

[14] Although both the city nondevelopable area and the refugee shares have little or no time variation, the total number of refugees increased significantly over time. So, to rule out from the outset any confounding time-varying effects from this common shock, we hold the exposure measure constant at its presample value of 2008.

This city characteristic is a potentially relevant instrument for property prices in both sectors as sheltering a larger proportion of refugees in a given city means utilizing more residential and commercial real estate services otherwise available for consumption or production by other economic agents. However, while the property price impact of a higher refugee share is unambiguously positive in the commercial sector, in a simple model of city real estate markets that is presented in Appendix D, we show that the impact in the residential sector also depends on other city characteristics.

**Validity** The share of refugees is a new candidate instrument for property prices. So, it is important to discuss critical aspects of the institutional framework underpinning its validity. Germany allocates refugees across states and cities according to long-standing federal laws and state regulations governing asylum seeking, the granting of refugee status and their benefit entitlements, including housing.[15] The well-known federal *Koenigsteiner Schluessel* rule determines annually quotas for the distribution of refugees across *states* based on state population in total population (with a weight of one-third) and the percentage of state tax revenue in total revenue over the previous 2 years (with a weight of two-thirds).[16] Thus, because of the rule's dependency on past tax revenue, the *state* allocation of refugees could be endogenous to business cycle conditions.

Noteworthy here is that Berlin, Bremen, and Hamburg are city-states and do not have independent within-state allocation criteria. Berlin and Hamburg are also among the largest German cities and have the highest share of refugees (not reported). Because of the potential endogeneity of the refugee allocation to economic activity, therefore, we exclude these three city-states from the sample in our empirical analysis. Indeed, in unreported regressions, we find that including city-states biases our estimates downward.[17]

In all other states, as can be seen from Table 2, which summarizes the criteria for the city allocation of refugees within each state, the within-state allocation rules do not depend on tax revenue. Individual states have similar, but not identical city allocation systems. Although there is some heterogeneity, most states determine the refugee allocation across cities based only on state population shares. Some also use the state share of city area as a criterion. Neither criterion, however, depends on growth outcomes at the business cycle frequency. In particular, none of the non-city-state uses lagged tax revenue, although Brandenburg employs the number of employees as secondary criterion. It is important to stress also that cities have no influence

[15] Similar allocation rules apply in other Northern European countries, such as Norway and Denmark. See, for instance, Eckert, Walsh, and Hejlesen (2018) and Dustmann, Vasiljeva, and Damm (2019).

[16] This rule was established in 1949 and is used to allocate other contributions or resources across states, such as the share of federal funding to universities and research institutions.

[17] Note here that Bremerhaven is classified as an independent city in the city-state of Bremen in the national accounts and the refugee statistics and is included in our sample of 85 cities. However, all our empirical results below are robust if, in addition to Bremen itself, we also drop Bremerhaven.

**Table 2**
**Within-state refugee allocation criteria and housing solutions (%)**

| State | Allocation criteria (% of state total) | Refugees in independent accommodations (Self-reported) | Refugees in independent accommodations (official statistics) |
|---|---|---|---|
| Baden-Württemberg | Population | 35.0 | 33.5 |
| Bavaria | Population | 32.0 | 48.0 |
| Brandenburg | Population, employees | 30.0 | 34.1 |
| Hesse | Population | 50.0 | 45.5 |
| Lower Saxony | Population | 67.0 | 83.6 |
| Mecklenburg-Vorpommern | Population | 71.0 | 48.6 |
| North Rhine-Westphalia | Population, total area | 63.0 | 50.3 |
| Rhineland-Palatinate | Population | 78.0 | 90.6 |
| Saarland | Population | 79.0 | 42.7 |
| Saxony | Population | 53.0 | 34.2 |
| Saxony-Anhalt | Population | 72.0 | 45.8 |
| Schleswig-Holstein | Population | 62.0 | 90.9 |
| Thuringia | Population in 1998 | 57.0 | 49.1 |
| Berlin[a] | NA | 17.0 | 57.8 |
| Bremen[a] | NA | 60.0 | 71.6 |
| Hamburg[a] | NA | 25.0 | 64.5 |
| Average all cities | | 53.2 | 55.7 |
| Average city-states | | 34.0 | 64.6 |
| Average non-city-states | | 57.6 | 53.6 |

The table describes the refugee allocation criteria across *cites* within *all* 16 German states, based on Müller (2013). The table also shows the share of refugees housed in independent accommodations, such as apartments and single-family homes, as opposed to accommodation centers and other publicly run facilities using commercial real estate space, based on data provided by Baier and Siegert (2018) and Wendel (2014). "Self-Reported" refugees in independent accommodations is the estimate provided in Baier and Siegert (2018), based on a survey of about 4,500 asylum seekers that entered Germany between January 2013 and January 2016. "Official Statistics" refugees in independent accommodations is an estimate for 2013 based on official statistics reported by Wendel (2014). [a] Indicates that Berlin, Bremen, and Hamburg are city-states and do not have independent within-state allocation criteria.

on the characteristics of the allocated refugees, such as the country of origin, skill and education levels, or other characteristics.

The predictability and efficiency of this system are well-known with small deviations from the assigned quota at the state level (Katz, Noring, and Garrelts, 2016). While there is no hard evidence on secondary migrations, in Appendix C, we argue that significant mobility across cities within states is quite unlikely. Indeed, the panel correlation between the city share of state population in 2008, the critical criterion of the state allocation rules, and the annual de facto city share of state refugees during the period 2009–2014 is 0.91. Moreover, in Table 2, we can also see that the difference between the average share of refugees in independent accommodation based on self-reporting is very close to the measured one in non-city-states.[18] In contrast, in the city-states, there is a marked difference, consistent with media and policy reports about refugee mobility into the cities of Hamburg and Berlin (Katz, Noring, and Garrelts 2016; OECD 2017a).

[18] The difference between the self-reported and the measured share of refugees housed in independent accommodation in Table 2 is informative on this issue to the extent to which, as we argue in Appendix C, substantial illegal mobility would have to show up in a higher fraction of refugees in independent accommodations.

As we also detail in Appendix C, unlike other countries with pervasive illegal immigration, in Germany, refugees cannot have a significant impact on the labor market at business cycle frequency, even after they obtain a work permit or reach status, and hence constraining their mobility opportunities. The reasons are legal and language barriers to formal and illegal employment, respectively. For example, Brell, Dustmann, and Preston (2020) estimate that the rate of employment of refugees is essentially nil within a year, less than 14% within 2 years after they receive their work permit compared to 60% for other immigrants—the second lowest rate in Europe, after Finland—and increasing only to about 25% within 6 years. Katz, Noring, and Garrelts (2016) report that it takes 14 years for refugees to attain employment levels comparable to natives in Germany. Moreover, most of these jobs are temporary and low-skill, with wages at 40%–50% of the native level, compared to 90%–95% for other immigrants, making it very difficult to rent in the open market (see also OECD 2017a).

**Relevance** The key reason the city share of refugees can be a relevant instrument in the *commercial* sector is that refugee reception centers compete for the use of built-up structures and developable land for other uses, such as Hotels and Resorts (classified as Accommodation), Health Care, Transportation, and Warehousing. Theoretically, this is intuitive as a higher share of refugees housed in accommodation centers reduces the net supply of real estate services for other productive uses in the local economy.

Indeed, in Germany, the government provides both short-term shelter in accommodation centers for asylum seekers and long-term affordable housing for refugees who cannot self-sustain financially, ultimately reducing the net supply of both commercial and residential real estate services, which is inelastic in the short term. For instance, anecdotal evidence suggests that cities initially use up hotel capacity, school gymnasiums, retirement homes and other facilities devoted to social housing and idle spaces before being able to expand the affordable housing supply, which requires approval, design and construction and funding.

In Germany, there are institutional reasons to presume that the city share of refugees also could be a relevant instrument for property prices in the *residential* sector. Once asylum seekers reach status, refugees, who can self-sustain, must seek housing in the open market. Refugees, who cannot self-sustain, continue to be housed in collective living facilities or they can be granted the right to independent accommodation depending on the public interest and individual circumstances, with the decision at the discretion of the local government.

Nonetheless, as the simple model of local commercial and residential property markets that we present in Appendix D illustrates, the impact of a higher share of refugees on residential property prices is theoretically ambiguous if commercial real estate services are a factor of production. A higher share of refugees has an unambiguous positive impact on *commercial* property prices because a higher share of refugees reduces the net supply of real estate services for other uses in both sectors of the local economy. But a lower net supply of commercial services also lowers output and income. In addition, the fiscal transfer needed to finance refugee benefits might further affect disposable income of residents. As a result, while a higher share of refugees reduces the net supply of residential services, lower output and a higher fiscal transfer reduce local disposable income and hence the consumption of housing services with a countervailing effect on residential prices. The net effect of these two forces depends on parameter values. When we calibrate our simple model to capture the institutional details of the German policy framework (without city-states) and other salient features of our city-level data, we find that, with a higher share of refugees, residential property prices fall, rather than increase (Figure D1 in the appendix). Interestingly, this counter-intuitive result is consistent with neighborhood-level evidence from the German city of Hamburg in Dehos and Eilers (2018) and mixed evidence on the impact of immigration and refugees in the residential sector as in Saiz and Wachter (2011) and several other studies reviewed in Balkan et al. (2018).

In line with this analysis, we find that the association between the share of refugees and property prices is weakest in the residential sector of the West German cities (Table 1), thus strengthening the case for using the product of nondevelopable area and the share of refugees as exposure measure. Indeed, Table 1 shows that the product of the two variables does relatively well in both property sectors, and better than refugees alone in the commercial sector of the West and the residential one in the East, even though it does not strengthen the association with residential property prices in the West.[19]

In sum, the institutional details of the German policy framework for the allocation of refugees and the auxiliary evidence discussed suggest that our measure of exposure to real estate market tightness is a good candidate instrumental variable for our empirical analysis. In the next section, therefore, we will use this variable interacted with the GIPS spread to investigate the role of real estate markets, and property price changes more specifically, in the transmission of bank flow shocks to city output growth as captured by the GIPS spread.

## 3. Bank Flows, Real Estate Markets, and City Business Cycles

The hypothesis in the paper is that higher property prices, *triggered* by aggregate capital flow shocks, may have a stronger impact on output growth in cities with tighter real estate markets. In this section, we investigate this hypothesis empirically, for both the residential and the commercial sectors, exploiting

[19] Appendix Table A3 reports the 2008 values of the share of refugees, nondevelopable area, and their product for all 85 cities.

the quasi-random variation across cities in our measure of real estate market tightness to achieve identification.

Our main "instrument" is the interaction of the aggregate bank flow, as captured by the GIPS spread, with the city-level measure of exposure to real estate market tightness in 2008. While the GIPS spread is likely endogenous to economic conditions in individual German cities in which banking activity is concentrated, its interaction with the exposure measure, whose city distribution is assumed to be orthogonal to local and aggregate economic conditions, provides an exogenous source of variation in the intensity with which the bank flow shock impacts different cities' level of economic activity.

### 3.1 Reduced-form estimates

Equipped with a proxy for bank flows and a measure of city exposure to real estate market tightness, we start by estimating the following simple city-level reduced-form regression that does not distinguish between commercial and residential sectors:

$$\Delta GDP_{c,t} = \alpha_c + \alpha_t + \beta \cdot (\text{Spread}_{t-1} \times \text{Exposure}_{c,2008}) + \varepsilon_{c,t}, \tag{1}$$

where $GDP_{c,t}$ is log real GDP per capita in city $c$ at time $t$, $\text{Spread}_{t-1}$ is our proxy for bank inflows at time $t-1$, and $\text{Exposure}_{c,2008}$ is the value of our exposure measure to local real estate market tightness in 2008. The latter is assumed to be uncorrelated with the error term, $\varepsilon_{c,t}$. Heteroscedasticity-robust standard errors are clustered at the city level.

Table 3 displays the regression results. As a benchmark, column 1 reports an estimate of the coefficient of interest, $\beta$, without time or city fixed effects. The regression in Column 2 is saturated with city and time fixed effects to control for the direct influence of city-specific factors and common shocks, such as city size and common business cycle factors across cites. City size is particularly important because larger cities tend to grow faster due to agglomeration forces. In addition to fixed effects, columns 3–6 control for the interactions with other common shocks that might also transmit through the local real estate markets. With these latter specifications, we ensure that the GIPS spread does not capture the impact on city output growth of the ECB's monetary policy response to the European crisis or the drop in the German Bund term premium that we can see in panel B of Figure 1, which also drove the fall in German mortgage rates (not reported) during the sample period. This is important because these confounding factors are likely to be correlated with the capital flow shock and could affect cities differently depending on the degree of exposure to real estate market tightness. Finally, columns 7 and 8 add to the specification in column 3 the interaction of the GIPS spread with the city-level share of population or population density. These last two specifications rule out that the exposure measure identifies differences between large and small cities or congested and noncongested cities, rather than real estate market pressure as intended.

**Table 3**
**Bank flows, real estate exposure, and city output growth: Reduced-form estimates**

| | (1) ΔGDP | (2) ΔGDP | (3) ΔGDP | (4) ΔGDP | (5) ΔGDP | (6) ΔGDP | (7) ΔGDP | (8) ΔGDP |
|---|---|---|---|---|---|---|---|---|
| $\text{Spread}_{t-1}$ | −0.131** | - | - | - | - | - | - | - |
| | (0.062) | | | | | | | |
| $\text{Exposure}_{c,2008} \times \text{Spread}_{t-1}$ | 0.004** | 0.004** | 0.010** | 0.009*** | 0.008*** | 0.008*** | 0.007* | 0.011** |
| | (0.002) | (0.002) | (0.004) | (0.003) | (0.003) | (0.003) | (0.004) | (0.004) |
| $\text{Exposure}_{c,2008} \times \text{Bund}_{t-1}$ | | | 0.026* | | | | 0.026* | 0.026* |
| | | | (0.014) | | | | (0.014) | (0.014) |
| $\text{Exposure}_{c,2008} \times \text{Eonia}_{t-1}$ | | | | 0.018** | | | | |
| | | | | (0.008) | | | | |
| $\text{Exposure}_{c,2008} \times \text{ECB Rate}_{t-1}$ | | | | | 0.030** | | | |
| | | | | | (0.014) | | | |
| $\text{Exposure}_{c,2008} \times \text{VIX}_{t-1}$ | | | | | | 0.002* | | |
| | | | | | | (0.001) | | |
| $\text{Pop.}_{2008} \times \text{Spread}_{t-1}$ | | | | | | | 0.263 | |
| | | | | | | | (0.160) | |
| $\text{Pop. Dens.}_{2008} \times \text{Spread}_{t-1}$ | | | | | | | | −0.000 |
| | | | | | | | | (0.000) |
| $\text{Exposure}_{c,2008}$ | −0.036*** | – | – | – | – | – | – | – |
| | (0.011) | | | | | | | |
| Time FE | No | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| City FE | No | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Obs. | 510 | 510 | 510 | 510 | 510 | 510 | 510 | 510 |
| $R^2$ | .010 | .396 | .398 | .403 | .401 | .398 | .399 | .398 |

The regressions are based on annual city-level data over the period 2009–2014 for our sample of 85 matched cities and implemented with STATA's reghdfe command. The dependent variable is real GDP per capita growth. The regressors are the lagged values of the GIPS spread, the 2008 city-level value of our exposure measure, and the interaction between the two. The regression in column 1 is without city or time fixed effects. Column 2 includes both. Columns 3–6 control for the interactions between the German 10-year Bund yield, the Eonia interbank rate, the ECB policy rate, and the VIX index and our exposure measure, respectively. Columns 7 and 8 add to the specification in Column 3 a control for the interaction between the GIPS spread and the 2008 level of city population or population density. Heteroscedasticity-robust standard errors clustered at the city level are shown in parentheses. *$p<.1$; **$p<.05$; ***$p<.01$.

