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Dollar funding and housing markets:
The role of non-US global banks

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Abstract

House prices co-move considerably across countries. We show how non-US global banks and their exposure to US dollar funding conditions help explain this comovement. When the dollar appreciates, mortgage lending and house prices decrease more in borrower countries whose non-US creditor banks are more exposed to dollar funding conditions. As US dollar funding conditions vary, borrowing country pairs with higher joint exposure to US dollar funding conditions via their non-US creditor banks exhibit a higher synchronization of mortgage credit and house price growth. We capture the exposure to dollar funding conditions by the bilateral treasury basis between the currency of the non-US global creditor banks' nationality and the US dollar, a choice that we motivate in a simple value-at-risk model. Our results identify a novel international spillover channel of US dollar funding conditions. Because it works through heterogenous dollar funding exposures among creditors, this new channel is neither linked to common-lender exposures nor to currency mismatches on borrower countries' balance sheets, typically associated with the financial channel of the exchange-rate.
Key words: house prices, synchronization, US dollar funding, dollar cycle, US treasury basis, convenience yield, capital flows, global banks, global banking network
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1 Introduction

House prices co-move considerably across countries. Figure 1 shows the pairwise rolling-window correlation of house price growth between 35 advanced and emerging economies in our sample. Average house price synchronization varies substantially over time and peaked in the run-up to the 2008 Global Financial Crisis, and again in the euro area debt crisis. Importantly, the degree of synchronization varies significantly across country pairs as measured by the interquartile range.

Understanding international house price synchronization is highly policy-relevant. In most countries, housing wealth represents the largest component of net household wealth and is the single most important collateralizable asset. Identifying the drivers of the international synchronization of house prices is therefore paramount to understanding macro-financial linkages and financial stability at the global level.

We show that the variation in US dollar funding conditions drives the international synchronization of house prices. House price growth becomes synchronized as any two countries' housing markets are jointly exposed to US dollar funding variations through non-US global banks' lending to these countries. This is what we call "dollar co-dependence". We show that this dollar co-dependence is the key link between US dollar funding conditions and housing markets worldwide, explaining the time and cross-country variation of house price synchronization (Figure 1).

Dollar co-dependence combines two linkages, which reflect the structure of the global dollar banking network: global banks' sensitivity to dollar funding conditions, and borrowing country pairs' dependence on credit from global non-US banks. If dollar funding conditions ease, global non-US banks have access to cheaper funding and increase their foreign lending, which is mostly denominated in US dollars. We show that the additional international lending translates into higher mortgage credit in the recipient country and ultimately higher house prices. The magnitude of the effect on house prices, however, differs across borrowing countries. The higher the dependence of a borrowing country on foreign lending from global banks, combined with a higher sensitivity of their non-US global creditor banks' to US dollar funding conditions, the stronger is the effect on mortgage credit and thus house prices. The higher this exposure to dollar funding conditions for a pair of any two borrowing countries—the higher the dollar co-dependence—the higher the co-movement of house prices.

The first key element of this spillover mechanism is non-US global banks' sensitivity to US dollar funding conditions. Non-US banks are key intermediaries in the global financial system as they account for the overwhelming share of international bank lending globally (Aldasoro and Ehlers (2019)). At the same time, non-US global banks significantly depend on funding in US dollars to finance their US dollar-denominated loans. This dependence on US dollar funding is what makes non-US global banks sensitive to variations in US dollar funding conditions, notably the US dollar exchange rate. As we show in a stylized model of international bank lending, the sensitivity of non-US global banks' foreign lending to US dollar funding conditions is a function of the bilateral treasury basis between non-US global banks' home currency and the US dollar. The bilateral treasury basis is the difference between the return on a US treasury bond and the synthetic dollar return on a foreign government bond of the same maturity in domestic currency. The recent literature (Krishnamurthy and Lustig (2019)) interprets the US dollar treasury basis as a convenience yield: investors are willing to forego some yield in return for the liquidity and safety of dollar-denominated US government securities.

This paper emphasizes a particular implication of the convenience yield interpretation of the treasury basis. Specifically, the bilateral treasury basis captures the cost disadvantage of non-US global banks relative to US banks when procuring US dollar-denominated funding synthetically, using their home currency deposit base. Non-US global banks with a higher bilateral US treasury basis face higher synthetic US dollar funding costs. Thus, non-US global banks will fund more of their foreign lending directly in US dollar funding markets through wholesale funding or by issuing dollar deposits. However, direct dollar funding exposes non-US global banks to exchange rate risk and therefore ties up balance sheet capacity. Notably, a (transitory) appreciation of the US dollar lowers future expected returns (in the non-US bank's home currency) on dollar-denominated lending. We show in a value-at-risk (VaR) framework that this reduces non-US banks' risk-taking capacity, forcing them to reduce their foreign lending. In our stylized theoretical model, a bank optimally trades off costs of synthetic funding against the cost of balance sheet capacity. The model predicts that non-US global banks' reduction in foreign lending is stronger for banks with a higher bilateral US treasury basis as it implies a higher cost disadvantage in synthetic US dollar funding.

Hence, when US dollar funding conditions ease (tighten), this frees (ties) up balance sheet capacity of non-US global banks leading them to extend more (less) foreign credit. Counterparty banks in the borrowing countries absorb this expansion (contraction) of foreign credit and expand (reduce) domestic mortgage credit. This results in upward (downward) pressure on house prices. As this pattern replicates itself across borrowing countries, house prices become internationally synchronized. We measure a borrowing country's exposure to this mechanism—to which we refer as dollar dependence—as the market-share weighted average of the bilateral US treasury bases of their respective foreign creditor banks. We obtain the market-share weights by drawing on granular bilateral foreign lending data from the BIS consolidated banking statistics (CBS). Hence, our measure of dollar dependence can be interpreted as an "effective" (i.e. foreign-lending weighted) treasury basis of the borrowing country. This effective treasury basis reflects a combination of non-US global lender banks' exposure to US dollar funding conditions—as measured by their respective bilateral treasury bases vis-à-vis the US dollar—and borrowing countries' heterogeneous exposures to their respective non-US global creditor banks as measured by the market shares of these banks in providing foreign credit to the borrowing country.

Our focus on the role of non-US global banks in transmitting dollar funding conditions globally builds on the double-decker structure of the global banking system first emphasized by Bruno and Shin (2014) and also highlighted in Hale and Obstfeld (2016). To our knowledge, ours is the first paper to explore empirically how the structure of the global banking network affects the synchronization of real outcomes, and in particular of real estate markets. A key feature of the global banking network is that banks headquartered in a few advanced non-US economies, notably Germany, France, the UK, the Netherlands, Switzerland and Japan, account for the bulk of global foreign credit as well as for the largest-sized bilateral lending flows between countries (Aldasoro and Ehlers (2019)). By showing how dollar funding conditions are transmitted to borrowing countries through this network, our analysis contributes to a recent literature that documents the central role of non-US global banks in the international financial system (Ivashina et al. (2015), Borio et al. (2017, 2016), Du et al. (2018a); Iida et al. (2018); Barajas et al. (2019)).[1]

Our empirical analysis proceeds in two steps. First, we show that individual countries' house price growth depends on dollar funding conditions measured by the US dollar exchange rate, and

[1]While the non-US global banks US dollar dependence is a key analytical feature, our analysis does account for the role of US banks in the construction of our measure of dollar dependence. The US treasury basis of US banks is zero by definition, but still enters borrowing countries' dollar dependence computed as the market-share weighted average of the bilateral US treasury bases.

their indirect US dollar funding exposure as measured by our concept of dollar dependence. Vindicating the model's predictions, the analysis confirms that the foreign lending of non-US banks with wider bilateral US treasury basis is more sensitive to variations in US dollar funding conditions. Therefore, easing (tightening) US dollar funding conditions loosen (tighten) the leverage constraint of non-US banks by more the wider the bilateral US treasury basis, leading them to provide more (less) credit to counterparty banks in various foreign borrowing countries. Turning to house price synchronization next, we show analytically that the comovement between house prices of any two borrowing countries is determined by the product of their respective dollar dependencies. This product constitutes our measure of dollar co-dependence. We show empirically that dollar co-dependence is a key driver of house price synchronization. To shed light on the transmission mechanism, we show the same link between dollar co-dependence and mortgage credit growth and synchronization, respectively.

Our empirical implementation is based on the framework by Landier et al. (2017), which we expand to take account of heterogeneous exposures to US dollar funding shocks. Landier et al. (2017) document that banking liberalization in the United States in the period 1970 to the mid-1990s increased the synchronization of house price movement across states because, as banks integrated across state borders, mortgage lending across states became more exposed to idiosyncratic shocks to the same banks, leading to more house price synchronization.[2]

Importantly, in our framework, the synchronization of house prices between two arbitrary borrowing countries will depend not only on whether they are exposed to common lender banks, as emphasized by Landier et al. (2017), but also on their lender banks' sensitivity to US dollar funding conditions. To see the gist of our argument, consider an extreme case in which country A borrows exclusively from lender banks C and country B from lender banks D, respectively. Hence, the two countries A and B have no common lender. Idiosyncratic shocks to lender banks C affect only country A, and idiosyncratic shocks to lender banks D only affect country B. Therefore, uncorrelated lender bank-specific shocks will not lead to co-movement in the foreign lending supply to A and B. However, if both C and D have correlated funding sources because both are exposed

[2] Instead of focusing on individual banks, our analysis focuses on entire *banking systems*, i.e. the country level aggregate of banks providing and receiving foreign credit. BIS CBS data allow us to construct bilateral country-level exposures of a borrowing country to the banks headquartered in the countries providing foreign credit, henceforth also called *lending banking systems*. For ease of exposition, we will, however continue to use the term *lender bank* or just *bank* instead of *banking system* whenever this does not lead to ambiguity.

to dollar funding risk, then fluctuations in US dollar funding conditions will affect both C and D and therefore lead to synchronized outcomes for countries A and B. Hence, A and B are effectively co-dependent on US dollar funding although they do not share common lender banks. One key feature of our framework is that we can empirically separate this impact of dollar funding shocks on house price synchronization via non-US global banks from the impact of common-lender specific shocks—including shocks to US banks. As we show, it is indeed the former dollar funding channel that accounts for the bulk of the variation in international house price synchronization.