The estimated $\beta$ coefficient in columns 1 and 2 is 0.004, is statistically significant at the 5% level and is not affected by saturating the regression with fixed effects. Remarkably, Table 3 shows that the estimated impact of the capital flow shock on city output growth becomes stronger when we control for confounding factors. In particular, the impact of the capital flow shock is more than twice as large if we control for the confounding impact of the drop in the German Bund yield (Column 3), which can be affected by the ECB policy rate, interbank conditions, as well as the Bund term premium. Restricting attention to the European interbank market (in Column 4) or the ECB's monetary policy (in Column 5), as measured by the Eonia rate and the ECB policy rate, respectively, we see that the size of the GIPS spread impact remains above 0.008 and is estimated even more precisely. Thus, the results show that monetary policy and other common factors in the European interbank or the German Bund markets do not absorb the effects of the capital flow shock itself. Controlling for the VIX index and its interaction with our exposure measure, as is often done in the extant literature, does not weaken our results, consistent with our claim that the GIPS spread captures inflows from the rest of the euro area.

Columns 7 and 8 are the most conservative specifications, as they not only hold the concomitant presence of all other common shocks considered in columns 3–6 constant but also control for demographic factors in the interaction term. The results in columns 7 and 8 show that the regression coefficient on the interaction between the GIPS spread and our exposure measure has its highest value when we hold population density constant, and it is only slightly smaller than in columns 3–6 when we control for population shares, even though the statistical significance decreases in column 7, reflecting possible collinearity between population shares and refugee shares.[20] Note, however, that neither the population share nor the population density plays a separate role in explaining the differential impact of the capital flow shock on city growth, conditional on the bank flow shock that we identify.

**3.1.1 The economic magnitude of the estimated impacts.** The economic magnitude of the estimated impact is sizable. Considering the lower bound of our estimates in Column 2, our results imply that, for every 100-bp increase in the GIPS spread, output growth in cities at the 75th percentile of the exposure distribution (e.g., Kassel) is, on average over the sample period, 8 bp higher than in cities at the 25th percentile (e.g., Bottrop).[21] The average increase in the GIPS spread over our sample period, which was 390 bps (from an average value of 0.7% in 2008 to an average of 4.6% during 2009–2014), implies that cities most exposed to real estate market tightness might have grown 31 bps more on average per year than the least exposed cities during that period, or 1.9 percentage points more cumulatively between 2009 and 2014, a cumulative growth differential that rises up to 5 percentage points cumulatively if we consider the point estimate in column 8.

**3.1.2 Placebo regressions and other robustness checks.** The reduced-form results are remarkably robust. Although our treatment variable is continuous and in the paper, we did not pursue a difference-in-differences research design, in Table 4, we report results for an estimated placebo regression in this spirit. We do so by examining the reduced-form impact of the interaction between the GIPS spread and our exposure measure on city-level output growth in a presample of the same length as our sample period, before the beginning of the German real estate boom.[22] We consider two ending years, 2007 and 2008, to control for the timing of the GFC. The placebo regressions control for the

[20] Recall here that the refugee state allocation criteria use city shares of state population, while here we are holding constant city shares of country population.

[21] Cities at the 75th percentile of the distribution have an exposure value of 27.23. Thus, the output growth effect of a 100-bp GIPS spread change is 11=(100*27.23*0.004) bps. In contrast, cities at the 25th percentile have an exposure value of 8.35. Hence, in this case, the impact is 3=(100*8.35*0.004) bps.

[22] As Figure 1 illustrates, the price-to-rent ratio turned around in 2009 in the commercial sector and only in 2012 in the residential one.

**Table 4**
**Bank flows, real estate exposure, and city output growth: Placebo regressions**

| | 2009–2014 | 2009–2014 | 2009–2014 | 2002–2007 | 2002–2007 | 2003–2008 | 2003–2008 | 2009–2014 | 2009–2014 |
|---|---|---|---|---|---|---|---|---|---|
| | (1) $\Delta$GDP | (2) $\Delta$GDP | (3) $\Delta$GDP | (4) $\Delta$GDP | (5) $\Delta$GDP | (6) $\Delta$GDP | (7) $\Delta$GDP | (8) $\Delta$GDP | (9) $\Delta$GDP |
| $\text{Exposure}_{c,2008} \times \text{Spread}_{t-1}$ | 0.004** | 0.007 | 0.011** | −0.020 | 0.095 | −0.418 | 0.130 | | |
| | (0.002) | (0.004) | (0.004) | (0.229) | (0.144) | (0.376) | (0.373) | | |
| $\text{Exposure}_{c,2008} \times \text{Bund}_{t-1}$ | | 0.026* | 0.026* | −0.032 | −0.032 | −0.032 | −0.032 | 0.015* | −0.002 |
| | | (0.014) | (0.014) | (0.020) | (0.020) | (0.022) | (0.022) | (0.009) | (0.007) |
| $\text{Pop.}_{2008} \times \text{Spread}_{t-1}$ | | 0.328** | | 14.731 | | 51.128* | | 0.512*** | |
| | | (0.161) | | (13.841) | | (29.655) | | (0.188) | |
| $\text{Pop. dens.}_{2008} \times \text{Spread}_{t-1}$ | | | −0.000 | | 0.002 | | 0.001 | | −0.000 |
| | | | (0.000) | | (0.004) | | (0.009) | | (0.000) |
| $\text{Random exposure}_{2008} \times \text{Spread}_{t-1}$ | | | | | | | | −0.001 | −0.001 |
| | | | | | | | | (0.001) | (0.001) |
| Time FE | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| City FE | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Obs. | 510 | 510 | 510 | 510 | 510 | 510 | 510 | 510 | 510 |
| $R^2$ | .471 | .474 | .473 | .195 | .194 | .241 | .238 | .404 | .400 |

The regressions are based on annual city-level data over different time periods for our matched sample of 85 cities and implemented with STATA's reghdfe command. The dependent variable is nominal GDP per capita growth, except in columns 8 and 9 in which it is real terms as in Table 3. The regressors are as in Table 3, except in columns 8 and 9, in which the city-level exposure measure is the product of nondevelopable area and randomly assigned refugee shares drawing from a uniform distribution with a range matching the maximum and minimum in the data. All regressions include city and time fixed effects. Heteroscedasticity-robust standard errors clustered at the city level are shown in parentheses. * $p$ <.1; ** $p$ <.05; *** $p$ <.01.

interactions between exposure and the Bund yield as well as between the GIPS spread and either the city-level population shares or the city-level population density, as in columns 7 and 8 of Table 3.

As some states do not have CPI data prior to 2005, we need to perform this analysis using nominal per capita GDP growth rates. Columns 1–3 of Table 4, however, replicate Columns 2, 7, and 8 of Table 3 and show that this makes essentially no difference for the baseline results, apart from a marginal decline in the statistical significance of the parameter of interest once we control for the interaction between population levels and the GIPS spread in column 2. Columns 4–7 report the estimated placebo regressions and show that the estimated impact of capital flow shocks is statistically insignificant in both presample periods considered. This is evidence that, during periods in which local real estate markets were not tight, our identified capital flow shocks did not explain city differences in output growth.

In columns 8 and 9 of Table 4, we also report the results of a synthetic placebo regression constructed by randomizing the share of refugee assignment in our exposure measure and generating it from a uniform distribution with a range matching the maximum and minimum in our data. Again, the interaction term between the GIPS spread and this synthetic exposure measure loses its statistical significance.

As we already mentioned in Section 1.1, most reduced-form results above are also robust to extending the sample to all 127 cities for which there are commercial price indexes or the 107 urban areas in the national accounts (results not reported to conserve space, but available on request). In addition, note that our results get stronger when we drop Munich from the sample, which has the highest exposure value in Table A3. Finally, in unreported regressions, we also controlled for the differential impact of the bank flow shock across East and West German cities by interacting an East Germany dummy with the GIPS spread and our exposure measure. The triple interaction was insignificant, indicating that there is no statistically different impact left between East and West German cities.

## 3.2 Instrumental variables estimates

The reduced-form estimates in Table 3 yield evidence on the importance of real estate market tightness for output growth, but are silent on the specific role that residential or commercial property prices may play in the transmission of capital flow shocks. To shed light on this question, we now regress city output growth on property prices in either the residential or the commercial sector, instrumenting the latter with the interaction of the GIPS spread with our main exposure measure.[23]

[23] The setup is similar to the one in Chaney, Sraer, and Thesmar (2012), who interact the aggregate mortgage interest rate (our GIPS spread) with the housing supply elasticity of Saiz (2010) (our exposure measure) and then use the

**Table 5**
**Bank flows, real estate exposure, and property prices: First-stage results**

| | (1) $RREP$ | (2) $CREP$ | (3) $RREP$ | (4) $CREP$ | (5) $RREP$ | (6) $CREP$ |
|---|---|---|---|---|---|---|
| $\text{Spread}_{t-1} \times \text{Exposure}_{c,2008}$ | 0.012 | 0.029*** | | | 0.002 | 0.022*** |
| | (0.008) | (0.004) | | | (0.007) | (0.004) |
| $\text{Spread}_{t-1} \times \text{Share 25–30}_{c,2008}$ | | | 0.597*** | 0.517*** | 0.589*** | 0.431*** |
| | | | (0.080) | (0.092) | (0.079) | (0.082) |
| Time FE | Yes | Yes | Yes | Yes | Yes | Yes |
| City FE | Yes | Yes | Yes | Yes | Yes | Yes |
| F-statistic | 2.3 | 44.3 | 56.2 | 31.8 | 29.6 | 38.1 |
| Obs. | 510 | 510 | 510 | 510 | 510 | 510 |
| $R^2$ | .783 | .787 | .830 | .799 | .830 | .814 |

The table reports the estimation results for the first stage. The regressions are based on annual city-level data for 85 cities from 2009 to 2014 and implemented with STATA's reghdfe command. The dependent variables are the residential and commercial real property price indexes. The regressors are the lagged value of the GIPS spread interacted with the 2008 value of the exposure measure in columns 1 and 2, and the GIPS spread interacted with the 2008 value of the city share of population aged 25–30 in columns 3 and 4. Columns 5 and 6 include both candidate instruments in each equation. All regressions include city and time fixed effects. The heteroscedasticity-robust standard errors clustered at the city level are shown in parentheses. $p<.1$; $^{**}p<.05$; $^{***}p<.01$.

**3.2.1 First stage.** Table 5 reports a battery of alternative first-stage regressions. The specification of the first stage is

$$\text{REP}_{c,t} = \alpha_c + \alpha_t + \gamma \cdot \text{INST}_{c,t-1} + \eta_{c,t}, \tag{2}$$

where $\text{REP}_{c,t}$ is either the commercial property price index ($CREP$) or the residential price index ($RREP$), and $INST_{c,t-1}$ is the instrumental variable. All regressions are saturated with time and city fixed effects. In columns 1 and 2, the instrumental variable is our exposure measure interacted with the GIPS spread, ($\text{Spread}_{t-1} \times \text{Exposure}_{c,2008}$), as it was used in the reduced-form estimates reported in Table 3. It is the same variable in both sector regressions. The results in columns 1 and 2 show that our exposure measure is a good predictor of property prices in the commercial sector, with an F-statistic well above the norm, even after controlling for city and time fixed effects. Unfortunately, however, once interacted with the GIPS spread, the exposure measure is not a valid instrument for the residential sector, despite the statistically significant panel correlation with property prices in Table 1. In the residential sector, in fact, the F-statistic is far below acceptable levels.

As we discussed earlier, this result is theoretically plausible and consistent with the mixed evidence on the impact of immigration on residential real estate markets. So, to run a horse race between the commercial and residential sector in accounting for our reduced-form results, we need to find an alternative instrument. For this purpose, we use the city share of population in the 25–30 age bracket, called the city share of young people in 2008. The share of young people is a plausible alternative because its city distribution is stable over time

predicted component of local real estate prices to estimate their mediating effect on firm investments in response to the common shock.

and unlikely to respond to short-to-medium term business cycle conditions.[24] This variable is also a potentially relevant instrument because younger people are more likely to rent. Higher rents, in turn, can attract buy-to-let domestic and (deep-pocket) foreign investors that might have played an important role in igniting the German residential housing boom.[25]

In columns 3 and 4, therefore, we report estimates for the same specification as in equation 2, but using the interaction between the GIPS spread and the share of population in the 25–30 year age bracket, denoted ($\text{Spread}_{t-1} \times$ Share $25\text{–}30_{c,2008}$), as instrument for both sectors. The results show that this alternative instrument has good predictive power in both sectors, even though the F-statistic is now weaker in the commercial sector.[26]

In columns 5 and 6 of Table 5, we use both instruments at the same time in each property sector. The results show that only the city share of people aged 25–30 in 2008 interacted with the GIPS spread predicts differences in residential property prices across cities, while both candidate instruments are good predictors of price differences across cities in the commercial sector. In both the residential and the commercial sector, the F-statistic is well above conventional levels. Note here that the magnitude of the coefficient on the exposure measure in column 6 is only slighter smaller than the coefficient in column 2, while the $R^2$ is marginally higher. In contrast, in the residential sector, when we use both instruments, the effect of our exposure measure is soaked up by the share of young population. Equipped with these first-stage results, we proceed to report our instrumental variable estimates.

**3.2.2 Second stage.** In the second stage, we estimate the impact of property price changes triggered by capital flow shocks on output growth. We consider three alternative specifications. The first, estimated with two-stage least squares (2SLS) and following Chaney, Sraer, and Thesmar (2012), evaluates the role of residential and commercial property prices separately:

$$\Delta GDP_{c,t} = \alpha_c + \alpha_t + \delta \cdot \text{REP}_{c,t} + \varepsilon_{c,t}, \tag{3}$$

where the instrument for $\text{CREP}_{c,t}$ is ($\text{Spread}_{t-1} \times \text{Exposure}_{c,2008}$) and the instrument for $\text{RREP}_{c,t}$ is ($\text{Spread}_{t-1} \times$ Share $25\text{–}30_{c,2008}$). This first

[24] The panel correlation between the share of young people and its own lag is 98%. For example, the city with the lowest share of people in this age bracket in our sample is Suhl in East Germany, with a 2008 value of 5.1%. In 2014, this city had a 5.2% share. The city with the highest proportion of young people is Wuerzburg in West Germany, with values of 10.3% in 2008 and 11.1% in 2014.

[25] The hypothesis is consistent with the decline (increase) of a full percentage point in the German home ownership (tenancy) rate over our sample period that is reported by the Eurostat and the German Federal Statistical office data. See also Kindermann et al. (2020).

[26] Note here that we obtain these results only for the city share of population aged 25–30, but not with the share of people aged 18–25. The results are also weaker when we use the next age bracket (age of 30–50). One interpretation is that people aged 18–25 typically go to university in Germany, either living with parents or on campus, and are less likely to rent in the open market. As the age bracket increases to age 30–50, the demographic composition is more likely to shift toward buyers.

specification is also estimated instrumenting each property price index with both instrumental variables, so that the parameter of interest, $\delta$, is overidentified. Implicit in this specification is the assumption that the error terms of the residential and the commercial price equations are uncorrelated.

The second specification is

$$\Delta GDP_{c,t} = \alpha_c + \alpha_t + \delta^P \cdot \text{RREP}_{c,t} + \delta^C \cdot \text{CREP}_{c,t} + \varepsilon_{c,t}, \tag{4}$$

where the instruments are ($\text{Spread}_{t-1} \times \text{Share } 25\text{–}30_{c,2008}$) for residential prices and ($\text{Spread}_{t-1} \times \text{Exposure}_{c,2008}$) for commercial prices. This second specification, also estimated with 2SLS, permits to assess the relative importance of the residential and commercial sectors in the transmission of the bank flow shocks captured by the GIPS spread.

The third specification, estimated with 3SLS, is based on the following system of seemingly unrelated regressions:

$$\text{RREP}_{c,t} = \alpha_c + \alpha_t + \gamma^R \cdot (\text{Spread}_{t-1} \times \text{Share } 25\text{–}30_{c,2008}) + \alpha^{RR} \cdot \text{RREP}_{c,t-1} + \alpha^{RC} \cdot \text{CREP}_{c,t-1} + \eta_{c,t} \tag{5}$$

$$\text{CREP}_{c,t} = \alpha_c + \alpha_t + \gamma^C \cdot (\text{Spread}_{t-1} \times \text{Exposure}_{c,2008}) + \alpha^{CR} \cdot \text{RREP}_{c,t-1} + \alpha^{CC} \cdot \text{CREP}_{c,t-1} + \eta_{c,t} \tag{6}$$

$$\Delta GDP_{c,t} = \alpha_c + \alpha_t + \delta^R \cdot \text{RREP}_{c,t} + \delta^C \cdot \text{CREP}_{c,t} + \varepsilon_{c,t}, \tag{7}$$

which takes into explicit account the contemporaneous or lagged correlation between the residential and the commercial real estate sectors.

Table 6 reports the estimation results. The results suggest that both commercial and residential property prices variations predicted by changes in the GIPS spread can affect city output growth, although the magnitude of the impact is larger and estimated more precisely in the commercial sector (columns 1 and 2). The comparison is unaffected when we overidentify the parameters of interest by using both instruments in each regression (columns 3 and 4). Note here, however, that, for the residential sector, the test for overidentification restrictions rejects the null that those are valid at the 6% level, as compared to 58% in the commercial one.

Column 5 reports the 2SLS estimate for the output equation when we horse race the two sectors including both the residential and commercial property prices at the same time, as in Equation (4), and assigning one relevant instrument to each variable. When we put both residential and commercial prices in the regression, we find that the impact through the residential sector vanishes, with the residential sector coefficient losing its economic and statistical significance. This result is corroborated by the 3SLS estimates in column 6, where the coefficient on the residential sector even turns negatively significant at the 10% level. Although the precision of the estimates diminishes somewhat in the horse race in columns 5 and 6, because of the high correlation between

**Table 6**
**City output growth and property prices: Instrumental variable results**

| | 2SLS | 2SLS | 2SLS | 2SLS | 2SLS | 3SLS |
|---|---|---|---|---|---|---|
| | (1) ΔGDP | (2) ΔGDP | (3) ΔGDP | (4) ΔGDP | (5) ΔGDP | (6) ΔGDP |
| $RREP_t$ | 0.092* | | 0.096* | | −0.049 | −0.085* |
| | (0.054) | | (0.054) | | (0.091) | (0.048) |
| $CREP_t$ | | 0.142** | | 0.120** | 0.163* | 0.087* |
| | | (0.056) | | (0.050) | (0.088) | (0.045) |
| Time FE | Yes | Yes | Yes | Yes | Yes | Yes |
| City FE | Yes | Yes | Yes | Yes | Yes | Yes |
| Over ident. ($p$-val.) | – | – | 0.06 | 0.58 | – | – |
| Obs. | 510 | 510 | 510 | 510 | 510 | 510 |

This table reports instrumental variable estimates. The regressions are based on annual city-level data for 85 cities from 2009 to 2014 and implemented with STATA's reghdfe command. The dependent variable is real per capita GDP growth. The main regressors are the real commercial and residential property price indexes. In column 1, we instrument residential prices by the interaction term between the lagged GIPS spread and the 2008 value of the city share of people aged 25–30. In column 2, commercial prices are instrumented by the interaction between the lagged GIPS spread and the 2008 value of our exposure measure. Columns 3 and 4 employ both instruments at the same time. In column 5, we include both residential and commercial property prices in the same regression as in Equation (4), using the relevant instrument for each of these two variables. In column 6, we estimate the same regression as in column 5, but using 3SLS on the system of seemingly unrelated regressions in Equations (5) and (7). All regressions include city and time fixed effects. The heteroscedasticity-robust standard errors clustered at the city level are shown in parentheses. $*p<.1$; $**p<.05$; $***p<.01$.

commercial and residential prices in the sample (with an unconditional panel value of 0.38 and a $p$-value of .0 in our sample), the evidence is strong that commercial property prices are more important than residential prices in the transmission of the bank flow shock we focused on.[27]

### 3.3 Discussion

We saw earlier that a GIPS spread increase leads to higher real GDP growth rates in cities most exposed to real estate market tightness. The second-stage estimates reported in Table 6 suggest that commercial property price increases triggered by GIPS spread changes can account for all of this differential. To see this, for example, multiply the first-stage coefficient in column 2 of Table 5, which is 0.029, by the second-stage estimate in column 2 of Table 6, which is 0.142. The resultant product is 0.004, which is the same as the smallest of the reduced-form estimates we obtain without additional controls in column 2 of Table 3. Here, it is noteworthy that, when we estimate our reduced-form regression in Table 3 for the 401 cities and rural areas for which we have only residential property prices (because of lack of sufficient commercial real estate activity), we find that the capital flow shock has no differential impact

[27] The empirical finding begs the question of why, in contrast to the experience of the United States and several other countries, the residential real estate sector does not seem to affect output growth during the German housing boom. While answering this is beyond the scope of this paper, one possible interpretation is that the German housing boom was not associated with a credit boom through the end of our sample period in 2014. An interesting open research question, therefore, is whether this is a more general feature of cash-based housing booms, as, for instance, when they are driven by deep-pocket foreign investors or domestic commercial investors or buyer-to-let.

(results not reported). This is in line with the evidence in the second stage of the instrumental variable analysis showing that the residential sector is not part of the transmission mechanism of the capital flow shock in our baseline sample of urban areas.

Overall, these findings suggest that (a) tighter real estate markets, as captured by our exposure measure, are associated with a stronger impact of bank inflows on local economic activity, and (b) commercial property price differences across cities triggered by bank flow shocks can account for the bulk of this differential impact. These findings suggest that commercial property prices are at the heart of the transmission mechanism of the bank retrenchment episode that we studied, possibly reflecting the working of a collateral channel of transmission. In light of this, when we open up the black box of the transmission mechanism, in the rest of the paper, we will focus only on the commercial sector, exploring empirically the collateral channel on the firm side.

Before proceeding, we want to take another look at our instrument's relevance from a different perspective. We specifically ask how a small fraction of a city population may produce a seemingly sizable differential impact of capital flow shocks across cities, for a given increase in the GIPS spread. This paper does not claim that the city share of refugees or tight city boundaries drove the German real estate boom, but it contends that the housing of refugees in accommodation centers can have a statistically significant impact on commercial property prices.

Appendix D of the paper shows that this contention is plausible for two reasons. First, refugees housed in reception centers took up a nontrivial share of the annual turnover in relevant segments of the commercial real estate market. In fact, a back-of-the-envelope estimate suggests that annual new refugee inflows housed in reception centers might have absorbed roughly 5.3% of the annual average turnover in the hotel segment of the commercial real estate market during the 2009–2014 period, with possible demand spillovers to other segments of the commercial real estate sector. Second, when we feed to our model in Appendix D the actual 2008 city-level share of refugees used to construct our exposure measure, the model predicts that the impact of the city allocation of refugees on commercial property prices that has the same sign as in the data.[28]

## 4. Bank Flows, Real Estate Collateral, and Credit Supply

In this section, we study the role of real estate collateral in the credit allocation to firms following a bank flow shock. Figure 5 plots average credit growth

[28] Whether the strength of this effect in the model is smaller or larger than in the data is an open question. As we illustrate in the Appendix D.5, the model predicts a stronger correlation than in the data once we standardize the price changes. However, the model implies a smaller regression slope than in the data when we do not control for the fact that the variance of price changes in the model is much smaller than in the data.

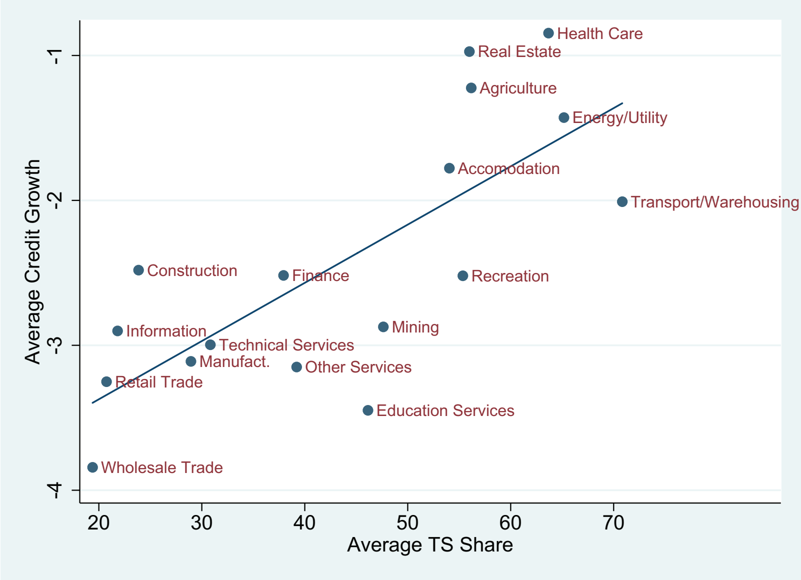


**Figure 5**
**Credit growth and real estate collateral by sector**
The figure plots average credit growth against average tangible fixed assets over total assets (TS) by industry. The industry classification corresponds to the two-digit NAICS code, with the following adjustments: Manufacturing refers to codes 31–33; Retail Trade equals refers to 44–45; Transport and Warehousing refers to codes 48–49; and Technical Services refers to 54–56. The correlation between the two variables is 75% with a $p$-value of 0. The sample period is 2009–2014.

during the 2009–2014 period by two-digit NAICS classification against the average share of tangible assets in the same sector ($TS$). The figure shows that, on average during this period, all sectors experienced a sharp credit contraction, consistent with the aggregate picture in panel D of Figure 1. However, it also suggests a positive and tight association between faster (slower) credit growth (decline) and the availability of real estate collateral, with sectors typically using land and structures more intensively, such as Agriculture, Real Estate itself, Transport and Warehousing, Accommodation, and Recreation, experiencing higher (lower) credit growth (decline).

Consistent with micro evidence on the role of collateral in financial contracting (e.g., Benmelech, Garmaise, and Moskowitz (2005); Benmelech and Bergman (2008)), we conjecture that an increase in the domestic credit supply associated with repatriation of foreign assets should benefit more firms and sectors with more real estate collateral, as this form of lending is safer and easier to screen, price and monitor. This hypothesis also accords well with related evidence on the impact of capital flow shocks on the domestic credit supply reviewed before.

As we show in Appendix B, a higher GIPS spread is associated with a reduction in German bank holdings of foreign assets in the rest of the euro area and a lower aggregate domestic lending-deposit spread. Here, we focus on the allocation of domestic credit at the bank-firm level associated with changes in the GIPS spread. Our bank-firm-level proxy for real estate collateral is the share of tangible *fixed* assets in total assets, or "share of tangible assets" for brevity, which we discussed in Section 1. To address endogeneity concerns, at this step of the empirical analysis, we rely on the microeconomic nature of our bank-firm-relationship data, assuming that no such individual relationship can affect the GIPS spread, and that the *quantity* of real estate collateral at the bank-firm-relationship level is predetermined with respect to lending decisions.

Specifically, to assess the role of real estate collateral in the credit allocation to firms, following Behn, Haselmann, and Wachtel (2016), who use German credit register data to study the impact of capital regulation on credit supply, we estimate the following reduced-form regression:

$$\Delta \mathrm{L}_{i,j,t} = \alpha_{i,t} + \alpha_{j,year} + \beta \cdot (\mathrm{SPREAD}_{t-1} \times \mathrm{TS}_{j,t-4}) + \varepsilon_{i,j,t}, \tag{8}$$

where $\Delta L_{i,j,t}$ is the log-change in loan volume of bank $i$ to firm $j$ in quarter-year $t$, and $(\mathrm{SPREAD}_{t-1} \times \mathrm{TS}_{j,t-4})$ is the lagged interaction term between the GIPS spread and firm $j$'s share of tangible assets, $\mathrm{TS}_{j,t-4}$.[29] To control for unobserved time-varying heterogeneity at the bank level, we include bank-year-quarter fixed effects ($\alpha_{i,t}$). To control for year-on-year changes in loan demand and for the location of firms' headquarters that might influence firms' credit access, we include firm-year fixed effects ($\alpha_{j,year}$). Finally, by clustering the standard errors at the bank-firm level, we allow the observations to be correlated across bank-firm relationships. The coefficient of interest is $\beta$, which captures the differential firm credit access in response to the bank flow shock.