Our empirical specifications for house price synchronization allow us to control for a rich set of confounders. In particular the inclusion of borrower country-time-specific effects effectively rules out that our results are driven by shifts in credit demand in borrowing countries. To further buttress the causal interpretation of our results, we also eliminate any unobserved, time-varying country-pair specific influences possibly leading to reverse causality between house price synchronization and dollar co-dependence. Such feedbacks could arise, for example, if two borrowing countries specialize in a particular export industry in which US dollar financing is particularly prevalent or if they engage in predominantly US dollar-denominated trade with each other. Then the joint (country-pair specific) exposure to the same US dollar demand factors could lead to time-varying co-movement in foreign borrowing and house prices, while also affecting the US dollar borrowing of the country pair's global creditor banks. To address this possibility we build on Gabaix and Koijen (2024) and construct a granular instrumental variable (GIV) that purges lender banks' dollar dependence of the potential feedback from common demand factors in borrowing countries. Thus, our approach extends the methodology of Landier et al. (2017) to settings where quasi-natural experiments are not readily available for identification. To our knowledge this constitutes the first application of the GIV framework to the synchronization of economic variables.

Our analysis contributes to the literature on international capital flows and house prices such as Aizenman and Jinjarak (2009); Ferrero (2015); Hoffmann and Stewen (2020) and Sá et al. (2014). With only a few exceptions (Alter et al. (2018), Milcheva and Zhu (2016)), this literature has not focused on international correlations in house prices or explored the role of the global banking network in transmitting dollar funding conditions to real estate markets.

To our knowledge, ours is the first study to link international housing markets with the literature on the global financial cycle (Bruno and Shin (2015); Cerutti et al. (2017); Habib and Venditti

(2019); Miranda-Agrippino and Rey (2020); Rey (2015)). This literature has shown that global capital flows are driven by a few dominant common factors that can directly be related to shocks to the balance sheets of globally active financial intermediaries. Recent research has singled out the US dollar exchange rate as one particularly important such common factor (Avdjiev et al. (2019); Gopinath et al. (2020); Boz et al. (2017); Gopinath and Stein (2021, 2018)), identifying a financial channel through which the exchange rate affects cross-border capital and trade flows. According to the financial channel of the exchange rate, international lending in dollars leads to currency mismatch on borrowers' balance sheets which makes firms and households in borrowing countries vulnerable to a dollar appreciation. The deterioration of borrowers balance sheets then also reduces the risk-taking capacity of both non-US and US global banks, making the dollar exchange rate a common factor in cross-border lending. In our setting, a (temporary) appreciation of the dollar lowers the expected returns of dollar lending in terms of non-US global banks' home currency, reducing their risk-taking capacity and thus their dollar lending. This mechanism is independent of currency mismatch on borrower countries' balance sheets and it is present only for non-US global banks. Hence, while the identity of the lenders is irrelevant in the classical version of the financial channel, in our setting it matters from whom (i.e. which lenders) a country borrows. Our approach therefore allows us to identify the causal impact of dollar funding conditions on a given borrowing country by exploiting the heterogeneity of lenders' exposure to these conditions (as indicated by their treasury basis). Hence, in our mechanism the lending network of non-US global banks—rather than internationally active US or ultimate borrowing country banks—takes center stage and we show that this mechanism affects house prices worldwide.

The remainder of the paper is organized as follows. Section (2) introduces the concept of dollar (co-)dependence and provides a first look at the data. Section (3) explains the analytical framework used for empirical analysis, while section (4) presents details on the data. Section (5) presents and discuses our main results, including our instrumental variable estimates. Section (6) provides additional robustness checks. Section (7) concludes.

2 Dollar (co-)dependence and house prices: a first look

To study how variations in US dollar funding conditions affect house price growth through non-US global banks, we introduce the concepts of dollar dependence and dollar co-dependence. These concepts formalize the exposure of borrowing countries to US dollar funding conditions via their respective lender banks' sensitivity to US dollar funding conditions.

Formally, let $\mathcal{B}(i)$ be the set of (creditor) banks lending to borrowing country i and let λ_t^b an indicator of the sensitivity of lender bank b to changes in US dollar funding conditions. Then we define the dependence of borrowing country i to dollar funding conditions—henceforth labeled "dollar dependence" as

$$\text{DD}_t^i = \sum_{b \in \mathcal{B}(i)}^{N} \omega_t^{b,i} \lambda_t^b \tag{1}$$

where $\omega_t^{b,i}$ is the market share of lender bank b in total foreign bank lending to borrowing country i at time t. Our measure of λ_t^b is the bilateral treasury basis—the deviation between government bond yields denominated in the home market currency of bank b and US government bond yields—defined as

$$\lambda_{n,t}^b = i_{n,t}^b - i_{n,t}^{\$} - \rho_{n,t}^b$$

where $i_{n,t}^b$ is the n-year home-currency government bond yield in lending banking system b, $i_{n,t}^{\$}$ is the n-year US treasury bond yield, and $\rho_{n,t}^b$ is the n-year market-implied forward premium for hedging currency i against the US dollar.

While we further motivate this choice in section 2.1 and provide a formal theoretical foundation for it in Appendix B, the intuition is as follows: an increase in λ_t^b implies higher funding costs for non-US banks using hedged positions funded from their domestic (home-currency denominated) deposit base as a source of their US dollar lending.[3] This induces non-US banks to borrow US dollars directly through wholesale funding or by issuing dollar deposits. Unlike synthetic funding, this ties up balance sheet capacity because it exposes non-US banks' (home currency denominated) balance sheets to unhedged exchange rate risk. Thus, the bilateral treasury basis is a suitable

[3] Note that our definition of the treasury basis follows Du et al. (2018b), so that an increase in λ means that synthetic dollar funding becomes more expensive. We will generally refer to this increase as a "widening" of the basis. Note that this differs from the normalization in Krishnamurthy and Lustig (2019) who define the treasury basis as $-\lambda$.

measure for $\lambda^b_{n,t}$, as it captures non-US bank's sensitivity to US dollar funding shocks.

For borrowing country i, dollar dependence DD^i is constructed as an "effective" treasury basis across all its lender banks. It is a weighted average of the bilateral treasury bases of lender bank b of country i, with the market shares $\omega^{b,i}_t$ of banks b providing foreign credit to country i serving as weights.[4]

To illustrate how the transmission between the dollar and house prices is modulated by DD^i, we run a sequence of cross-sectional regressions on quarterly data from 2000 to 2020:

$$\Delta\text{HP}^i_t = \zeta_t \times \text{DD}^i_{t-1} + \text{constant}_t + \varepsilon^i_t$$

where ΔHP^i_t measures house price growth in country i.[5] Figure 2 plots the sequence of estimated coefficients $\{\zeta_t\}$ against the four-quarter change in the effective US dollar exchange rate, an important measure of US dollar funding conditions (Avdjiev et al. (2019)). The strong negative correlation between the two time series, at -0.4, suggests that house prices rise as the US dollar depreciates, and vice-versa. This link is stronger for countries with higher dollar dependence.

Our identification strategy relies on cross-country heterogeneity in DD^i_t. Figure D.1 plots DD^i_t (relative to its cross-county, time t mean) for a selection of borrowing countries in our sample. Note that DD^i_t varies considerably both across time and across borrowing countries i and that countries change their relative positions quite frequently. This variation is driven by a combination of the heterogenous exposure of borrowing countries i to lender banks b as given by $\omega^{b,i}_t$, as well as by the heterogenous exposure of lender banks b to variations in US dollar funding as given by λ^b_t.

The analytical framework that we propose in section 3 allows us to explore the implications of dollar dependence for the synchronization of house price growth across borrowing countries. It is in this context that we introduce the notion of dollar co-dependence. We define the dollar co-dependence between any two borrowing countries i and j as the product of the individual countries' dollar dependencies:

$$\text{CoDD}^{i,j}_t = \text{DD}^i_t \times \text{DD}^j_t \tag{2}$$

[4]Note that the set of lender banks $\mathcal{B}(i)$ includes the United States. However, the bilateral CIP-deviation of the US vis-à-vis itself is zero. Thus, by construction, DD^i_t captures how dollar funding conditions affect borrowing country i through non-US lending banking systems.

[5]We provide a detailed discussion of our data below in section 4.

As we will show both theoretically and empirically, the synchronization of house price growth in two arbitrary borrowing countries i and j increases in $\text{CoDD}_t^{i,j}$. For a pair of borrowing countries to have a high level of dollar co-dependence the individual dollar dependencies of both countries need to be relatively high.

Figure 3 provides a first illustration of the link between house price synchronization and dollar co-dependence. As shown in Figure D.1, the dollar dependence of individual countries varies considerably over time and relative to other countries. Therefore, in Figure 3, in each quarter, we sort our sample of country pairs by ascending dollar co-dependence into portfolios. We then compute the mean co-dependence and mean house price synchronization for each portfolio over our sample period. Figure 3, which plots these means against each other clearly shows that higher dollar co-dependence is associated with higher house price synchronization. Country pairs with the highest dollar co-dependence at any given point in time display the highest synchronization of house prices.[6]

Note that high levels of dollar co-dependence and thus a high synchronization of house price growth can occur between borrowing countries with exposure to entirely distinct sets of lender banks. What matters for dollar co-dependence is that borrowing countries are dependent on lender banks that are themselves highly exposed to variations in US dollar funding.

2.1 Measuring lender banks' sensitivity to US dollar funding conditions

The US treasury basis proxies the cost disadvantage that a non-US bank faces relative to US banks when it raises US dollar denominated funds for repayment in n years synthetically by raising deposits in its own currency and then entering a foreign exchange swap for US dollars, as opposed to raising US dollar-denominated funding directly in the market for wholesale dollar funding or deposits. To see why the bilateral US treasury basis may be a useful measure of non-US banks' exposure to changes in US dollar refinancing conditions, consider the options a non-US bank faces when it finances a foreign US dollar-denominated loan.

The first option for the non-US bank would be to use its domestic base of insured deposits denominated in domestic currency to fund US dollar lending positions. Financial stability regulation will generally require positions financed by insured deposits to be fully hedged (Ivashina et al.

[6] We find this conclusion to be robust to changing the number of portfolios.

(2015)). Hence, the bank will only be able to use its home currency deposits for synthetic US dollar funding, which embodies the hedging of currency risk by definition. The bilateral US treasury basis $\lambda^b_{n,t}$ captures the costs of this hedge.

The second option for the non-US bank is to fund US dollar denominated lending with US dollar denominated liabilities raised directly in the US dollar funding market. We refer to this option as direct US dollar funding. The non-US bank will incur capital charges for this foreign currency position, as they are subject to exchange rate risk. For home country regulation, the aggregate balance sheet of the non-US bank is denominated in its non-US home currency. Note that the exchange rate risk arises solely due to the fact that the bank has a foreign currency position and is independent of whether US dollar-denominated lending is matched by direct US dollar-denominated borrowing. This exchange rate risk ties up balance sheet capacity of the non-US global bank and imposes a shadow cost unique to non-US banks.[7]

The non-US bank optimally trades off the cost of both funding options. In a model provided in appendix B we formalize this trade-off for a non-US bank that operates under a value-at-risk (VaR) constraint. Intuitively, the model predicts that the bank equates the marginal cost of hedging (captured by the bilateral treasury basis) with the shadow cost of balance sheet capacity tied up by a marginal unit of direct dollar funding. A wider (narrower) bilateral treasury basis therefore increases (lowers) the share of the non-US global banks' directly funded dollar lending.

Importantly, our model also implies that the non-US global bank becomes more sensitive to variations in US dollar funding conditions when the treasury basis increases. The intuition is that lender banks with a wider treasury basis will have a higher share of direct dollar funding which translates into higher shadow costs of balance sheet capacity. This makes them particularly sensitive to changes in US dollar funding conditions, such as an increase in US interest rates or of an appreciation of the US dollar exchange rate (Avdjiev et al. (2019)).

[7] Non-US global banks' cost disadvantage might be further aggravated, as non-US banks lack a broad base of insured deposits in the US. Hence, non-US banks are perceived as riskier than US banks, which in turn raises their direct funding costs relative to US banks (Ivashina et al. (2015))

3 Analytical framework

We adapt and extend the methodological framework of Landier et al. (2017) for our analysis. The authors show that an increase in the co-movement of house prices across US states between the late 1970s and the mid 1990s can be associated with the emergence of multi-state banks in the wake of the US interstate banking liberalization implemented over the same period. The key mechanism in their framework is a common lender effect: house prices in US states in which multi-state banks have relatively large market shares exhibit higher co-movement as these states are relatively more exposed to the idiosyncratic shocks of multi-state banks.

Relative to their setting, we innovate along two dimensions. First, we take their setup to the international level and analyze the effect on house price co-movement across countries. That is, our unit of analysis are entire country-level banking systems, i.e. the aggregate of all banks headquartered in a country.

Second, we uncover that the international synchronization of house price growth between borrowing countries depends on their respective lender banks' heterogenous exposure to refinancing conditions in US dollars, as captured by borrowing countries' dollar co-dependence. The lender banks that two arbitrary borrowing countries are exposed to do not need to be common lenders. For the effect of US dollar refinancing conditions on house price growth synchronization to be increasing in borrowing countries' dollar co-dependence, it is sufficient to consider borrowing countries' exposure to dollar funding variations via their lender banks' sensitivity to these variations. In addition, our framework also encompasses a common lender effect as the theoretical setup allows for borrowing countries' exposure to idiosyncratic shocks of common lender banks. Empirically, however, results in section 5 suggest that the exposure to dollar funding conditions is the key channel in our international setup.

Following Landier et al. (2017), we conjecture that foreign bank credit supply to banks in borrowing country i drives house price growth $\frac{\Delta \mathrm{HP}_t^i}{\mathrm{HP}_{t-1}^i}$ in borrowing country i with an elasticity of α, so that

$$\frac{\Delta \mathrm{HP}_t^i}{\mathrm{HP}_{t-1}^i} = \alpha \frac{\Delta L_t^i}{L_{t-1}^i} + \varepsilon_t^i \tag{3}$$

where L_t^i are aggregate foreign claims on country i, ε_t^i is a shock specific to borrowing country i

and captures credit demand, and $\alpha > 0$ is the elasticity of house prices to lending.

Furthermore, for the foreign lending supply provided by lending banking system b to banks in country i we posit that

$$\frac{\Delta L_t^{b,i}}{L_{t-1}^{b,i}} = \gamma_t + \lambda_{t-1}^b \zeta_t + \eta_t^b \tag{4}$$

where $L_t^{b,i}$ measures the bilateral foreign claims of lending banking system b on borrowing country i, γ_t is a global factor that is homogeneous in its impact across borrowing countries and lending banking systems alike, and where η_t^b is an idiosyncratic shock specific to lending banking system b. Our analysis in this paper focuses on the role of ζ_t, to which we assign the role of a common US dollar funding shock.

Importantly, lending banking systems differ in their exposure to this dollar funding shock ζ_t. This assumption drives the empirical implications of our theory for the impact of the dollar co-dependence on the synchronization of housing markets. The heterogenous exposure is given by λ_{t-1}^b.

Using that $L_t^i = \sum_{b \in \mathcal{B}(i)} L_{t-1}^{b,i}$, we can consolidate equations (3) and (4) to obtain

$$\frac{\Delta \mathrm{HP}_t^i}{\mathrm{HP}_{t-1}^i} = \alpha \left(\sum_{b \in \mathcal{B}(i)}^{N} \left(\lambda_{t-1}^b \zeta_t + \eta_t^b + \gamma_t \right) \omega_{t-1}^{b,i} \right) + \varepsilon_t^i$$

or equivalently

$$\frac{\Delta \mathrm{HP}_t^i}{\mathrm{HP}_{t-1}^i} = \alpha \gamma_t + \alpha \left(\sum_{b \in \mathcal{B}(i)}^{N} \omega_{t-1}^{b,i} \eta_t^b \right) + \alpha \underbrace{\left(\sum_{b \in \mathcal{B}(i)}^{N} \omega_{t-1}^{b,i} \lambda_{t-1}^b \right)}_{\text{Dollar dependence}} \times \zeta_t + \varepsilon_t^i \tag{5}$$

where we have used that the market share of lender bank b in country i is given by $\omega_t^{b,i} = L_t^{b,i} / L_t^{b,i}$.

As indicated by the under-braced term, equation (5) establishes a direct link between house price growth of borrowing country i and US dollar funding conditions depending on country i's US dollar dependence. Assuming that the lending banking system specific supply shocks, η_t^b, the borrowing country specific shock, ν_t^i, the global factor γ_t and the factor ζ_t are mutually uncorrelated, we can derive an expression for the time-varying conditional covariance of house price growth between any two borrowing countries i and j:

$$\text{HPcov}_{t-1} = \alpha^2\sigma_\gamma^2 + \alpha^2\sigma_\eta^2 \underbrace{\left(\sum_{b\in\mathcal{B}(i)\cup\mathcal{B}(j)}^{N} \omega_{t-1}^{i,b}\omega_{t-1}^{j,b}\right)}_{\text{co-Herfindahl}} + \alpha^2\sigma_\zeta^2 \underbrace{\left(\sum_{b\in\mathcal{B}(i)}^{N} \omega_{t-1}^{i,b}\lambda_{t-1}^{b}\right)\left(\sum_{b\in\mathcal{B}(j)}^{N} \omega_{t-1}^{j,b}\lambda_{t-1}^{b}\right)}_{\text{dollar co-dependence}} \tag{6}$$

The first under-braced term on the right hand side captures the effect on synchronization that stems from the idiosyncratic shocks affecting common lending banking systems, i.e. the common lender effect. Landier et al. (2017) refer to this term as the co-Herfindahl index. For lending banking system specific shocks to have a big impact on house price growth synchronization, a lending banking system must have high market shares in both borrowing countries i and j so that the product of the market shares $\omega_{t-1}^{i,b}$ and $\omega_{t-1}^{j,b}$ becomes big.

The second under-braced term is the focus of this paper. This term captures the dollar co-dependence as defined in equation (2) above. The term reflects the impact of any two borrowing countries' simultaneous indirect exposures to fluctuations in US dollar funding conditions through their respective lender banks on the synchronization of house price growth.