Table 7 summarizes the baseline results, as in Equation (8), and reports a number of robustness checks. The positive and highly statistically significant estimate of $\beta$ in column 1 suggests that a higher GIPS spread leads to more bank lending to firms with more real estate collateral, controlling for loan demand with firm-year fixed effects. The magnitude of this effect is economically significant: a 100-bp GIPS spread increase raises (slows) the quarterly rate of credit growth (decline) of high-TS firms (at the 75th percentile of the distribution) by 74 bps more than the corresponding growth rate of low-TS firms (at the 25th percentile).[30]

Column 2 of Table 7 interacts the GIPS spread with the average share of tangible assets across industries. The motivation behind this specification is that industry-specific characteristics can affect the level and nature of firms' real estate asset holdings (see, for instance, Rochdi 2015). In addition, the average industry share of tangible assets is less likely to be endogenous with

[29] The tangible asset ratio is lagged by four quarters because firm-level data are annual, while bank data are quarterly.

[30] We calculate these magnitudes as follows. The 25th percentile of the distribution of $TS$ is 8.74%. The corresponding value for the 75th percentile is 65.52%. Thus, the credit growth difference between both types of firms is (65.52-8.74)*0.013=0.74.

**Table 7**
**Bank flows, real estate collateral, and firm credit access**

| | (1) $\Delta$L | (2) $\Delta$L | (3) $\Delta$L | (4) $\Delta$L | (5) $\Delta$L |
|---|---|---|---|---|---|
| $\text{Spread}_{t-1} \times \text{TS}_{t-4}$ | 0.013*** | | | | |
| | (0.003) | | | | |
| $\text{Spread}_{t-1} \times \text{TS}_{\text{Industry},t-4}$ | | 0.014*** | | | |
| | | (0.003) | | | |
| $\text{Spread}_{t-1} \times (\text{TS}_{2008} * \text{CREP}_t)$ | | | 0.010** | | |
| | | | (0.005) | | |
| $\text{Spread}_{t-1} \times \text{TS}_{t-4} \times \text{Interbank}_{t-1}$ | | | | 0.467** | |
| | | | | (0.231) | |
| $\text{TS}_{t-4} \times \text{Interbank}_{t-1}$ | | | | −3.832** | |
| | | | | (1.813) | |
| $\text{Spread}_{t-1} \times \text{TS}_{t-4} \times \text{Net foreign assets}_{t-1}$ | | | | | 0.030* |
| | | | | | (0.017) |
| $\text{TS}_{t-4} \times \text{Net foreign assets}_{t-1}$ | | | | | −0.370*** |
| | | | | | (0.122) |
| Firm-year FE | Yes | Yes | Yes | No | No |
| Firm-year-quarter FE | No | No | No | Yes | Yes |
| Bank-year-quarter FE | Yes | Yes | Yes | Yes | Yes |
| Obs. | 573,985 | 707,742 | 192,698 | 387,734 | 514,985 |
| $R^2$ | .141 | .145 | .148 | .456 | .430 |

The regressions are based on quarterly bank-firm-relationship-level data over the period 2009:Q1–2014:Q4 and implemented with STATA's reghdfe command. The dependent variable is the log-difference in loan volume of bank $i$ to firm $j$ in quarter-year pair $t$. The independent variable is the GIPS spread interacted with firms' share of tangible assets. The latter is replaced by its industry mean in column 2. In column 3, we inflate the predetermined 2008 share of firm-level tangible assets with the city-level real commercial real estate price index. Columns 1–3 include bank-time and firm-year fixed effects. Column 4 interacts the GIPS spread with firm-level tangible assets and the interbank funding over retail deposits ratio. Column 5 interacts the spread with tangible assets and banks' net foreign assets versus euro area countries over total assets. In columns 4 and 5, we replace firm-year with firm-year-quarter fixed effects. The standard errors are clustered at the bank-firm level and shown in parentheses. *$p$ <.1; **$p$ <.05; ***$p$ <.01.

respect to other firm characteristics or lending at the bank-firm level. Column 2 indicates that our results are robust to using the industry average of tangible assets. In particular, we find that banks shift credit toward firms in industries with higher shares of tangible assets. The results are virtually unchanged if we use the median industry value instead (not reported).

The specification in column 3 holds the firm share of tangible assets fixed at its 2008 level and inflates it with the city-level commercial real property price index, assuming that firms own most of their real estate assets in the city where their headquarters are located, in a manner similar to Chaney, Sraer, and Thesmar (2012) and Doerr (2018). Again, we find that banks shift their credit supply toward firms with more real estate collateral, even though the $\beta$ coefficient is now estimated less precisely. This might be the case because of the lower number of observations in this experiment, as the variable CREP is not available for all regions and cities covered by the German credit registry.

The role of two important bank characteristics is explored in columns 4 and 5. Specifically, first we examine whether the role of collateral is stronger for banks with higher interbank-to-deposit ratios. As this type of funding is more likely to be exposed to international capital market fluctuations, banks with a higher share of nondeposit funding should be more affected by changes in the GIPS spread, which captures also changes in global financial conditions. Second, we also examine the role of individual-bank pre-GFC exposure to the rest of the euro area as captured by the net foreign assets position vs the rest of the euro area as a share of total assets in 2006. If the GIPS spread is capturing bank retrenchment from Southern Europe, we should find that banks with a higher pre-GFC exposure to the rest of the euro area respond more to the spread change. To this end, we include two triple interaction terms. The first is the interaction between the GIPS spread, the share of tangible assets, and the lagged interbank-to-deposit funding ratio (Column 4). The second term interacts the spread with the share of tangible assets and the lagged value of the share of net foreign assets (Column 5).

In these two additional regressions, the granularity of the credit register data permits us to restrict the sample to firms with multiple bank relationships, and hence allowing us to include firm-year-quarter fixed effects ($\alpha_{j,t}$), as opposed to firm-year fixed effects as before. As shown by Khwaja and Mian (2008), this strategy fully absorbs firm-specific loan demand shocks.[31] The estimation results in column 4 indicate that the sensitivity of credit supply to real estate collateral is stronger for banks with a higher non-core-funding ratio, as can be gauged from the positive and statistically significant coefficient on the triple interaction term. The results in column 5 suggest that lending also might be affected by the initial euro area exposure, even though this effect is statistically significant only at the 10% level.

Table 8 reports additional robustness checks. The results in Table 7 are robust to augmenting the baseline regression with the interaction between the GIPS spread and other firm-level controls that are likely to be correlated with the firm share of tangible assets (columns 1–3). This additional experiment ensures that the baseline results are not driven by a correlation between the share of tangible assets and other firm-level characteristics. The results show that, if anything, adding more firm-level controls interacted with the GIPS spread increases the economic magnitude of the estimated coefficient on the key interaction term between the GIPS spread and tangible assets. In column 4 of Table 8, we drop observations during the 2009–2010 period. This might be important because the German government, after the GFC, provided guarantees to certain firms and sectors. To the extent to which these guarantees are correlated with the tangible asset ratio of firms, our baseline results could be biased. Column 4 shows that, even excluding 2009–2010, a higher GIPS spread leads to a shift in credit supply toward high-tangible asset firms. Finally, in column 5, we document that our results are also robust to employing a time-invariant level of TS, measured in 2008, without inflating the initial level with commercial property price changes.

[31] Recall that we can employ this identification strategy because most of the firms in our sample borrow from more than one bank. Note here that, in these two specifications, firm-time fixed effects absorb the double interaction term between the GIPS spread and the tangible asset ratio, which therefore cannot be included separately.

**Table 8**
**Bank flows, real estate collateral, firm credit access: Additional robustness checks**

| | (1)<br>ΔL | (2)<br>ΔL | (3)<br>ΔL | (4)<br>ΔL | (5)<br>ΔL |
|---|---|---|---|---|---|
| $\text{Spread}_{t-1} \times \text{TS}_{t-4}$ | 0.013*** | 0.015*** | 0.016*** | 0.014*** | 0.014*** |
| | (0.003) | (0.003) | (0.003) | (0.003) | (0.003) |
| $\text{Spread}_{t-1} \times \text{EQ}_{t-4}$ | −0.008* | −0.004 | −0.009 | | |
| | (0.003) | (0.003) | (0.003) | | |
| $\text{Spread}_{t-1} \times \text{ROA}_{t-4}$ | | −0.047*** | −0.045*** | | |
| | | (0.003) | (0.003) | | |
| $\text{Spread}_{t-1} \times \text{TA}_{t-4}$ | | | 0.128*** | | |
| | | | (0.003) | | |
| Firm-year FE | Yes | Yes | Yes | Yes | Yes |
| Bank-time FE | Yes | Yes | Yes | Yes | Yes |
| Obs. | 568,128 | 410,649 | 410,649 | 387,734 | 512,985 |
| $R^2$ | .145 | .145 | .145 | .136 | .141 |

The regressions are based on quarterly data from 2009:Q1 to 2014:Q4 and implemented with STATA's reghdfe command. The dependent variable is the log difference in loan volume of bank $i$ to firm $j$ in quarter-year $t$, as in Table 7. In Columns 1–3, in addition to $\text{Spread}_{t-1} \times \text{TS}_{t-4}$, the specification includes the interactions between the GIPS spread and the firm-level capital-to-asset ratio (Column 1), the firm-level return on assets (Column 2), and firm size measured with the log of total assets (Column 3). Column 4 drops observations for 2009 and 2010, a period in which the government intervened in the credit market via guarantees. Column 5 replaces $TS_{t-4}$ with the 2008 firm-level value of the tangible asset ratios, $TS_{2008}$, without inflating it with commercial property prices. All regressions include bank-quarter and firm-year fixed effects. The standard errors in parentheses are clustered at the bank-firm level. *$p$ <.1; **$p$ <.05; ***$p$ <.01.

The results are also robust to aggregation by sectors. Results reported in previous versions of this paper show that credit increases (declines) the fastest (the slowest) in industries with the highest shares of tangible assets. As we saw earlier, these sectors are those in which land and buildings are used more intensively in the production of their output (Figure 5). Specifically, credit growth is highest in Agriculture, Energy and Utility, Transport and Warehousing, and Real Estate itself, all of which are industries with the highest average shares of tangible assets in total assets. In contrast, the additional results show that the information sector, which has one of the lowest shares of tangible fixed assets and whose production function is intensive in *intangible* assets, receives a significantly lower share of credit in response to bank inflows.

In sum, Table 7 documents the role of real estate collateral in firms' access to credit in response to bank flow shocks. The empirical results suggest that banks allocate more credit to firms with more real estate collateral, even after controlling for loan demand. This effect is stronger for riskier banks with higher interbank funding ratios or a greater net foreign asset exposure inside the euro area. The estimation results are robust to using the industry rather than the individual share of tangible assets, to employing the initial level of tangible assets, or to excluding from the sample the period of government stimulus via credit guarantees. This evidence is in strong accord with a transmission mechanism of bank flow shocks to output in which real estate collateral plays a critical role, as established in Section 3.

## 5. Firm and Industry Outcomes

Having established that retrenching banks supplied more credit to firms with more real estate collateral, in this final section, we want to evaluate the role of collateral in determining the differential impact of a bank flow shock on firm- and industry-level outcomes. We focus particularly on employment, investment, total factor productivity (TFP), and borrowing costs. We then also evaluate whether bank flow shocks are associated with capital misallocation.

We measure borrowing cost changes, $\Delta INTEXP$, with the log difference of firm interest expenses as a share of total debt as in Bernile, Bhagwat, and Rau (2017). Employment growth, $\Delta EMPL$, is the rate of growth in the total number of employees. Investment, $\Delta K$, is the change in firm *total* fixed assets as a share of total assets, so as to make sure that the results are not driven by firm size. TFP growth, $\Delta TFP$, is constructed by estimating a production function based on our firm-level data, aggregated at the industry level at the second digit of the NAICS code, following Wooldridge (2009). Specifically, TFP is the residual of a regression of firm-level log real value added on log labor input (the log of the real wage bill), and log capital input (the log of the real book value of total fixed assets), where firm value added and the wage bill are deflated with the two-digit industry price deflators from the OECD STAN database. The capital stock is deflated by the price of investment goods. Some of the firm-level variables take extreme values. We therefore winsorize at the 1% level $\Delta INTEXP$, as well as all variables used for the estimation of TFP before taking logs. Other firm variables are not winsorized, but the results would be essentially unchanged if we did winsorize them.

To evaluate the role of real estate collateral in the transmission of bank flow shocks to firms, we specify the following firm-level reduced-form regression:

$$\Delta \mathrm{Y}_{j,t} = \alpha_j + \alpha_t + \gamma \cdot \Delta \mathrm{Y}_{j,t-1} + \upsilon \cdot (\mathrm{SPREAD}_{t-1} \times \mathrm{TS}_{j,t-1}) + \varepsilon_{j,t}, \qquad (9)$$

where $Y$ denotes alternative firm outcomes. As in the credit regression before, the main independent variable in all specifications is the $GIPS$ spread interacted with the firm-level share of tangible assets. To mitigate endogenity concerns, we continue to rely on the microeconomic nature of the data, including firm and time fixed effects. To address concerns that firms may be on different trend paths, we include the lagged dependent variable in all regressions.

First, we expect bank flow shocks to reduce the borrowing costs of firms with a higher share of real estate collateral since these firms can obtain more credit on possibly better terms (Benmelech, Garmaise, and Moskowitz 2005; Benmelech and Bergman 2008).[32] Second, for consistency with our findings in Section 3, we should also find a positive coefficient on employment growth and investment in the transmission of a bank flow shock through the commercial

[32] Unfortunately, the German credit registry does not include information on the price of credit contracted.