To obtain our empirically testable hypothesis, from (5) we write the conditional variance of house price growth as

$$\sigma^2\left(\frac{\Delta\text{HP}_t^i}{\text{HP}_{t-1}^i}\right) = \sigma_\varepsilon^2 + \alpha^2\sigma_\gamma^2 + \alpha^2\sigma_\eta^2\text{CoHFI}_{t-1}^{i,i} + \alpha^2\sigma_\zeta^2\text{CoDD}_{t-1}^{i,i} \tag{7}$$

where CoHFI is the co-Herfindahl index and CoDD is the dollar co-dependence as defined in (6) above. In appendix (C), we show how to use (6) and (7) to obtain a linearized expression for the house price correlation between countries i and j of the form

$$\text{HPcorr}_t^{i,j} = \kappa + a \times \text{CoHFI}_{t-1}^{i,j} + b \times \text{CoDD}_{t-1}^{i,j} + n_{t-1}^{i,i} + n_{t-1}^{j,j} \tag{8}$$

where κ is a constant, a, and b are positive functions of the parameters α, σ_ε, σ_η, σ_γ, and σ_ζ and n_{t-1}^i and n_{t-1}^j are country-specific nuisance terms. In our empirical specification the latter will be absorbed by country-time fixed effects. Similar to the expression for the covariance in equation (6) above, the first term, κ, captures the relative importance of the common shocks γ_t, and ζ_t and of the idiosyncratic shock ε_t. The higher the volatility of the common shocks relative to the idiosyncratic shock, the higher will be the house price correlation. The interpretation of the second and third terms remains unchanged relative to equation (6) above. Equation (8) provides the empirically testable hypothesis investigated in section (5.2).

4 Data

Our sample comprises a quarterly panel of 35 OECD borrowing countries from 2000Q1 to 2019Q4, covering house price growth, mortgage growth and our measure of dollar dependence DD_t^i. Table A.1 in the appendix provides summary statistics for these main variables as well as as for the key factor variables driving US dollar funding conditions. From these main variables, we then compute our various synchronization measures for house price and mortgage growth as well as the dollar co-dependence $\mathrm{CoDD}^{i,j}$ and we also construct the co-Herfindahl index. Here, we briefly discuss the sources of house price and mortgage data and the data on bilateral treasury bases, λ^b and market shares, $\omega^{b,i}$, that we use to compute our dollar dependence measure DD_t^i.

House prices and mortgage credit: We measure house price growth over four quarters ahead based on a country-level residential real house price index available from the OECD for 35 borrowing countries.[8] Similarly, mortgage credit growth is computed over four quarters ahead based on the time series of credit to households and non-profit institutions serving households, provided by the BIS. The sample period for both data sets is 2000Q1-2019Q4 . For each borrowing-country pair, the international synchronization of house price growth is measured as the 16-quarter-ahead rolling-window correlation of house prices, and analogously for the synchronization of mortgage credit growth. Because house price and mortgage growth are themselves measured four quarters ahead, this results in a five-year window. As a result, synchronization regressions reported in the paper will be based on 595 unique country pairs, effectively covering the period 2000Q1-2014Q4.

[8]Our sample covers the following borrowing countries: Australia, Belgium, Canada, Chile, China, Colombia, Czech Republic, Denmark, Finland, France, Germany, Greece, Hungary, Indonesia, Ireland, Israel, Italy, Japan, Korea, Luxembourg, Mexico, Netherlands, Norway, New Zealand, Portugal, Russia, Slovakia, Slovenia, South Africa, Spain, Sweden, Switzerland, Turkey, UK, United States. For further details on the OECD house price index refer to appendix (A.1).

Bilateral treasury bases: To measure λ_t^b, the bilateral treasury basis, we use the "covered interest parity deviations between government bonds" data compiled by Du and Schreger (2016); Du et al. (2018b) and updated in January 2020 (v2). These data are kindly made available on Jesse Schreger's website.[9] We use the bilateral treasury basis at the five-year tenor as our baseline measure to reflect the exposure to dollar funding conditions relevant to the typically longer maturities of mortgage lending and housing markets which are our focus here. The five-year horizon also lines up with the horizon at which we measure house price and mortgage credit growth comovements as discussed in the previous paragraph.

Lending banking systems' market shares in borrowing countries: The market shares $\omega^{i,b}$ and $\omega^{j,b}$ of lending banking systems are essential inputs for the empirical counterparts of the co-Herfindahl index $\text{CoHFI}^{i,j}$ and the dollar co-dependence $\text{CoDD}^{i,j}$. We compute these market shares based on bilateral positions of outstanding foreign claims recorded in the consolidated banking statistics (CBS) on immediate counterparty basis, maintained as part of the international banking statistics (IBS) by the BIS.[10] The bilateral CBS statistics are confidential.

The CBS provide a uniquely suitable database to capture the network structure of lending banking systems' foreign claims as it records banking groups' consolidated "foreign claims". "Foreign" refers to the fact that these claims capture international credit by banks that are headquartered in a country other than the borrowing country, i.e. banks that are of foreign nationality, irrespective of whether this credit is cross-border or extended by a local subsidiary or branch. A consolidated view of international bank lending is most suitable to our research question, as US dollar funding conditions affect a banking group as a whole, regardless of the location of its offices. Internationally active banking groups obtain US dollar funding through various channels — notably deposits, debt securities issuance, wholesale funding, FX derivatives — and from various locations (Aldasoro and Ehlers (2018)). Moreover, they actively shift US dollar funds across offices in different locations (Cetorelli and Goldberg (2012)). The CBS record bank claims at a group level and thus abstract from interoffice positions that mainly reflect the internal shifting of funds within a banking

[9] https://sites.google.com/view/jschreger/CIP

[10] Foreign claims in the BIS terminology are the sum of international credit and local credit in local currency. International credit is defined as the sum of cross-border credit in both local and foreign currency and local credit in foreign currency. Local credit is defined as credit extended by a foreign banking group's affiliates located in the borrowing country itself.

group. Foreign claims reflect the full foreign credit exposure of a bank, as they not only comprise loans, but also debt securities holdings and net derivative exposures. We use data on the bilateral country-level claims of 28 lending banking systems on the 35 borrowing countries in our sample.[11]

5 Main empirical results

We first establish that US dollar funding conditions affect house prices globally and that the strength of this effect depends on the dollar dependence of borrowing countries. To this end, we take equation (5) on first conditional moments of house price growth to the data, and run borrowing country-level regressions. In a second step, we take equation (6) to the data, to show at the country-pair level that dollar co-dependence translates into (time-varying) house-price synchronization across borrowing countries.

5.1 Country-level evidence: house price growth, dollar dependence and dollar funding conditions

We test equation (5) by running the following panel regression:

$$\text{HPgrowth}_t^i = \boldsymbol{\beta}' \textbf{DF}_t \times \text{DD}_{t-1}^i + \text{CONTROLS}_t^i + \xi_t^i \tag{9}$$

where HPgrowth_t^i is the rate of house price growth over four quarters ahead in borrowing country i, DD_{t-1}^i is country i's dollar dependence as defined in section (2), and $\textbf{DF}_t$ denotes a broad range of variables that could potentially drive US dollar funding conditions: i) the dollar factor, i.e. four-quarter changes in the real effective exchange rate of the US dollar, as shown by Avdjiev et al. (2019) to be an important driver of cross-border investment and also suggested by our theoretical model in appendix B, ii) the four-quarter change in the US federal funds rate to account for changes in the stance of US monetary policy, and, iii) net treasury flows into the United States. As for iii), Krishnamurthy and Lustig (2019) show that treasury inflows drive the multilateral US treasury basis, an important reference for US dollar funding conditions. Hoffmann and Stewen (2020) have shown that capital inflows into US safe assets can be interpreted as a positive liquidity supply

[11] For further details on the computation of the market shares, please refer to appendix (A.2). Appendix (A.3) provides a detailed view on the suitability of the CBS.

shock that improves refinancing conditions and relaxes leverage constraints for banks borrowing in the US money market. We further include as measures of the global financial cycle iv) a measure of US broker dealer leverage, and v), the VIX as an index of global investor sentiment (Rey (2015)).

The vector β' contains our coefficients of interest with signs such that an improvement in US dollar funding conditions loosens non-US banks' balance sheet capacity, increases cross-border capital flows into foreign mortgage markets, and increases house prices in borrowing countries. The vector of controls CONTROLS contains the stand-alone term DD^i_{t-1} and a range of fixed effects. These include borrowing country as well as time fixed effects. We also allow the impact of the time fixed effect to vary across borrowing countries depending on the market share of US banks in the respective country. This controls for potentially heterogeneous confounding effects of global or US factors on borrowing countries and ensures the validity of our shift-share design.[12] We also include GDP growth defined over the same horizon as house price growth to make sure that our findings for house prices do not just reflect the influence of business cycles. Note that the stand alone term of the vector $\mathbf{DF}_t$ is absorbed by the time fixed effects. We cluster standard errors by both the country and time dimension to account for the possible correlation of residuals across borrowing countries at each point in time as well as within borrowing countries over time.

Panel A (columns 1-6) of Table 1 shows the results for regression (9). The results for the individual factors in columns (1) to (5) are in line with expectations. Consistent with our mechanism, an appreciation of the dollar lowers house price growth more for more dollar dependent economies. Higher capital inflows and broker-dealer leverage lead to higher house price growth while an increase in the federal funds rate lowers house prices. The VIX is not individually significant. When we consider all factors jointly in column (6), the dollar factor and broker-dealer leverage retain their significance and the associated coefficients remain stable relative to the specifications in the previous columns. Again this is consistent with our model in which leverage and changes in the dollar exchange rates are the key determinants of bank lending and thus house price growth.