**Table 9**
**Firm and industry outcomes**

| | Firm-level | Firm-level | Firm-level | Firm-level | Industry-level | Industry-level |
|---|---|---|---|---|---|---|
| | (1) $\Delta$INTEXP | (2) $\Delta$EMPL | (3) $\Delta$K | (4) $\Delta$TFP | (5) $\Delta$TFP | (6) SD(TFP) |
| $\text{Spread}_{t-1} \times \text{TS}_{t-1}$ | −0.00021* | 0.009*** | 0.144* | −0.002 | - | - |
| | (0.00) | (0.00) | (0.08) | (0.01) | | |
| $\text{Spread}_{t-1} \times \text{TS}_{\text{Industry},t-1}$ | - | - | - | - | −0.016 | 0.005 |
| | | | | | (0.02) | (0.03) |
| $\text{TS}_{t-1}$ | 0.004** | −0.042 | 1.225*** | 0.386*** | - | - |
| | (0.00) | (0.07) | (0.34) | (0.08) | | |
| $\text{TS}_{\text{Industry},t-1}$ | - | - | - | - | 1.308 | 0.335 |
| | | | | | (0.86) | (1.14) |
| $\text{Y}_{t-1}$ | −0.415*** | −0.484*** | 0.031 | −0.425*** | −0.358** | −0.024 |
| | (0.01) | (0.03) | (0.03) | (0.03) | (0.12) | (0.21) |
| Firm FE | Yes | Yes | Yes | Yes | - | - |
| Time FE | Yes | Yes | Yes | Yes | Yes | Yes |
| Industry FE | – | – | – | – | Yes | Yes |
| Obs. | 20,681 | 41,827 | 50,774 | 15,143 | 64 | 64 |
| $R^2$ | .326 | .364 | .031 | .389 | .340 | .735 |

The regressions in Columns 1–4 are based on annual firm-level data over the period 2009-2014. The dependent variables are: changes in firm borrowing costs ($\Delta INTEXP$), measured as the log difference of firm interest expenses as a share of total debt in column 1; employment growth ($\Delta EMPL$), measured as the rate of growth in the total number of employees in column 2; investment ($\Delta K$), measured as the change in firm *total* fixed assets as a share of total assets in column 3; TFP growth ($\Delta TFP$), constructed by estimating a production function based on our firm-level data aggregated at the second digit of the NAICS code, following Wooldridge (2009), in column 4. The independent variable is the GIPS spread interacted with the firm share of tangible assets. All specifications of columns 1–4 include the respective lagged dependent variable, $LDV$, in addition to firm and time fixed effects. The standard errors are clustered at the industry-year level and are shown in parentheses. Columns 5 and 6 are industry-level regressions in which the dependent variables are industry-level average total factor productivity growth, and the industry-level dispersion in total factor productivity. The main regressor here is the GIPS spread interacted with the industry *average* share of tangible assets. Both specifications include the lagged dependent variable, as well as time and industry fixed effects. Standard errors are clustered at the industry-year level in the first four regressions and at the industry level in the last two regressions. All regressions are implemented with STATA's reghdfe command. *$p$ <.1; **$p$ <.05; ***$p$ <.01.

real estate sector. In contrast, we do not have a firm prior on the impact on TFP of the bank flow shock.

Table 9 reports the results. Column 1 shows that bank flow shocks not only increase the credit supply to high-tangible asset firms, as shown before, but also reduce their borrowing costs, in line with the aggregate evidence reported in Table B1. Moreover, columns 2 and 3 show that bank flow shocks increase employment and investment of high-tangible asset firms, with a statistically significance at the 1% and 10% level, respectively. Column 4 suggests that a higher firm share of tangible assets has a strong positive linear effect on TFP growth. However, there is no differential impact of the bank flow shock on high-tangible asset firms. This is evident from the estimated coefficient on the level of lagged firm TS, which is positive and highly significant, and the insignificant coefficient on the interaction term. In other words, column 4 illustrates that bank flow shocks are not associated with a disproportionate increase or reduction in TFP growth of high-tangible asset firms. This result suggests that there was no capital misallocation during the bank repatriation episode that we study.

To investigate further this hypothesis, we first aggregate our firm-level data at the NAICS2 code industry level as in Doerr (2018). We then regress average industry TFP *growth* rates on the interaction between the GIPS spread and average industry-level shares of tangible assets, controlling for lagged TFP growth in addition to time and industry fixed effects. Column 5 shows that, as in the firm regression in column 4, there is no significant association between the bank flow shock and a disproportionate change in TFP of high-tangible asset industries.

Second, we also regress the industry-level TFP *dispersion* on the interaction between the GIPS spread and the industry-level average share of tangible assets. Following Hsieh and Klenow (2009), the idea here is that, if credit growth leads to capital misallocation, TFP dispersion across firms in the same industry should increase with the bank flow shock, especially in industries with more real estate collateral that obtained a more than proportional share of the declining credit volumes during the 2009–2014 period. Column 6 shows that the estimated coefficient on this interaction term is positive, but clearly statistically insignificant. Even though this last regression is run with very few observations, the finding is consistent with the results of Gopinath et al. (2017), who do not uncover evidence of misallocation in Germany in the run up to the GFC when German capital was flowing abroad through its banking system.[33] In an unreported regression, we also regressed TFP dispersion on the GIPS spread, without interacting the latter with industry-level average tangible asset ratios, finding no association between the GIPS spread and industry-level TFP dispersion.

These results stand in sharp contrast to some other findings in the misallocation literature, specifically focused on housing booms. For example, Doerr (2018) shows that, when property and land prices increase, firms with larger real estate holdings hire, invest, and produce more, but are less productive than firms with smaller real estate holdings in a sample of U.S. public companies. Similarly, Martin, Moral-Benito, and Schmitz (2018) find that Spanish banks more exposed to a real estate bubble initially lend relatively more to housing firms than nonhousing firms. However, as the bubble persists, this composition effect disappears because housing credit repayments raise banks' net worth, supporting the credit access of all firms. In contrast, we show that bank retrenchment, while causing higher property prices with varying intensities across cities, is not associated with lower productivity growth at the firm or industry-level, or increased TFP dispersion. In the absence of a credit boom, the bank flow shock does not appear to be associated with lower TFP growth and capital misallocation. In fact, as we have noted before, aggregate

[33] The result is unchanged if we replace TFP dispersion with the dispersion of the marginal product of capital. This is not surprising because, with the constant factor shares and a Cobb-Douglas production function, TFP dispersion and the dispersion of the marginal product of capital are proportional to each other (see Gopinath et al. (2017)).

credit declined in real terms during the period we consider (see, for instance, panel D in Figure 1).

In sum, this last set of results accords well with the findings of Section 3 based on city-level data. The estimated impact of a bank flow shock on firm and industry-level outcomes provides additional evidence that real estate collateral plays a significant role. Real estate collateral not only seems critical for the differential access of firms and sectors to the increased credit supply triggered by bank inflows but also is associated with increased hiring and investment, thus contributing to higher levels of local economic activity, without evidence of misallocation.

## 6. Conclusions

This paper studies the role of real estate markets in the transmission of capital flow shocks to city business cycles in Germany by using a new matched city-level and bank-firm-level data set. Germany is an interesting laboratory because, during and after the global financial crisis, it experienced a sizable bank retrenchment episode with a real estate boom, without a credit boom, well captured by the sovereign bond spread of Southern European countries (the so-called "GIPS spread").

To identify the differential impact of bank flow shocks on output growth across German cities, we exploit the quasi-random geographic variation in a city-level measure of real estate market tightness or exposure. This measure is the product of the city share of refugees in total refugees, which is determined by long-standing government rules and regulations, and a measure of nondevelopable area, which is determined by geography and land use regulations.

We find that the output growth impact of a bank flow shock, as measured by the sovereign bond spread of Southern European countries, is more significant in cities that are more exposed to local real estate markets tightness. Our estimates imply that the cities most exposed to real estate market tightness, on average, grew 31 bps more per year than the least exposed ones during 2009–2014, or 1.9 percentage points more cumulatively. Moreover, the differential response of commercial property prices across cities to the bank flow shock accounts for the bulk of this growth differential.

The transmission mechanism that we uncover works through a collateral channel on the firm side, with commercial real estate playing a critical role. We document the importance of real estate collateral for firm credit access and bank behavior by showing that German banks repatriated gross foreign assets from the rest of Europe after the GFC and lent disproportionately more to domestic firms and sectors with more tangible fixed assets. We also show that firms with more tangible assets hired and invested more in response to the bank flow shock. Consistent with the extant literature on Germany, but differently from studies of housing booms with credit booms, we do not find evidence of capital misallocation associated with the transmission of bank flow shocks across German cities.

While we find that property prices in the residential sector have no role in the transmission of bank flow shocks to city output growth during the retrenchment episode that we study, the paper is silent on the causes of the German residential boom and its economic effects. Both the creditless nature of the boom with falling household leverage and the reduced-form evidence on the predictive power of the city share of young people that we report are consistent with a cash-based buy-to-let hypothesis. Exploring drivers of residential house prices in the context of a portfolio rebalancing framework like in Flavin and Yamashita (2002) and exploring its impact on household consumption is an interesting area of future research.

## A. Variable Definitions, Data Sources, and Summary Statistics

**Table A1**
**Definitions and sources of all variables**

| Variable | Definition | Unit | Source |
|---|---|---|---|
| $RREP$ | City c's residential real estate price index, deflated by state-level CPI | 2008=100 | Bulwiengesa |
| $CREP$ | City c's commercial real estate price index, deflated by state-level CPI | 2008=100 | Bulwiengesa |
| Nondevelopable area | City c's nondevelopable area (open space) relative to the total area in 2008 | % | IOER Monitor |
| Refugees | 2008 share of refugees allocated by the government to city *c* relative to all refugees in 2008 | % | Fed. Stat. Off. |
| Exposure | City c's product of the share of refugees and the share of nondevelopable area | % | Own calculation |
| Share 25–30 | City c's share of people aged 25–30 in total population in 2008 | % | BBSR Bonn (INKAR)[a] |
| $\Delta GDP$ | City c's growth in GDP per capita, deflated by state-level CPI | % | BBSR Bonn (INKAR) |
| Pop. dens. | City c's number of inhabitants per square kilometer in 2008 | - | BBSR Bonn (INKAR) |
| Pop. | City c's population in total German population in 2008 | % | GENESIS |
| GIPS spread | The average 10-year government bond spread of Greece, Italy, Portugal, and Spain relative to Germany | % | FRED |
| Bund | The German 10-year government bond yield | % | FRED |
| EONIA | The European interbank market rate | % | FRED |
| ECB rate | The ECB's monetary policy rate | % | FRED |
| VIX | The CBOE volatility index | % | FRED |
| Net bank inflows outside the eurozone | Change in net foreign liabilities outside the eurozone of German BIS reporting banks | % of GDP | BIS |

(*Continued*)

**Table A1**
**(Continued)**

| | | | |
|---|---|---|---|
| Net bank inflows inside the eurozone | Change in net foreign liabilities inside the eurozone of German BIS reporting banks | % of GDP | BIS |
| Gross bank inflows inside the eurozone | Change in gross foreign liabilities inside the eurozone of German BIS reporting banks | % of GDP | BIS |
| Gross bank outflows inside the eurozone | Change in gross foreign assets inside the eurozone of German BIS reporting banks | % of GDP | BIS |
| Lending-deposit spread | Lending rate is the interest rate charged by banks on short- and medium-term loans to the private sector<br>Deposit interest rate is the interest rate offered by commercial banks on 3-month deposits | % | IMF IFS |
| TS | Firm j's tangible fixed assets (Bureau van Dijk code TFAS) as a share of total assets (TOAS) | % | Amadeus |
| $TS_{Industry}$ | The arithmetic mean of TS by two-digit NAICS code | % | Amadeus |
| $TS_{2008} * CREP$ | Firm j's 2008 tangible over total assets, inflated by the city-level real commercial real estate price index | % | Own calculation |
| EQ | Firm j's capital-to-asset ratio (CAPI/TOAS) | % | Amadeus |
| ROA | Firm j's return on assets (EBIT/TOAS) | % | Amadeus |
| TA | Firm j's logarithm of total assets (TOAS) | ln(Euro) | Amadeus |
| $\Delta INTEXP$ | Change in firm j's logarithm of interest expenses over total debt (INTE/LOAN) | % | Amadeus |
| $\Delta EMPL$ | Change in firm j's logarithm of the number of employees (EMPL) | % | Amadeus |
| $\Delta K$ | Change in firm j's total fixed assets (FIAS) scaled by total assets (TOAS) | % | Amadeus |
| $\Delta TFP$ | Change in firm j's logarithm of TFP computed by following Wooldridge (2009) | % | Amadeus |
| $\Delta L_{i,j,t}$ | Log-difference of the stock of loans of bank $i$ to firm $j$ in quarter-year $t$ | % | Credit Register |
| Bank share of gross foreign assets | Bank i's gross foreign assets over total assets | % | Bundesbank |
| Net foreign assets | Bank i's net foreign assets vis-à-vis the euro area over total assets | [0,1] | Bundesbank |
| Interbank | Bank i's interbank funding-to-deposits ratio | [0,1] | Bundesbank |
| Capital ratio | Bank i's regulatory capital-to-asset ratio | % | Bundesbank |
| Size | Bank i's logarithm of total assets | ln(Euro) | Bundesbank |
| Liquidity | Bank i's liquid assets over total assets | % | Bundesbank |
| ROA | Bank i's return on risk-weighted assets | % | Bundesbank |
| NPL | Bank i's nonperforming over total loans | % | Bundesbank |
| Loans | Bank i's loans over total assets | % | Bundesbank |

[a] All data from BBSR Bonn are subject to Data license Germany – BBSR Bonn – Version 2.0

**Table A2**
**Summary statistics for all variables used in the empirical analysis (% unless otherwise noted)**

| Variable | Observations | Mean | Median | SD | 25th | 75th |
|---|---|---|---|---|---|---|
| $RREP$ | 510 | 104.66 | 101.55 | 9.95 | 97.84 | 109.70 |
| $CREP$ | 510 | 103.37 | 102.64 | 9.71 | 98.01 | 108.01 |
| Nondevelopable area | 510 | 59.44 | 58.60 | 13.55 | 49.70 | 70.90 |
| Refugees | 510 | 0.43 | 0.27 | 0.54 | 0.14 | 0.43 |
| Exposure | 510 | 21.58 | 16.54 | 19.54 | 8.35 | 27.23 |
| Share 25–30 | 510 | 7.22 | 7.10 | 1.29 | 6.30 | 8.20 |
| $\Delta GDP$ | 510 | 0.57 | 0.63 | 4.40 | −1.53 | 2.66 |
| Pop. dens. | 510 | 1487.08 | 1357.00 | 745.37 | 951.00 | 1843.60 |
| Pop. | 510 | 0.27 | 0.20 | 0.25 | 0.13 | 0.31 |
| GIPS spread | 60 | 2.04 | 0.46 | 2.84 | 0.19 | 3.22 |
| Bund | 60 | 3.36 | 3.63 | 1.26 | 2.42 | 4.30 |
| EONIA | 60 | 2.08 | 2.07 | 1.57 | 0.36 | 3.34 |
| ECB rate | 60 | 2.2 | 2.0 | 1.39 | 1.0 | 3.38 |
| VIX | 60 | 20.95 | 19.34 | 8.39 | 14.90 | 24.99 |
| Net bank inflows outside the eurozone | 60 | −2.08 | −4.28 | 12.31 | −8.52 | 1.96 |
| Net bank inflows inside the eurozone | 60 | −2.32 | −1.17 | 6.05 | −5.33 | 1.06 |
| Gross bank inflows inside the eurozone | 60 | 0.13 | 0.78 | 4.65 | −1.96 | 2.64 |
| Gross bank outflows inside the eurozone | 60 | 2.46 | 2.03 | 6.69 | −1.39 | 4.97 |
| Lending-deposit spread | 48 | 2.15 | 2.05 | 0.70 | 1.61 | 2.50 |
| TS | 72,290 | 38.04 | 29.63 | 32.03 | 8.74 | 65.52 |
| $\text{TS}_{\text{Industry}}$ | 90,483 | 36.41 | 27.09 | 18.53 | 22.35 | 54.36 |
| $\text{TS}_{2008} * \text{CREP}$ | 20,468 | 35.31 | 22.12 | 34.84 | 5.45 | 61.48 |
| EQ | 73,948 | 26.36 | 23.85 | 27.77 | 8.56 | 42.76 |
| ROA | 42,275 | 5.48 | 4.51 | 12.29 | 0.75 | 10.30 |
| TA | 75,076 | 21.09 | 23.06 | 4.68 | 15.62 | 24.02 |
| $\Delta INTEXP$ | 27,266 | −0.02 | −0.04 | 1.07 | −0.36 | 0.29 |
| $\Delta EMPL$ | 65,776 | 4.31 | 0.00 | 54.85 | −0.80 | 6.24 |
| $\Delta K$ | 62,043 | 33.44 | 0.04 | 476.42 | −0.24 | 0.65 |
| $\Delta TFP$ | 28,813 | −0.52 | 0.04 | 35.51 | −9.15 | 9.49 |
| $\Delta L_{i,j,t}$ | 723,296 | −2.84 | −0.83 | 63.54 | −7.34 | 2.04 |
| Bank share of gross foreign assets | 89,651 | 58.07 | 59.82 | 15.74 | 50.23 | 67.78 |
| Net foreign assets | 29,606 | 0.0009 | −0.00006 | 0.07 | −0.001 | 0.002 |
| Interbank | 27,338 | 0.0003 | 0.00 | 0.009 | 0.00 | 0.00 |
| Capital ratio | 28,769 | 18.90 | 17.03 | 14.40 | 14.59 | 20.48 |
| Size | 29,615 | 20.76 | 20.59 | 1.32 | 19.87 | 21.41 |
| Liquidity | 29,615 | 20.91 | 17.06 | 15.33 | 12.62 | 23.32 |
| ROA | 28,583 | 2.12 | 1.99 | 3.46 | 1.56 | 2.49 |
| NPL | 27,344 | 3.90 | 3.30 | 3.85 | 2.05 | 4.89 |
| Loans | 29,610 | 58.52 | 60.23 | 15.80 | 51.01 | 68.17 |

The table reports summary statistics for all variables employed in our analysis after any outlier correction. See Table A1 for the variable definitions and data sources.

**Table A3**
**List of cities, refugee share, land shares, and exposures in 2008 (%)**

| | | | | |
|---|---|---|---|---|
| Aschaffenburg (0.2, 64.3, 11.3) | Duisburg (0.8, 42.9, 35.7) | Heilbronn (0.4, 68.0, 28.3) | Magdeburg (0.3, 62.1, 20.3) | Remscheid (0.1, 58.6, 7.1) |
| Augsburg (0.6, 58.5, 34.3) | Düsseldorf (1.1, 43.2, 48.5) | Herne (0.2, 25.6, 6.0) | Mainz (0.4, 56.2, 23.5) | Rosenheim (0.1, 61.3, 4.2) |
| Bamberg (0.1, 53.7, 4.4) | Eisenach (0.0, 84.6, 3.0) | Ingolstadt (0.1, 67.2, 7.3) | Mannheim (0.5, 47.3, 25.5) | Rostock (0.3, 70.4, 19.8) |
| Bayreuth (0.2, 63.4, 10.0) | Erfurt (0.3, 75.8, 22.3) | Jena (0.1, 76.3, 7.1) | Mönchengladbach (0.4, 53.7, 19.5) | Salzgitter (0.2, 75.4, 12.1) |
| Bielefeld (0.7, 59.7, 41.0) | Erlangen (0.1, 63.9, 8.7) | Kaiserslautern (0.1, 75.4, 6.4) | Müheim an der Ruhr (0.4, 44.9, 16.5) | Schweinfurt (0.1, 49.7, 6.3) |
| Bochum (0.6, 34.0, 19.3) | Essen (1.4, 37.5, 52.7) | Karlsruhe (0.5, 58.0, 31.5) | München (4.0, 27.1, 107.4) | Schwerin (0.2, 78.5, 19.6) |
| Bonn (0.7, 50.6, 36.4) | Flensburg (0.1, 51.2, 4.3) | Kassel (0.6, 47.2, 27.5) | Münster (0.4, 71.6, 28.1) | Solingen (0.2, 52.5, 9.7) |
| Bottrop (0.1, 62.4, 8.3) | Frankfurt (Oder) (0.1, 79.4, 5.8) | Kempten (Allgäu) (0.0, 70.4, 3.1) | Neumünster (0.2, 52.5, 8.6) | Stuttgart (1.3, 50.0, 66.3) |
| Brandenburg an der Havel (0.1, 81.0, 7.9) | Frankfurt am Main (1.7, 47.5, 79.4) | Kiel (0.4, 45.5, 18.9) | Nürnberg (1.2, 42.9, 53.0) | Suhl (0.0, 81.2, 3.1) |
| Braunschweig (0.2, 58.2, 13.7) | Freiburg im Breisgau (0.4, 71.7, 30.0) | Koblenz (0.3, 68.2, 18.1) | Oberhausen (0.4, 32.9, 11.9) | Trier (0.2, 73.5, 13.7) |
| Bremerhaven (0.2, 54.7, 9.2) | Fürth (0.2, 56.9, 9.2) | Krefeld (0.3, 48.1, 12.9) | Offenbach am Main (0.4, 55.6, 20.8) | Ulm (0.3, 70.9, 19.4) |
| Chemnitz (0.3, 59.7, 19.4) | Gelsenkirchen (0.5, 34.2, 17.1) | Köln (1.8, 44.1, 78.5) | Oldenburg (Oldenburg) (0.4, 42.5, 16.0) | Weimar (0.0, 76.2, 3.3) |
| Cottbus (0.1, 71.0, 7.9) | Gera (0.1, 77.6, 4.3) | Landshut (0.1, 67.5, 5.7) | Osnabrück (0.2, 53.6, 12.1) | Wiesbaden (0.5, 68.3, 34.2) |
| Darmstadt (1.0, 68.0, 69.7) | Hagen (0.4, 66.2, 23.8) | Leipzig (0.6, 51.3, 31.0) | Passau (0.0, 72.3, 3.3) | Wilhelmshaven (0.1, 63.1, 6.7) |
| Dessau-Roßlau (0.1, 84.1, 6.8) | Halle (Saale) (0.4, 58.4, 24.8) | Leverkusen (0.3, 46.0, 11.5) | Pforzheim (0.4, 72.2, 27.2) | Wolfsburg (0.2, 74.6, 13.7) |
| Dortmund (1.3, 44.7, 56.3) | Hamm (0.2, 70.9, 14.4) | Ludwigshafen am Rhein (0.4, 47.0, 17.3) | Potsdam (0.2, 72.1, 16.3) | Wuppertal (1.0, 52.9, 50.6) |
| Dresden (0.4, 58.6, 24.4) | Heidelberg (0.2, 70.9, 16.1) | Lübeck (0.3, 66.9, 17.0) | Regensburg (0.2, 47.7, 8.8) | Würzburg (0.3, 56.3, 17.2) |

The table lists the 85 urban areas included in our empirical analysis. The numbers in parentheses are the 2008 values of the share of refugees, the nondevelopable area, and our exposure measure (the product of the two), respectively, rounded to the first decimal digit. See Table A2 for summary statistics of these variables.

**Table A4**
**Matched versus unmatched firms (%)**

| Variable | Mean (matched) | Median (matched) | SD (matched) | Mean (unmatched) | Median (unmatched) | SD (unmatched) | $t$-test of difference |
|---|---|---|---|---|---|---|---|
| TA | 21.1 | 23.1 | 4.7 | 6.8 | 0.2 | 420.5 | $t=-69.3$ |
| ROA | 5.5 | 4.5 | 12.3 | 7.2 | 4.8 | 19.1 | $t=19.3$ |
| EQ | 26.4 | 23.9 | 27.8 | 39.2 | 36.9 | 41.4 | $t=38.3$ |
| TS | 38.0 | 29.6 | 32.0 | 23.3 | 10.8 | 34.7 | $t=-1,610.7$ |

The table shows the means, medians, and standard deviations for selected firm-level variables for the matched sample of firms and nonmatched firms. The variables are the logarithm of total assets, the return on assets, the equity-to-asset ratio, and the share of tangible fixed over total assets.

## B. Measuring Capital Flows, Cross-Border Bank Flows, and the GIPS Spread

In this appendix, we document formally that the GIPS Spread is a good proxy variable for cross-border bank flows as an auxiliary step in our empirical analysis. To quantify the relevance of the GIPS spread as predictor of bank flows, we run a battery of regressions for alternative bank flow measures on the GIPS spread. The frequency is quarterly and the sample period is 2000:Q1–2014:Q4 to make sure that the spread can capture both phases of the boom-bust cycle. The estimated equation is specified as follows:

$$\mathrm{BF}_t = \gamma \cdot \mathrm{Spread}_t + \varepsilon_t, \tag{B1}$$

where $\mathrm{BF}_t$ represents alternative measures of bank flows, and "$\mathrm{Spread}_t$" denotes the GIPS spread. We distinguish between *net* flows from *outside* and *inside* the euro area. We then break down net flows from the rest of the euro area into *gross* inflows and outflows. Following Larrain, and Stumpner (2017), we also examine the impact of the GIPS spread on the domestic lending-deposit interest rate spread. If bank flows increase domestic credit supply, we should observe a negative effect on the domestic lending-deposit spread. Finally, we use our *bank-level* data to evaluate the predictive ability of the GIPS spread for *individual* banks' gross foreign assets as a share of total assets, controlling for bank fixed effects. The last regression is important as concerns regarding reverse causation from bank flows to the GIPS spread are mitigated by the use of bank-level data.

Table B1 reports the results. Columns 1 and 2 show that a higher GIPS spread is positively associated with net bank flows into Germany from both outside and inside the euro area. The relation, however, is statistically significant only for net bank flows originating from the rest of the euro area. The results in Columns 3 and 4 also illustrate that net bank flows are driven by lower gross bank outflows, rather than higher gross bank inflows. These regressions, therefore, taken together, suggest that a GIPS spread increase is associated with a repatriation of bank foreign assets from the rest of the euro area, which Forbes and Warnock (2012) call "retrenchment" of capital.

The evidence of retrenchment is further corroborated by column 6, which shows that a higher GIPS spread is associated with a smaller share of gross foreign assets in total bank assets at the level of individual banks. This last regression suggests that, in economic terms, a 100-bp increase in the GIPS spread reduces banks' share of foreign assets in total assets by almost 25 bps. Put differently, this estimate implies that, during the peak of the European crisis, the German banking system shifted lending from foreign to domestic borrowers amounting to about 1.6% of its aggregate balance sheet, or 1.9% of GDP.[34]

[34] The GIPS spread averaged 6.5% during the acute phase of the European crisis, from 2010:Q1 to 2012:Q3, compared to a value close to zero right before the GFC. Hence, the impact of the crisis is quantified as 6.5%*0.246=1.60%. According to FRED data, total assets held by deposit money banks compared to GDP were approximately 120% in 2009. As a result, the estimated shift in banking assets is 1.60%*1.2=1.92% of GDP.

**Table B1**
**The GIPS Spread and Bank Flows**

| | Country-Level | Country-Level | Country-Level | Country-Level | Country-Level | Bank-Level |
|---|---|---|---|---|---|---|
| | (1) Net Bank Inflows Outside Eurozone | (2) Net Bank Inflows Inside Eurozone | (3) Gross Bank Inflows Inside Eurozone | (4) Gross Bank Outflows Inside Eurozone | (5) Lending-Deposit Spread | (6) Bank Share of Gross Foreign Assets |
| GIPS Spread$_t$ | 0.526 | 0.639*** | −0.116 | −0.755*** | −0.115** | −0.246*** |
| | (0.316) | (0.115) | (0.075) | (0.130) | (0.044) | (0.030) |
| Bank FE | — | — | — | — | — | Yes |
| Obs | 60 | 60 | 60 | 60 | 48 | 89,651 |
| $R^2$ | .040 | .247 | .014 | .283 | .247 | .844 |

*Note*. All regressions are based on quarterly data over the period 2000:Q1-2014:Q4, except for the regression in Column (5) for which the data are not available before 2003. The dependent variable in Columns (1) and (2) is *net* bank flows into the German banking system from the rest of the world outside the euro area and from the rest of the euro area, respectively. In Columns (3) and (4), the dependent variable is *gross* inflows and outflows from the rest of the euro area, respectively. In Column (5), the dependent variable is the difference between the domestic lending and deposit interest rate. In Column (6), the dependent variable is the share of individual banks' gross foreign assets over total assets. The regression in Column (6) includes individual bank fixed effects. See the Data Appendixes for variable definitions and data sources. Newey-West standard errors with 3 lags (columns (1)-(5)) and bank-level clustered standard errors (column (6)) are shown in parentheses. *, ** and *** indicate statistical significance at the 10%, 5%, and 1% level, respectively.

**Table B2**
**The GIPS Spread and Bank Flows: Robustness**

| | Country-Level | Country-Level | Country-Level | Country-Level | Country-Level | Bank-Level |
|---|---|---|---|---|---|---|
| | (1) Net Bank Inflows Outside Eurozone | (2) Net Bank Inflows Inside Eurozone | (3) Gross Bank Inflows Inside Eurozone | (4) Gross Bank Outflows Inside Eurozone | (5) Lending-Deposit Spread | (6) Bank Share of Gross Foreign Assets |
| Spread$_t$ | 0.238 | 0.532*** | −0.009 | −0.541** | −0.133* | −0.059*** |
| | (0.448) | (0.180) | (0.156) | (0.237) | (0.070) | (0.020) |
| Bank FE | — | — | — | — | — | Yes |
| Obs | 32 | 32 | 32 | 32 | 32 | 46,125 |
| $R^2$ | .007 | .209 | .000 | .196 | .260 | .920 |

*Note*. The regressions in this table are the same as in Table B1 in the paper. However, they are based on quarterly data from 2007:Q1 to 2014:Q4. All variables are defined as in the Data Appendix. Newey-West standard errors with 3 lags (columns (1)-(5)) and bank-level clustered standard errors (column (6)) are shown in parentheses. *, ** and *** indicate statistical significance at the 10%, 5%, and 1% levels, respectively.

As shown in column 5, a higher GIPS spread is also associated with a lower domestic lending-deposit spread, suggesting that the German bank retrenchment episode we consider is associated with looser domestic financial conditions and an aggregate increased supply of credit. Moreover, as we will report and discuss in Section 5, a higher GIPS spread leads to lower debt service costs at the firm level. This transmission, therefore, is in line with the hypothesis that a bank flow shock can loosen domestic financial conditions and increase the domestic credit supply.