The estimated coefficients also suggest that the effect of the dollar on house prices is not only statistically significant but also economically important. The standard deviation of DD across all

[12] The dollar-dependence DD can be interpreted as a shift-share variable, in which, however, one of the shocks is non-randomly assigned because the treasury basis of the dollar with itself is identically zero. Borusyak et al. (2022) show that in order to ensure the validity of the shift-share design, in such cases it is important to control for a time fixed effect interacted with the share of non-randomly assigned shocks, as we do here.

countries and across the entire sample is 8 basis points. Our estimate of -1.45 therefore implies that after a 10 percent dollar appreciation, house prices drop by around 1.2 percent more (relative to the average borrowing country) in a borrowing country that has one standard deviation higher dollar dependence.

Our mechanism predicts that dollar funding conditions affect house price growth through mortgage lending growth in the borrowing economies. As mortgage growth constitutes the bulk of household credit, we therefore look at household credit growth instead of house price growth as the dependent variable in regression (9).[13] The results in Panel B (columns 7-12) of Table 1 are consistent with this mechanism in that they show the same pattern we have documented for house prices: mortgage growth is more sensitive to fluctuations in dollar funding conditions in more dollar-dependent borrowing countries. Also, the estimated coefficient on the dollar factor is quantitatively in line with the previous estimates for house prices and economically meaningful. At -1.0, it implies that a borrowing country with dollar dependence one standard deviation above the mean will have 0.8 percent lower mortgage growth after a 10 pecent dollar depreciation.

5.2 House price synchronization and dollar co-dependence

In the next step, we explore the implications of our framework for house price synchronization. We translate equation (6) from the theoretical setup into the following panel regression

$$\text{HPcorr}_t^{i,j} = \beta \times \text{CoDD}_{t-1}^{i,j} + \delta \times \text{CoHFI}_{t-1}^{i,j} + \text{CONTROLS}_t^{i,j} + \theta_{i,j} + \mu_t^i + \delta_t^j + \epsilon_t^{i,j} \tag{10}$$

where $\text{HPcorr}_t^{i,j}$ denotes the conditional correlation of house price growth between borrowing countries i and j. We compute $\text{HPcorr}_t^{i,j}$ using a forward rolling window of 16 quarters from period t. Our coefficient of interest is the one on the dollar co-dependence term, $\text{CoDD}_{t-1}^{i,j}$. This coefficient β should be unambiguously positively signed as an increase in the dollar co-dependence implies that borrowing countries i and j are simultaneously more exposed to their lender banks' reaction to fluctuations in US dollar funding conditions, strengthening the link between US dollar funding conditions and the international synchronization of house price growth. The second term is again

[13] Note that mortgage lending constitutes about 90 percent of household lending in the borrowing countries of our sample. The literature on the effect of capital inflows on house prices provides further evidence for the central role of the domestic banking sector in translating capital inflows into mortgage credit (Aizenman and Jinjarak (2009), Sá et al. (2014), and Hoffmann and Stewen (2020)).

the co-Herfindahl index $\text{CoHFI}_{t-1}^{i,j}$ and captures the common lender effect adapted from Landier et al. (2017).

The vector $\text{CONTROLS}_t^{i,j}$ comprises variables controlling for the time-varying bilateral integration between borrowing countries i and j, notably bilateral trade integration to control for demand-driven house price co-movement generated by bilateral trade. We also control for time-varying bilateral output growth correlations to ensure that our results are not driven by a correlation in business cycles.

Equation (10) is saturated with a full set of fixed effects which results in a demanding specification that allows us to control for most conceivable confounders, strengthening the causal interpretation of our results. Specifically, the pairwise panel structure of the data allows us to control for observed or unobserved time-invariant country-pair specific variation which gets absorbed by the country-pair fixed effect θ^{ij}. Furthermore, any time-varying country-i or country-j specific shocks — including any country-specific demand- or supply shocks for housing and foreign-funded credit— are controlled for by saturating the regression with country-time fixed effects μ_t^i and δ_t^j. These country-time fixed effects also absorb all nuisance terms that arise in the log-linearization underlying equations (8) and its empirical counterpart (10).

Table 2 shows our estimates of equation (10). The coefficient on the dollar co-dependence is positive and statistically significant at the 1 percent significance level and stable across specifications, in line with the theoretical prediction that a higher dollar co-dependence strengthens the link between the variation in US dollar funding conditions and the synchronization of house prices. The standard deviation of CoDD across all periods and country pairs is around 0.07 so that the estimate of the coefficient on CoDD of 1.76 implies that a one standard deviation increase in dollar co-dependence raises the bilateral correlation in house price growth for a given country pair by about 12 percentage points.

The effect of dollar co-dependence on house price synchronization is economically sizable. In contrast, we do not find that the transmission of lender-banking system specific shocks —i.e. traditional common lender effects—have a measurable impact on house price synchronization in our international context. The coefficient estimate for δ on the co-Herfindahl index $\text{CoHFI}_t^{i,j}$ is an order of magnitude smaller than our estimate of β and insignificant throughout.

We again find the exact same patterns for mortgage credit. In Table 3, we use mortgage credit synchronization as the dependent variable. The construction of mortgage credit synchronization follows that of house price synchronization, applying a 16-quarter-ahead rolling-window correlation. Dollar co-dependence is strongly significant in all specifications. The estimated coefficient of around 0.64 implies that an increase in CoDD of around one standard deviation (0.07) increases the bilateral correlation between mortgage growth rates by around 4.5 percentage points.

5.3 Addressing reverse causality: a granular IV approach

Our results show that house prices of country pairs co-dependent on dollar financing conditions tend to co-move. Our specifications contain a full set of fixed effects which allow us to rule out that unobserved time-invariant country-pair specific, or time-varying country-specific shocks could drive our results.

However, there is a potentially large global component in countries' exposure to dollar funding conditions which could affect house price synchronization in borrowing countries through other channels than non-US banks.[14] This could lead to reverse causality in the synchronization regression (10). Assume that some global factor affects the bilateral treasury basis so that

$$\lambda_t^b = f_t + u_t^b \tag{11}$$

and the same factor affects house price synchronization in borrowing country pair i, j so that the residual in equation (10) is of the form

$$\epsilon_t^{i,j} = \psi^{i,j} f_t + \nu_t^{i,j}$$

Then $\epsilon_t^{i,j}$ will be correlated with $\text{CoDD}_{t-s}^{i,j}$ and OLS estimates of $\alpha\sigma_\zeta^2$ would be biased.[15] In this setting, our OLS estimations would suggest that global dollar funding shocks affect borrowing-country outcomes through the differential exposure of lending banking systems (λ_t^b) while in reality it is global variation in this exposure (f_t) that drives the global transmission of dollar funding

[14] In fact, the bilateral treasury basis we use to measure lender banks' sensitivity to dollar funding shocks is known to have a large common component—the multilateral basis, defined as the equal-weighted average of bilateral treasury base (Krishnamurthy and Lustig (2019)).

[15] Note that OLS is biased only if the loading $\psi^{i,j}$ is country-pair specific. If the loading was country-specific only, such that $\epsilon_t^{i,j} = \psi^i f_t + \psi^j f_t + \nu_t^{i,j}$, the confounding effects of f_t would already be absorbed by the country i- time and country j-time effects in (10) above.

shocks—possibly through entirely different channels than the lending of non-US banks.

To address this issue, we propose to adapt the granular instrumental variable technique recently proposed by Gabaix and Koijen (2024) to study international comovement. In so doing, we also extend the approach of Landier et al. (2017) to settings in which no quasi-experimental exogenous institutional change is readily available as an instrument.[16] Applying the granular instrumental variable approach to the study of synchronization between economic variables constitutes a methodological contribution of our paper.

Suppose we know the residuals u_t^b of the factor structure (11) above. Then we can construct the following granular instrumental variable for CoDD:

$$\mathcal{G}_{t-1}^{\text{CoDD}} = \left(\sum_{b\in\mathcal{B}(i)} \Gamma_{t-1}^{i,b} u_{t-1}^b\right)\left(\sum_{b\in\mathcal{B}(j)} \Gamma_{t-1}^{j,b} u_{t-1}^b\right) = \mathcal{DD}_{t-1}^i \times \mathcal{DD}_{t-1}^j \tag{12}$$

where $\left\{\Gamma_{t-1}^{i,b}\right\}$is a set of weights to be defined below. We call $\mathcal{DD}_{t-1}^i = \sum_{b\in\mathcal{B}(i)} \Gamma_{t-1}^{i,b} u_{t-1}^b$ the granular dollar dependence and $\mathcal{G}^{\text{CoDD}}$ the granular co-dependence. $\mathcal{DD}_{t-1}^i$ is uncorrelated with f_t by construction while being correlated with λ_{t-1}^b via the residual u_{t-1}^b. This makes $\mathcal{G}_{t-1}^{\text{CoDD}}$ a valid instrument for CoDD in our main regression (10).