Table B2 shows that these results are similar if we restrict the sample to 2007–2014. In unreported regressions, we also obtain essentially the same results as in column 6 above by using net, rather than gross, individual bank foreign assets. Finally, the results are also unchanged if we include in the construction of the GIPS spread Ireland, or exclude Greece, as the sovereign bond spreads are highly correlated in crisis times.

In summary, these findings confirm that German banks experienced a sizable net inflow of capital from the rest of the euro area since the GFC, driven by banks' repatriation of foreign assets, consistent with available evidence on the behavior of Northern European banks before and after the GFC discussed in the main text. The evidence also shows that the GIPS spread is a good predictor of these bank flows.

## C. Refugee Allocation as an Instrument for Real Estate Demand

This appendix provides institutional details about the German framework for the city allocation of refugees underpinning its validity as an instrument for demand of real estate services. We focus particularly on refugee mobility across cities and the possible impact on labor supply.[35]

### C.1 Refugee Mobility across Cities

Refugees assigned to a given city, in the non-city-states in our sample, cannot easily relocate. Upon arrival, asylum seekers must apply for status at the *assigned* federal office for immigration and refugees (BAMF). After registration and asylum application, the refugees are distributed to the municipalities within the state according to the quota systems summarized in Table 2. A first-round decision on status is supposed to be taken within 6 months. While an application is pending, asylum seekers are required to stay at the initial accommodation center and cannot leave without permission. Only if and when BAMF grants status, refugees can relocate. However, as many applications are initially rejected, and most asylum seekers appeal in the courts, which typically takes a year or more, refugees usually remain confined to their initial city assignment for much longer than the minimum time necessary to obtain status. Of course, refugees could relocate if they find illegal work in another city before they reach status. However, illegal work is not widespread in Germany, also because of the language barrier. As a consequence, the number of refugees able to leave the initial city allocation is also bound to be minimal.

Once refugees obtain status, if they find a job and can self-sustain, they can relocate. As argued below though, access to the labor market is very limited in the short-to-medium term. Similarly, their ability to compete for rentals with domestic residents is constrained by the low wages relative to the natives', even if they are legally employed. As a result, the number of refugees that are able to relocate finding legal employment and self-sustaining after reaching status is also very small.

After status is reached, if a refugee cannot find a job and is not financially self-sufficient, the government continues to determine where subsidized shelter is provided. Subsidized shelter is provided in secondary accommodation centers or in independent accommodations, depending on family circumstances. On average, in the non-city-states in our sample, about 50% of the refugees are housed in accommodation centers, which is also more cost effective (Table 2). While on housing welfare, until 2015, refugees could request relocation. However, the administrative burden of the process and again the language barrier made this possibility not likely.

In sum, although there are no data on the actual degree of internal mobility of refugees, it seems highly unlikely to be pervasive. We also note that our sample excludes the city-states, Berlin, Hamburg, and Bremen, about which press and media reports of deviations from the initial state allocation rules often refer to. Indeed, the panel correlation between the city share of state population in 2008, the critical criterion of the state allocation rules, and the annual de facto city share of state refugees during the period 2009–2014 is 0.91, pointing to a very high persistence in the allocation across cities and suggesting that deviations from the initial allocations are minimal.

[35] The main sources of information on the institutional details of German refugee policy on which we relied upon are Müller (2013), Baier and Siegert (2018) and Nam and Steinhoff (2018).

### C.2 Refugees and Labor Supply

In Germany, refugees are quite unlikely to have any significant impact on the labor market at business cycle frequency, even after they reach status. In the short-to-medium term, this is because of formal restrictions to refugees' entry in the labor market even after they obtain a work permit or reach status. The main reason is a legal requirement of working knowledge of the German language for formal employment in most occupations that was in place until changes were introduced in 2015 and 2016, to facilitate integration. Before 2015, the law also entailed preferences toward German and European applicants, as well as other restrictions on work permits for refugees who did not complete vocational training.[36] For example, Brell, Dustmann, and Preston (2020) estimate that the rate of employment of refugees is essentially zero within 1 year, less than 15% within 2 years after they receive their work permit, and still below 25% six years after obtaining the work permit. Moreover, most of these jobs, are temporary and low-skill according to OECD (2017a), fetching wages that pay less than 50% of the native rates, even after 5 years from obtaining the work permit (Brell, Dustmann, and Preston 2020). Even in the long run, refugees' access to the labor market is limited in Germany, chiefly because of the language barrier. For example, Katz, Noring, and Garrelts (2016) report that, on average, it may take up to 15 years of residence to obtain a rate of employment comparable to that of natives.

## D. A Simple Model of City Real Estate Markets

In this Appendix, we set up a simple model of a city's real estate markets with an exogenously varying share of individuals that need shelter (the refugees) to rationalize the working of the refugee share as an instrument for real estate prices as we use it in our empirical analysis. In particular, we show that, while a higher share of refugees unambiguously leads to higher *commercial* real estate prices, the impact on *residential* real estate prices depends on parameter values. It is therefore plausible to find empirically that the instrument is relevant for the commercial sector, while it is weak in the residential sector, as in our first-stage results. We also use the model to evaluate the magnitude of the estimated elasticity of real estate prices to the refugee share.

### D.1 The Model

We model commercial and residential housing following Liu, Wang, and Zha (2013). We consider a collection of $N$ islands or city economies that are identical except for the city share of refugees. There are two types of agents in the local economy. One is the resident or native and the other is the refugee. Residents of a given city $c$ are a continuum of mass $p_c$ that can be normalized to 1 without loss of generality, while refugees are a continuum of mass $\pi_c$, with $\pi_c < p_c$. Thus, the total city population is $p_c + \pi_c$ while the normalized size of total population is given by $1+\frac{\pi_c}{p_c}$.

We omit the city subscript $c$ for simplicity and denote the ratio of city refugees to city population by $\kappa$. By construction, we have

$$\kappa = \frac{\pi_c}{p_c} = \frac{\frac{\pi_c}{\sum \pi_c} \times \sum \pi_c}{p_c}, \tag{D1}$$

where $\sum \pi_c$ is the total number of refugees. It is now easy to see that there is a one-to-one relationship between $\kappa$ and the city share of refugees as defined in our empirical analysis, $\frac{\pi_c}{\sum \pi_c}$, conditional on the total number of refugees, $\sum \pi_c$, and the city population, $p_c$. In our empirical framework, the city share of refugees, $\frac{\pi_c}{\sum \pi_c}$, varies across cities but is held constant at the 2008 value. Therefore, in the model below, the comparative statistic analysis with respect to $\kappa$, conditional on city-level population $p_c$, captures the effects of the variation across cities in the share of refugees.

[36] During our sample period, refugees were not allowed to work during the first 3 months after arrival. Between months 4 and 15, they were allowed to work only if the Federal Employment Agency agreed that no other German was equally suitable for the same position and that the wage offered was comparable to the market rate. Between month 16 and the end of the third year, they were allowed to work only if their wage is deemed market comparable. Starting from the fourth year, they could work without restrictions.

**D.1.1 Local Residents.** Consider a representative household-firm-entrepreneur with preferences

$$U_0 = \sum_{t=0}^{\infty} \beta^t \left[\log(c_t) + \omega \log\left(h_t^R\right)\right],$$

where $\beta$ is the discount rate, $c_t$ is consumption, $h_t^R$ is a flow of residential housing services, and $\omega$ is a parameter that governs the share of residential housing expenditure in total expenditure. Output is produced with the following simple technology

$$y_t = A_t \left(h_t^C\right)^{\alpha},$$

where $A_t$ is the aggregate productivity level, $h_t^C$ is a flow of commercial real estate services, and $\alpha$ is a parameter that governs the share of commercial housing income in total income.

Natives purchase residential and commercial housing services at the market prices $p_t^R$ and $p_t^C$, respectively. Their budget constraint is

$$c_t + p_t^R \left(h_{t+1}^R - h_t^R\right) + p_t^C \left(h_{t+1}^C - h_t^C\right) + \frac{b_{t+1}}{R} = y_t + b_t - T_t,$$

where $b_t$ is a one-period bond issued by the household-firm-entrepreneurs with an exogenous national interest rate $R = 1 + r$, and $T_t$ is a lump-sum transfer *from* domestic residents to refugees.

Borrowing is collateralized as follows:

$$-\frac{b_{t+1}}{R} \leq \phi y_t + \phi^C p_t^C h_t^C + \phi^R p_t^R h_t^R,$$

where $\left\{\phi, \phi^R, \phi^C\right\}$ are parameters that govern the tightness of the constraint on different types of collateral. The constraint captures the idea that agents can borrow to purchase residential services, commercial services, or to obtain working capital. We also assume that borrowers are impatient, that is, $\beta R < 1$, so that the borrowing constraint is always binding in the steady state.

The residents maximize utility by choosing $\{c_t, b_{t+1}, h_{t+1}^R, h_{t+1}^C\}$ subject to the budget and collateral constraints. The optimality conditions are, respectively, given by

$$\lambda_t = \frac{1}{c_t}$$

$$\lambda_t = \mu_t + \beta R \lambda_{t+1}$$

$$\lambda_t p_t^R = \beta \left( \frac{\omega}{h_{t+1}^R} + \lambda_{t+1} p_{t+1}^R + \phi^R \mu_{t+1} p_{t+1}^R \right)$$

$$\lambda_t p_t^C = \beta \lambda_{t+1} \left( \alpha A_{t+1} \left(h_{t+1}^C\right)^{\alpha-1} + p_{t+1}^C \right) + \beta \phi \mu_{t+1} \alpha A_{t+1} \left(h_{t+1}^C\right)^{\alpha-1} + \beta \phi^C \mu_{t+1} p_{t+1}^C.$$

where $\lambda_t$ and $\mu_t$ are the Lagrangian multipliers for the budget and collateral constraints.

**D.1.2 Refugees.** We assume that refugees are hand-to-mouth agents and, for subsistence purposes, receive a fixed per capita lump-sum transfer from the government (i.e., the residents), $\tau$. Thus, $T_t = \kappa \tau$. Furthermore, we assume that a fraction $\gamma < 1$ consumes a fixed amount of residential housing services, $\tilde{h}^R$, while a fraction $1 - \gamma < 1$ must consume commercial real estate services, $\tilde{h}^C$. For simplicity, we assume that both types of real estate services are paid directly by the government without entering market transactions. These assumptions capture the critical features of the German policy framework toward asylum seekers and refugees that we discussed in the paper and the previous appendix.

Note here that each refugee receives a cash allowance to purchase a subsistence level of consumption goods, including food and clothing. The government also provides shelter, either in kind of accommodation centers and other communal facilities or through vouchers and reimbursements for renting in the market. Since in the model both kinds of refugees are hand-to-mouth agents, we abstract from the choice between consumption and housing services of the refugees that rent in the market. Rather, their housing and nonhousing consumption expenditures are fixed and will be calibrated to the data.

**D.1.3 Real Estate Supply.** For simplicity, we assume that the total supply of residential ($H^R$) and commercial ($H^C$) housing services is fixed. For the same reason, we also assume that the two markets are segmented by zoning and land use regulations. These assumptions imply that the supply elasticity is zero in both sectors. Without loss of generality, we will normalize the total supply of residential and commercial real estate services to one and use the data to calibrate the sector share. Thus, the market clearing conditions for the real estate markets are

$$h_t^R + \kappa\gamma\tilde{h}^R = H^R$$

$$h_t^C + \kappa(1-\gamma)\tilde{h}^C = H^C.$$

These expressions show that, other things equal, in a city with a larger share of refugees, that is, the higher $\kappa$, there is a smaller supply of commercial and residential real estate services available for residents, with elasticities $-\gamma\tilde{h}^R$ and $-(1-\gamma)\tilde{h}^C$, respectively.[37]

**D.1.4 Steady state.** In the steady-state equilibrium, the following holds:

$$h^R = H^R - \kappa\gamma\tilde{h}^R$$

$$h^C = H^C - \kappa(1-\gamma)\tilde{h}^C$$

$$y = A\left[H^C - \kappa(1-\gamma)\tilde{h}^C\right]^\alpha$$

$$c = \frac{\Gamma_1 A\left[H^C - \kappa(1-\gamma)\tilde{h}^C\right]^\alpha - \kappa\tau}{1 + r\phi^R \frac{\beta\omega}{1-\beta-\beta\phi^R(1-\beta R)}}$$

$$p^C = \frac{\beta[1+\phi(1-\beta R)]}{1-\beta-\beta\phi^C(1-\beta R)}\alpha A\left[H^C - \kappa(1-\gamma)\tilde{h}^C\right]^{\alpha-1}$$

$$p^R = \frac{\beta\omega}{1-\beta-\beta\phi^R(1-\beta R)}\frac{c}{h^R}$$

$$= \frac{\beta\omega}{1-\beta-\beta\phi^R(1-\beta R)+r\phi^R\beta\omega}\frac{\Gamma_1 A\left[H^C - \kappa(1-\gamma)\tilde{h}^C\right]^\alpha - \kappa\tau}{H^R - \kappa\gamma\tilde{h}^R},$$

where $\Gamma_1 = \left(1 - \phi r - r\phi^C\alpha\beta\frac{1+\phi(1-\beta R)}{1-\beta-\beta\phi^C(1-\beta R)}\right)$.

We are interested in the differential impact on the local economy of a higher share of refugees. To this end, we conduct a comparative statistics exercise by calculating the two critical derivatives, $\frac{dp^C}{d\kappa}$ and $\frac{dp^R}{d\kappa}$, given by:

$$\frac{dp^C}{d\kappa} = (1-\alpha)(1-\gamma)\tilde{h}^C\frac{p^C}{H^C - \kappa(1-\gamma)\tilde{h}^C} > 0$$

$$\frac{dp^R}{d\kappa} = p^R\left(\frac{\gamma\tilde{h}^R}{H^R - \gamma\tilde{h}^R\kappa} - \frac{\alpha(1-\gamma)\tilde{h}^C\Gamma_1 A\left[H^C - \kappa(1-\gamma)\tilde{h}^C\right]^{\alpha-1} + \tau}{\Gamma_1 A\left[H^C - \kappa(1-\gamma)\tilde{h}^C\right]^\alpha - \kappa\tau}\right).$$

From these expressions, it is easy to see that the first derivative is always positive. Intuitively, for a given total supply of commercial real estate services, a higher share of refugees implies a lower supply to the local economy for other uses, and hence a higher price $p^C$.

37 As refugees compete with domestic residents for a fixed supply of residential or commercial real estate services, a higher share of refugees can be interpreted as a demand shock.

In contrast, in the residential sector, the sign of the derivative is ambiguous and depends on parameter values. By assumption, as in the commercial sector, a higher share of refugees has the same direct negative effect on the net supply to residents. However, a higher share of refugees also implies a larger lump-sum transfer from residents to refugees and a lower output due to the lower net supply of commercial real estate, which ultimately leads to lower consumption and thus lower residential real estate prices. In equilibrium, the overall impact of a higher refugee share on residential prices depends on the relative strength of these two forces, which in turn depends on parameter values.

### D.2 Calibration

To explore the sensitivity of the derivative of *residential* property prices to the share of refugees, we conduct a simple calibration exercise. We will then vary $\kappa$ and see how this affects the steady state.

Table D1 summarizes the calibration. The discount rate, $\beta$, is set to 0.98, a conventional value at annual frequency. The gross interest rate, $R$, is set to the average level of the German government bond yield between 2009 and 2014, deflated with the average ex post CPI inflation rate. The share of residential housing expenditure in total expenditure, given by $\frac{\omega}{1+\omega}$ in the model, is set to 56%, matching the average share of housing wealth in total assets between 2009 and 2014 from aggregate household balance sheet data from the Bundesbank. For the collateral constraint parameters, $\{\phi, \phi^R, \phi^C\}$, we choose a homogeneous 20% that is a conventional value in the literature. The parameter $\alpha$ is set to 38%, matching the share of tangible assets in total assets as in Table A2. The technology level, $A$, is set to normalize output so that $y=1$ when $\kappa=0$. Similarly, $H^C+H^R=1$. The split between $H^R$ and $H^C$ is set equal to the average share of residential and commercial built-up area in total built-up area from the IOER Monitor in 2008, that is, $\frac{H^R}{H^R+H^C}=56.74\%$. The share of refugees accommodated in residential housing, $\gamma$, is set to the average share of refugees housed in independent accommodations across the 13 non-city-states in Table 2, that is, 53.69%.

To calibrate the residential and commercial housing services consumed by refugees, $\{\tilde{h}^R, \tilde{h}^C\}$, and the per capita fiscal transfer, $\tau$, we first assume that the nominal per capita transfer for nonhousing consumption is 354 euros per month for both types of refugees, or 14% per year of the 2008 national per capita income, $\frac{12*354}{31,308}$.[38] We then assume that refugees have the same preferences as local residents, meaning that their ratio of housing expenditure to nonhousing expenditure is also given by $\omega$. Under this assumption, we obtain the total per capita government transfer as $\tau=\frac{12*354}{31,308}*(1+\omega)$. To obtain an estimate of the per capita residential (commercial) stock of housing services absorbed, we further assume that the price of housing services consumed by the refugees is a unit of their consumption. Thus, this stock is given by $\frac{12*354}{31,308}*\omega$, yielding $\{\tau=0.3, \tilde{h}^R=\tilde{h}^C=0.17\}$.

### D.3 Comparative Static Analysis

Figure D1 plots the steady-state value of the endogenous variables as we increase the share of refugees, $\kappa$. Panel A reports results for the baseline calibration described above. Consistent with our theoretical derivations, when the local economy has a larger refugee share, both stocks of commercial and residential real estate services available for other uses are lower, and the aggregate fiscal transfer, $T$, is higher. This translates into lower local consumption and output, with a higher level of *commercial* real estate prices. In contrast, in the baseline calibration, *residential* real estate prices decrease with a higher refugee share since the effect of lower consumption (because of a

[38] According to the pertinent legislation (AsylbLG), asylum seekers and refugees are entitled to reduced benefits for 15 months, after which they receive a more generous welfare package in line with the German unemployment benefits for natives. For the first 15 months, if accommodated in a reception center, they receive 135 euros per month in addition to food and accommodation provided at the center. If they are in private accommodations, they receive 354 euros to cover living costs, including food but excluding rent that is paid by the government separately. Here, we assume that living expenses, such as food provided by an accommodation center, have the same value as the difference between these two allowances.

**Table D1**
**Baseline calibration**

| Parameter | Definition | Value | Source and Target |
|---|---|---|---|
| $\beta$ | Discount rate | 0.98 | Conventional value |
| $R = 1+r$ | Gross interest rate | 1.0082 | Average real ex post German bond yield (2009–2014) |
| $\omega$ | Ratio of housing to nonhousing expenditure | 1.27 | Average share of housing wealth in total asset (2009–2014) |
| $\phi$ | Collateral constraint parameter | 0.2 | Conventional value |
| $\phi^R$ | Collateral constraint parameter | 0.2 | Conventional value |
| $\phi^C$ | Collateral constraint parameter | 0.2 | Conventional value |
| $\alpha$ | Share of commercial real estate income in total income | 38% | Share of tangible asset in total assets from Table A2 |
| $A$ | Technology level | 1.38 | Normalization such that output $y = 1$ |
| $H^C$ | Commercial real estate supply | 0.43 | Average share of commercial built-up in total built-up area in 2008 |
| $H^R$ | Residential real estate supply | 0.57 | Average share of residential built-up in total built-up area in 2008 |
| $\gamma$ | Share of refugees housed in residential sector | 0.54 | Average across 13 non-city-states in Table 2 |
| $\tilde{h}^R$ | Refugee residential services | 0.17 | Estimated housing expense for refugee |
| $\tilde{h}^C$ | Refugee commercial services | 0.17 | Estimated housing expense for refugee |
| $\tau$ | Fiscal transfer | 0.31 | Estimated total expense for refugee |

lower income and a higher fiscal transfer) dominates the effect of a lower residential housing stock available to residents, consistent with our empirical findings on the weaker relevance of the share of refugees for residential property prices.

However, panel B of Figure D1 also shows that it is possible to find a model parametrization for which residential prices are higher in cities with higher shares of refugees. For example, this is the case if we calibrate the share of residential built-up in total built-up land, $\frac{H^R}{H^R + H^C}$, to 20%, close to the 5th percentile of the 2008 distribution in the data, meaning that we consider a city in which the supply of residential services is a much smaller than the supply of commercial services. Under this alternative calibration, a higher share of refugees is associated with higher property prices in both sectors.

### D.4 Interaction with a Capital Flow Shock

To see how the share of refugees interacts with the capital flow shock, we conduct a second comparative statics exercise by increasing $\phi$, which loosens the collateral constraint.[39] The dashed lines in panels A and B represent the value of the endogenous variables when we increase the

39 In our simple model, without domestic financial intermediation, an increase in $\phi$ can capture either the credit supply or the credit demand effect of a capital flow shock, as the credit supply is assumed infinitely elastic. Here, we are assuming that, as we document in the paper empirically, a capital flow shock leads to an increase in domestic credit supply captured by looser collateral requirements.

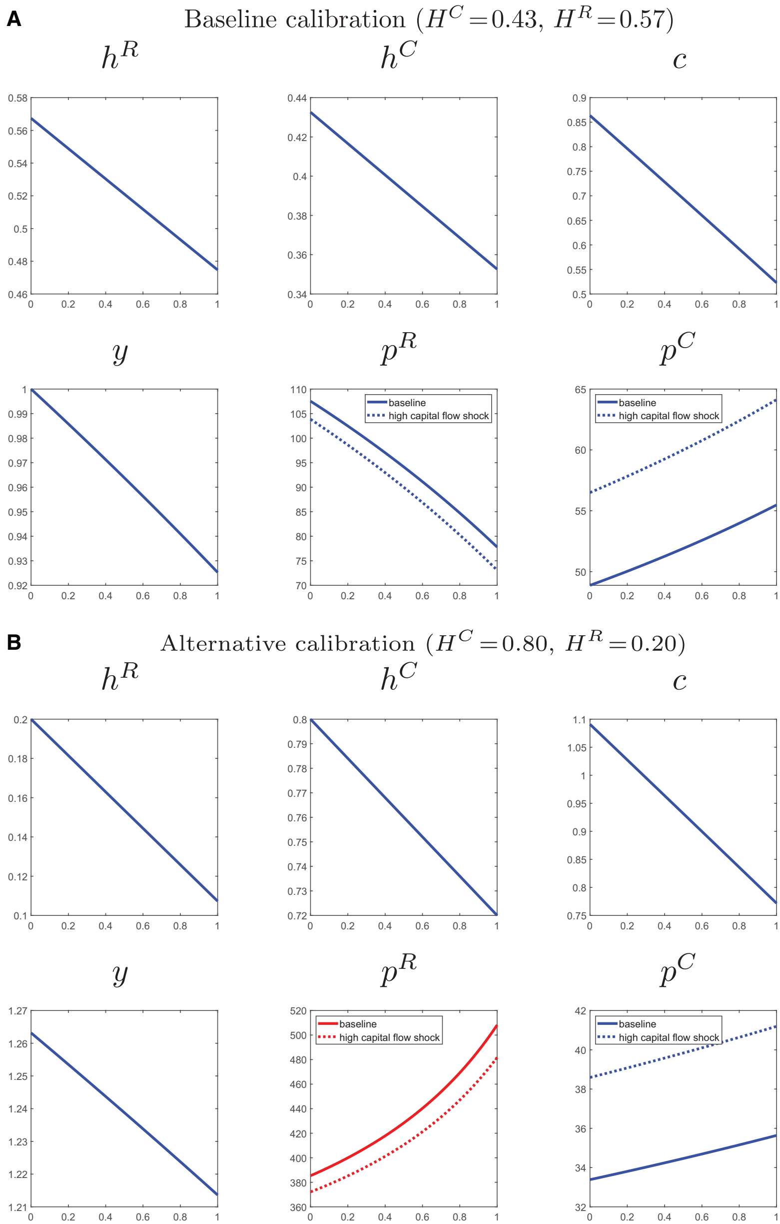


**Figure D1**
**City share of refugees and property prices: Comparative statistics**
The figure reports the comparative statistics of the steady-state equilibrium variables ($y$-axis) with respect to a change in the city share of refugees, $\kappa$ ($x$-axis). Panel A reports the baseline calibration as in Table D1. Panel B reports an alternative calibration in which we change the stock of residential and commercial real estate services $H^R$ and $H^C$. The solid lines represent the comparative statistics with the baseline value of the collateral constraint parameter, $\{\phi, \phi^C, \phi^R\}$=0.2, and the dashed lines represent a higher value of this parameter, $\{\phi, \phi^C, \phi^R\}$=0.4, which can be interpreted as a capital inflow shock as discussed in the text.

collateral constraint parameters $\{\phi, \phi^C, \phi^R\}$ from 0.2 to 0.4, a value that is at the upper end of the range used in the literature.

The capital flow shock increases commercial property prices while it decreases residential prices. In the current simple setup, without labor and capital, consumption falls because debt service increases, but output is unchanged following the shock. In principle, however, our model could generate a consumption boom if we were to introduce investment and endogenous labor supply. Moreover, labor and capital would not detract from the direct impact of the higher refugee share on the supply of commercial real estate services for other uses that drives the relevance of our exposure measure. As a result, the capital flow shock does not alter the relationship between the refugee share and real estate prices.

### D.5 Comparing Model-implied and Data Correlations

While our model is not suitable for quantitative analysis, we want to use it to evaluate the strength of the statistical relation between the refugee share and commercial property prices. To this end, we first feed the model with the distribution of the share of refugees in 2008 for the 85 cities in our sample, which we denote $\hat{\kappa}$, while keeping all other parameter values the same as in Table D1. We then evaluate the association between $\hat{\kappa}$ and the *model-implied* commercial real estate prices, which we denote $\hat{p}^C$. Finally, we compare the model-based estimate of the strength of this relation with the one in the data.

We compute both the correlation between these two variables and the regression slope of $\hat{p}^C$ on $\hat{\kappa}$, as $\hat{p}^C$ has a different variance in the model and in the data. In the data, $\hat{p}^C$ is the 2014 value of the property price indexes.[40] In the model, instead, $\hat{p}^C$ is a cross-section of steady-state values implied by the distribution of $\hat{\kappa}$ in the data. The independent variable in the model and the data is exactly the same.

The model-implied correlation between $\hat{\kappa}$ and $\hat{p}^C$ is approximately 1, compared with 0.32 in the data. Thus, the model-implied correlation is much larger than in the data for two reasons. First, the curvature in the comparative static relation in the model between $p^C$ and $\kappa$ is minimal given our calibration (Figure D1). Second, the distribution of the share of refugees in the data is centered on relatively low values, varying from 0.04% to 3.96% across the 85 cities in our sample. This is consistent with the fact that our results are not driven by a few cities with a high share of refugees. Note here that the model-implied correlation is close to the model-based estimate in Favilukis and Nieuwerburgh (2017), who document that a 10% increase of out-of-town buyers into a city can cause a 10% increase in housing prices, thus pointing to an elasticity of one.

Comparing correlations controls for the differences in standard deviations in the variables that do not coincide in the model and in the data. For robustness, we also compute the regression slope of $\log\left(\hat{p}^C\right)$ on $\log\left(\hat{\kappa}\right)$ in the model and data. In this case, the model predicts a much smaller slope than in the data, with values equal to 4.91% and 0.05%, respectively.

We conclude from this analysis that the sign of the impact of the refugee share on commercial property prices is fully consistent with the model-implied one, but its economic size cannot be evaluated properly. This is because we do not observe the absolute level of the commercial property prices before the boom, in 2008, when one could reasonably assume that local real estate markets were in a long-run equilibrium in Germany.

### D.6 Refugees' Shelter Needs and Commercial Real Estate Turnover

In the absence of suitable data on refugees' absorption of commercial real estate services, we attempt to estimate the significance of their demand from market sources and the parameters of the policy framework we discussed above. We start from the observation that the average annual

40 This is equivalent to the cumulative property price change between 2009 and 2014 because all city indexes are normalized to 100 in 2008. Here, also note that we focus on the *cross-section* correlation, while in Table 1 we report the panel correlation for consistency with the rest of the empirical analysis in the paper.

national flow of new refugees into Germany is about 90,000 according to official statistics.[41] We then note that the average space per refugee occupied is 7 sq. m.[42] We also know that the average share of refugees housed in accommodation centers is 0.46 in non-city-states (Table 2). Together, these parameters imply an estimated absorption of commercial real estate space of 289,800 sq. m. per year nationally.

For comparison, we look at the turnover in the hotel segment of the commercial real estate sector. A back-of-the-envelope estimate of the average number of square meters turned over, annually, in the German hotel industry during 2009–2014, is 5.5 million. This implies that refugees in accommodation centers represent 5.3% of the hotel industry turnover, which is a nontrivial share that can affect property prices. We obtain the 5.5 million sq. m. estimate as follows. We take the nominal annual national commercial transaction volume (about EUR 25 billion per year, on average).[43] We then consider the Hotel share of the industry, which is 8%.[44] Finally, we calculate an annual average price per square meter of hotel properties based on a 2019 estimate of the price of commercial land in the top-five German cities in 2019 (excluding city-states), deflated by using the nominal average annual commercial property price growth of 2.5%, as in our data.[45] This, on average during 2009–2014, is EUR 363 per sq. m. Thus, EUR 25 billion times 0.08 divided by 363 = 5.5 million.

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[41] By comparison, Germany received approximately 2.5 million applications during the period 1985–2007 according to Eurostat data.

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