For the factor structure of λ_t^b given in (11), where loadings are the same across different b, the variable $\mathcal{DD}_{t-1}^i$ can be constructed without having to estimate the individual u_{t-1}^b by choosing an appropriate set of weights $\Gamma_{t-1}^{i,b}$. To see this, define

$$\Gamma_{t-1}^{i,b} = \omega_{t-1}^{i,b} - \frac{1}{\#\mathcal{B}(i)}$$

where $\#\mathcal{B}(i)$ is the number of lender banks active in borrowing country i. Then

$$\mathcal{DD}_{t-1}^i = \underbrace{\sum_{b\in\mathcal{B}(i)}^{N} \Gamma_{t-1}^{i,b} f_{t-1}}_{=0} + \sum_{b\in\mathcal{B}(i)}^{N} \Gamma_{t-1}^{i,b} u_{t-1} \tag{13}$$

where the first term is zero since $f_{t-1} = \sum_{b\in\mathcal{B}(i)}^{N} \omega_{t-1}^{i,b} f_{t-1} = \frac{1}{N}\sum_{b\in\mathcal{B}(i)}^{N} f_{t-1}$. Hence, in the case of homogeneous loadings, a valid instrument can be constructed as the difference between the

[16] Landier et al. (2017) exploit the quasi-natural experiment of state-level banking deregulation in the US as an instrument.

market share-weighted (defined as in 1) and the equally-weighted dollar dependence (defined as $\text{DD}^{E,i}_{t-1} = \sum_{b\in\mathcal{B}(i)}^{N} \frac{\lambda^b_{t-1}}{\#\mathcal{B}(i)}$).

We construct $\mathcal{DD}^i_{t-1}$ according to (13), compute $\mathcal{G}^{\text{CoDD}} = \mathcal{DD}^i_{t-1} \times \mathcal{DD}^j_{t-1}$ and then use $\mathcal{G}^{\text{CoDD}}$ as an instrument for CoDD in the synchronization regression (10). Columns (1) to (3) of Table 4 report the results. The instrument is very strong as shown by the high first stage F-statistics. The estimated second-stage coefficient is significant and numerically very similar to the one obtained from the OLS regressions in Table 2. These findings allow us to rule out that global variation in borrowing countries' exposure to dollar funding conditions drives our results. Rather, global house price synchronization seems to be driven by the purely lender bank-specific component of the exposure to dollar funding shocks.

However, it could still be the case that some borrowing country-group specific factors feed back on the dollar funding conditions faced by some of their lender banks and thus on these lenders' bilateral treasury basis λ^b_t.[17] That is, factors with heterogeneous impact on various groups of borrowing countries and lenders could lead to biased results. Therefore, we allow the bilateral treasury basis to follow a more general factor structure of the form

$$\lambda^b_t = \sum_{r=1..R} \phi^{b,r}_t f^r_t + u^b_t \tag{14}$$

where the f^r_t for $r = 1...R$ is a set of unobserved ("regional") factors that could affect the synchronization between borrowing countries, such that the residual of the synchronization regression is given by $\epsilon^{i,j}_t = \sum_r \psi^{i,j}_r f^r_t + \nu^{i,j}_t$. The loading coefficients $\phi^{b,i}_t$ then capture the potential spillbacks of these factors on the lender banks' treasury bases and thus their exposure to dollar shocks.

Differently from the factor structure (11) above, the loadings $\phi^{b,r}_t$ can differ across lender banks b, which implies that $\mathcal{DD}^i_t$ cannot simply be constructed by a judicious choice of weights $\Gamma^{i,b}_{t-1}$. Instead, we have to estimate the individual u^b_t directly.

To obtain these estimates of u^b_t, following Gabaix and Koijen (2024), we could jointly estimate the loadings $\phi^{b,r}_t$ and the factors f^r_t using some atheoretical factor-analytical technique and then ex-

[17] For example, assuming strong regional concentration in lender banks' foreign lending, a regional economic slump could spill back to some geographically close lender banks, adversely affecting lenders' cost of borrowing in US dollars, thus driving up their bilateral treasury basis. The label "region" is a catch-all term for a characteristic shared by borrowing countries grouped together based on that common characteristic. For instance, this could be countries belonging to the same free trade agreement.

tract the residuals u_t^b. Alternatively, we could employ some economic theory to proxy the loadings $\phi_t^{\prime b,r}$ in terms of observable lender bank-specific characteristics. This would allows us to estimate the common factors by OLS as the series of coefficients on the interaction between $\phi_t^{\prime b,r}$ and a time-t-country-group-r dummy. The residual of this regression would then provide us with estimates for u_t^b. This theory-based approach has the advantage that it allows for a direct economic interpretation of the unobserved factors. This is the approach we take here.

Specifically, we suggest to interpret the f_t^r as geographical factors and it therefore seems natural to interpret the $\phi_t^{\prime b,r}$ as the share of region r in the international portfolio of lender country banking system b. Geographical proximity is known to be a good proxy for trade linkages between borrowing countries as well as for similarities in their industrial structure and plausibly for many other uncontrolled or unmodelled similarities between borrowing countries. We would also expect the impact of some regional factor f_t^r on lender banks b to increase with the exposure of b to the respective region. Our data set puts us in a unique position to calculate the portfolio shares $\phi_t^{\prime b,r}$ for each lender bank. In turn, this allows us to directly estimate the regional factors by OLS. We do so for a set of four geographical factors: besides the homogeneous global factor, we consider separate factors for advanced economies within and outside the euro area as well as for central and eastern Europe, respectively. We also allow for a lender bank-specific mean in the estimation of (14) in order to rule out that our results are driven by time-invariant unobserved characteristics of lender banks. We use the residuals of this model with multiple regional factors to construct $\mathcal{G}_{t-1}^{\text{CoDD}}$ according to (12), using our original market share weights, i.e. $\Gamma_{t-1}^{i,b} = \omega_{t-1}^{i,b}$.

Columns (4) to (6) of Table 4 show IV results for our main house price synchronization regression (10), with the instrument $\mathcal{G}_{t-1}^{\text{CoDD}}$ now constructed based on this model with multiple regional factors. Note first that that $\mathcal{G}_{t-1}^{\text{CoDD}}$ proves again a very strong instrument for CoDD. In all specifications the first stage F-statistics at the bottom of Table 4 remain far above the usual critical value of around 10. All our previous conclusions remain intact. Although the second-stage coefficient on CoDD is somewhat less significant than before, it remains numerically stable across specifications. It is also very similar to the coefficients obtained from the OLS specifications in Table 2 and from the previous IV specifications in columns (1)-(3).

In Table D.2 we report IV regressions for the synchronization of mortgage growth. While the second stage results are significant only at the 10-percent level, the results again mirror those for house price synchronization.

6 Robustness: alternative measures for house price synchronization and the Treasury basis

We provide additional robustness checks in Tables D.3 and D.4. First, we examine if our results hold up for alternative measures of synchronization. While our main results are based on correlations, we also re-run our synchronization regression (10) with pairwise covariances and on "pairwise average betas" as dependent variables. Following Landier et al. (2017), we construct the "pairwise average beta" as the mean of the (rolling-window) regression coefficients of house prices or mortgage growth in country i (j) on house price or mortgage growth in country j (i). Our results in Table D.3 remain largely unaffected.

Second, we examine the robustness of our conclusions with respect to treasury bases calculated at different maturities (tenors). While our baseline results are for 5-year tenors, we report results for the 1-year, 3-year and 10-year tenors in Table D.4. Again our results remain robust even though they are a little weaker at shorter maturities. This is to be expected because longer-term rates are likely to be more relevant for housing markets.

7 Conclusions

We document the role of non-US global banks in synchronizing house prices across countries. Because non-US global banks finance their cross-border lending largely in US dollars, variations in US dollar funding conditions induce an international synchronization of foreign lending and thus of mortgage credit growth and house price growth in borrowing countries. We show empirically and theoretically that the bilateral treasury basis between the currency of the non-US global creditor banks' headquarters and the US dollar represents non-US global banks' exposure to US dollar funding conditions. For each borrowing country, we construct a measure of dollar dependence as the weighted average of the treasury bases of its lenders and we show that house prices in countries with higher dollar dependence fall (rise) more as the dollar appreciates (depreciates).

For each borrowing country pair, we construct a measure of the joint exposure to US dollar funding conditions as the product of the individual countries' dollar dependence. We refer to this joint exposure as dollar co-dependence. Borrowing country pairs with higher dollar co-dependence exhibit higher house price synchronization, even after controlling for common-lender exposures. Our results identify a novel international spillover channel of US dollar funding conditions and shed new light on the globalization of real estate markets.

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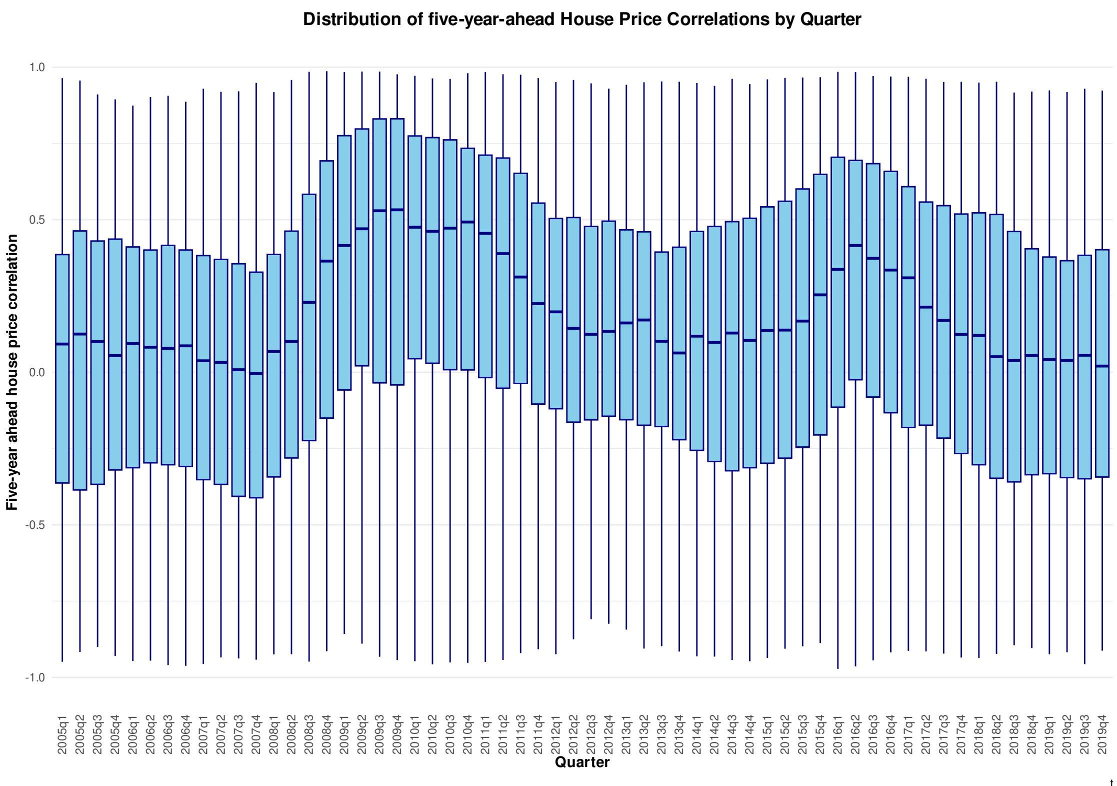
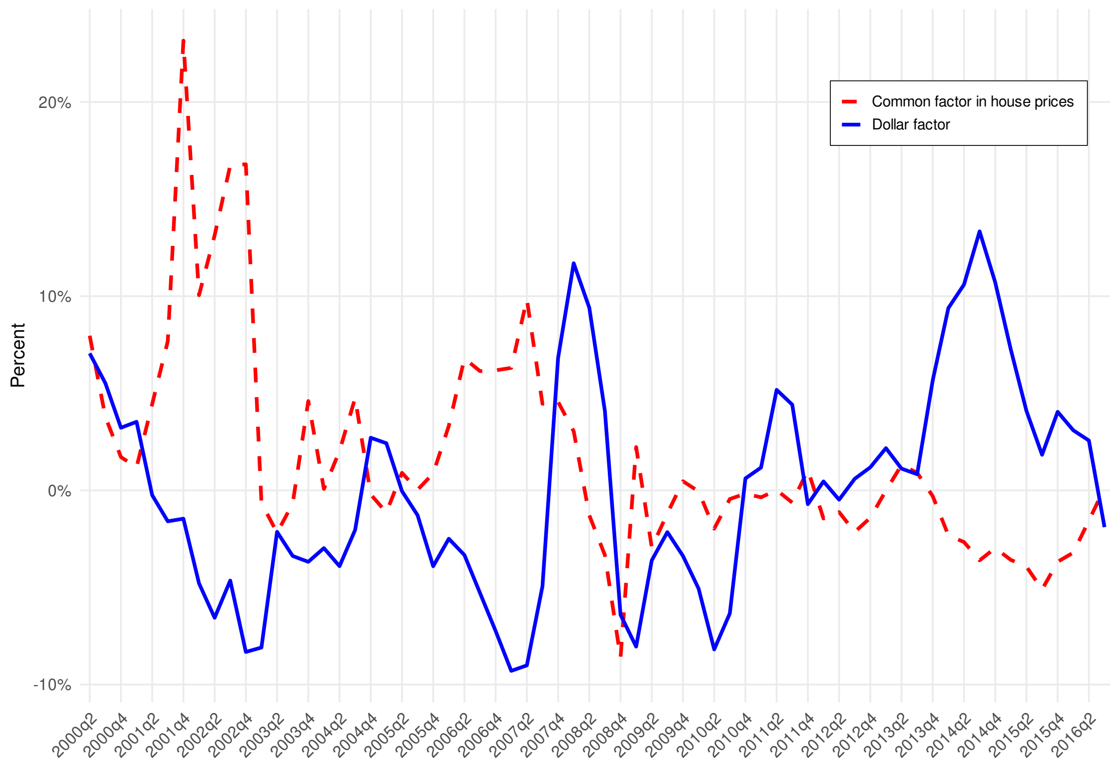
Figure 1: **House price synchronization, 2000Q1-2015Q1**



Note: This figure exhibits a box plot of quarterly pairwise international house price growth correlations over 595 country pairs for the period 2005Q1-2019Q4. The thick blue bars indicate the interquartile range. House price correlations of four-quarter-ahead house price growth rates are computed over a 16-quarter window ending in the quarter indicated on the horizontal axis. House price growth is calculated based on the country-wide residential real house price indices obtained for 35 countries from the OECD.

Figure 2: The dollar factor, dollar dependence, and house price growth across countries

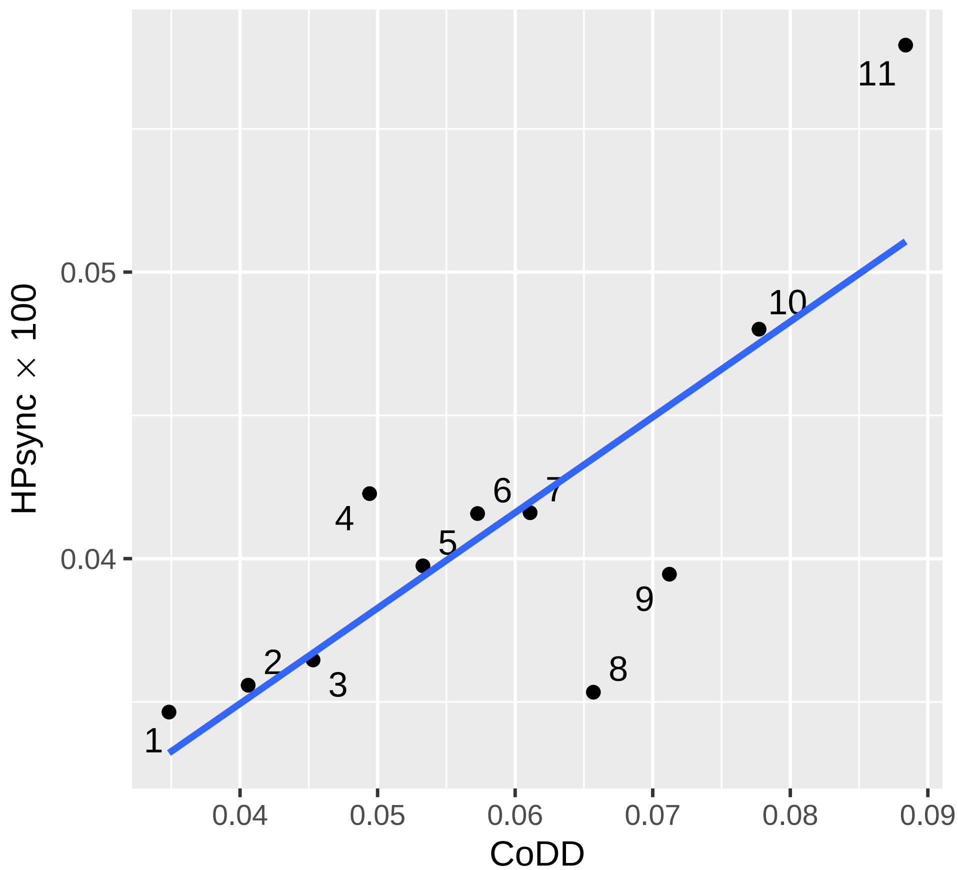


Note: the figure plots the coefficients ζ_t from the series of cross-sectional regressions

$$\Delta\text{HP}_t^i = \zeta_t \times \text{DD}_{t-1}^i + \text{constant}_t + \varepsilon_t^i$$

where $\Delta\text{HP}_t^i$is the four-quarter ahead house price growth in country i in period t and DD_{t-1}^i is dollar dependence in country i and period t (red, dashed line) against the dollar factor, i.e. the four-quarter ahead percentage change in the effective US nominal exchange rate (blue solid line). The estimated $\{\zeta_t\}$ have been rescaled to match the standard deviation of the dollar factor. The sample comprises 35 borrower countries over the period 2000Q1-2016Q3 (see main text for details).

Figure 3: House price synchronization and dollar co-dependence across country-pair portfolios



Note: the figure plots the average dollar co-dependence and the average house price synchronization for 11 portfolios formed from our 595 country pairs. The portfolios are re-sorted each quarter based on their dollar co-dependence. To control for outliers, in each period the highest and lowest 2 percent of observations are dropped from the sort. Each dot represents a portfolio and portfolios are numbered by ascending dollar co-dependence. The sample period is 2000-2016. House price synchronization for each country pair in each quarter is computed as the 5-year ahead covariance of house price growth and multiplied with 100.
The cross-sectional regression line in blue has slope 0.33 and a t-stat of 3.95.

Table 1: **House price growth, mortgage growth, and US dollar funding conditions**

	Panel A: house price growth						Panel B: mortgage growth					
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Variables												
$\text{DD}^i_{t-1}\times$ dollar factor	-1.45***					-1.03***	-1.00**					-0.788*
	(-4.27)					(-18.6)	(-2.34)					(-2.02)
$\text{DD}^i_{t-1}\times$ fed. fundsrate		-0.059***				-0.036		0.007				0.060
		(-2.75)				(-1.39)		(0.181)				(1.30)
$\text{DD}^i_{t-1}\times$ treasury inflows			1.53			-1.75			4.20			-2.13
			(0.883)			(-1.21)			(1.27)			(-0.910)
$\text{DD}^i_{t-1}\times$broker dealer lev.				0.010***		0.009**				0.011***		0.019**
				(3.13)		(2.08)				(2.99)		(2.61)
$\text{DD}^i_{t-1}\times$ VIX					-0.001	-0.003					-0.001	-0.002
					(-0.548)	(-1.37)					(-0.353)	(-0.557)
DD^i_{t-1}	0.049	0.044	0.039	-0.150**	0.074	-0.047	0.027	0.026	0.005	-0.189*	0.054	-0.279**
	(1.27)	(1.18)	(0.965)	(-2.44)	(1.18)	(-0.614)	(0.420)	(0.405)	(0.058)	(-1.93)	(0.406)	(-2.74)
GDP growth	0.971***	0.976***	0.972***	0.962***	0.976***	0.964***	1.57**	1.57**	1.56**	1.57**	1.57**	1.55**
	(5.24)	(5.21)	(5.18)	(5.24)	(5.22)	(5.23)	(2.39)	(2.40)	(2.38)	(2.38)	(2.40)	(2.38)
Fixed-effects												
country	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
date $\times\omega^{US,i}_{t-1}$	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Fit statistics												
Observations	2,032	2,032	2,032	2,032	2,032	2,032	1,983	1,983	1,983	1,983	1,983	1,983
R^2	0.49727	0.50141	0.49560	0.50127	0.49560	0.50554	0.50425	0.50398	0.50449	0.50629	0.50404	0.50822
Within R^2	0.14650	0.15352	0.14366	0.15328	0.14366	0.16052	0.13310	0.13263	0.13352	0.13667	0.13272	0.14003

Note: This table reports the results from estimating equation (9) for our panel of borrowing countries for the period from 2000Q1 to 2020Q4. In Panel A (columns 1-6), the dependent variable is the four-quarter-ahead growth rate of house prices, in Panel B (columns 7-12) the four-quarter-ahead growth rate of total consumer credit (of which around 90 percent are mortgages). The explanatory variables are country i's US dollar dependence DD^i_{t-1} lagged by one quarter, four-quarter-ahead GDP growth, and the interaction of the lagged US dollar dependence with the following common factors: the federal funds rate, the change in the effective US dollar exchange rate, net purchases by foreigners of US treasury securities normalized with foreign holdings of US treasuries, broker-dealer leverage and the VIX. All specifications include country and time fixed effects. Time effects are interacted with the US market share $\omega^{US,i}_{t-1}$ of borrowing country i. Standard errors are clustered by borrowing-country and quarter and t-statistics are shown in parentheses. *, ** and *** indicate statistical significance at the 10%, 5% and 1% level, respectively.

Table 2: **House price growth synchronization and dollar co-dependence**

Dependent Variable:		$\text{HPcorr}_t^{i,j}$	
	(1)	(2)	(3)
Variables			
$\text{CoDD}_t^{i,j}$	1.76^{***}	1.76^{***}	1.76^{***}
	(3.26)	(3.48)	(3.49)
$\text{CoHFI}_t^{i,j}$	0.175	0.174	0.176
	(0.328)	(0.311)	(0.316)
GDP growth corr.		0.154^{***}	0.154^{***}
		(3.14)	(3.14)
trade integration			-5.31
			(-0.238)
Fixed-effects			
CountryPair	Yes	Yes	Yes
country1-date	Yes	Yes	Yes
country2-date	Yes	Yes	Yes
Fit statistics			
Observations	27,767	26,894	26,894
R^2	0.53560	0.54960	0.54962
Within R^2	0.00150	0.00820	0.00823

Note: This table reports the results from estimating equation (10) for the period from 2000Q1 to 2020Q4. The dependent variable $\text{HPcorr}_t^{i,j}$ is the five-year ahead rolling window correlation of HPgrowth in countries i and j. The explanatory variables are US dollar co-dependence $\text{CoDD}_t^{i,j}$ and the co-Herfindahl index $\text{CoHFI}_{t-s}^{i,j}$. Standard errors are clustered two-way, by country i and country j, t-statistics are shown in parentheses. *, ** and *** indicate statistical significance at the 10%, 5% and 1% level, respectively.

Table 3: **Mortgage credit growth synchronization and dollar co-dependence**

Dependent Variable:	correlation of mortgage growth		
	(1)	(2)	(3)
Variables			
$\text{CoDD}_t^{i,j}$	0.433^{***}	0.644^{***}	0.643^{***}
	(7.39)	(17.4)	(4.64)
$\text{CoHFI}_t^{i,j}$	0.012	-0.191	-0.186
	(0.024)	(-0.414)	(-0.410)
GDP growth corr.		0.198^{***}	0.197^{***}
		(2.80)	(2.81)
trade integration			-4.87
			(-0.229)
Fixed-effects			
CountryPair	Yes	Yes	Yes
country1-date	Yes	Yes	Yes
country2-date	Yes	Yes	Yes
Fit statistics			
Observations	27,224	25,691	25,691
R^2	0.52796	0.55259	0.55260
Within R^2	6.69×10^{-5}	0.00899	0.00901

Note: This table reports the results from estimating equation (10) for the period from 2000Q1 to 2020Q4, but with the correlation of mortgage growth as the dependent variable. The correlation is computed as the four-year ahead rolling window correlation of 4 quarter-ahead mortgage growth in countries i and j. The explanatory variables are US dollar co-dependence $\text{CoDD}_t^{i,j}$ and the co-Herfindahl index $\text{CoHFI}_{t-s}^{i,j}$. Standard errors are clustered two-way, by country i and country j, t-statistics are shown in parentheses. *, ** and *** indicate statistical significance at the 10%, 5% and 1% level, respectively.

Table 4: **Instrumental variables regressions for house price synchronization**

Dependent Variable:	$\text{HPcorr}_t^{i,j}$					
GIV constructed using:	single global factor			multiple regional factors		
	(1)	(2)	(3)	(4)	(5)	(6)
Variables						
$\text{CoDD}_t^{i,j}$	1.76***	1.72***	1.72***	2.82**	2.52**	2.51**
	(4.20)	(4.59)	(4.61)	(2.23)	(2.18)	(2.18)
CoHFI_t^{ij}	0.175	0.174	0.177	0.161	0.163	0.166
	(0.329)	(0.313)	(0.318)	(0.302)	(0.293)	(0.299)
GDP growth corr.		0.154***	0.154***		0.155***	0.155***
		(3.14)	(3.14)		(3.11)	(3.11)
trade integration			-5.32			-5.23
			(-0.238)			(-0.235)
Fixed-effects						
CountryPair	Yes	Yes	Yes	Yes	Yes	Yes
country1-date	Yes	Yes	Yes	Yes	Yes	Yes
country2-date	Yes	Yes	Yes	Yes	Yes	Yes
Fit statistics						
F-test (1st stage), CoDD	248,437.4	240,366.3	240,352.7	1,882.3	1,852.6	1,852.1

Note: This table reports IV results equation (10) for the period from 2000Q1 to 2020Q4 using the granular instrument $\mathcal{G}^{\text{CoDD}}$ defined in (12). The dependent variable $\text{HPcorr}_t^{i,j}$ is the five-year ahead rolling window correlation of HPgrowth in countries i and j. The explanatory variables are US dollar co-dependence $\text{CoDD}_t^{i,j}$ and the co-Herfindahl index $\text{CoHFI}_{t-s}^{i,j}$. Two versions of $\mathcal{G}^{\text{CoDD}}$ are used: a version taking account of a single, homogenous global factor, constructed as the difference between the market share-weighted and the equally weighted US dollar dependence. Second, a version in which the granular residuals u_{t-1}^b are estimated as the residuals of a model with several regional factors as discussed in section 5.3.
Standard errors are clustered two-way, by country i and country j, t-statistics are shown in parentheses. *, ** and *** indicate statistical significance at the 10%, 5% and 1% level, respectively. First-stage F-statistics are reported in the last row of the table.

A Appendix

Table A.1: Descriptive statistics for main data

variable	units	mean	std. dev	5%	50%	95%
house price growth	percent	2.38	6.97	-8.60	2.13	13.49
mortgage credit growth	percent	7.66	13.96	-4.38	4.84	26.95
dollar dependence (DD_t^i)	percent	0.21	0.15	0.01	0.19	0.46
dollar factor	log 4-quarter change	0.02	2.23	-3.59	0.14	3.18
Federal funds rate	4-quarter change	-0.32	1.43	-3.16	-0.02	1.96
broker-dealer leverage	multiples of equity	22.89	5.83	14.48	22.33	33.63
VIX	percent	20.48	8.07	12.56	19.17	33.02

A.1 OECD house price index

Our analysis is based on country-level residential house price indices provided by the OECD. This data source is particularly suitable since the underlying house price data feeding into the index construction are of comparable quality, abstracting from differences in the definitions of the types of dwellings. Moreover, the relative homogeneity of OECD member countries in terms of structural features of their economies and financial market developments is advantageous for our identification strategy as time fixed effects in the regression analysis eliminate many time-varying confounding factors relevant to this country group. In addition to actual OECD member countries, the house price indices are also available for Brazil, China, Russia and South Africa. The price indices of Estonia, Latvia and Lithuania are available but excluded due to a relatively short time series characterized by extreme variation. Our final sample consists of 34 borrowing countries.

A.2 Computation of market shares

To define the market shares $\omega_t^{i,b}$, we argue that a lender bank's share in a borrowing country's market for foreign credit, i.e. credit provided by all foreign lender banks, is a more appropriate choice than the share in the market for total credit, i.e. foreign credit plus domestic credit provided by borrowing country banks. In this paper, our focus is on the effect of the foreign credit supply from lender banks induced by fluctuations in US dollar funding conditions. To isolate the effect of foreign as opposed to domestic credit on house price growth, our identification strategy in equation (10) employs country-time fixed effects that eliminate borrowing country specific economic and financial market developments, including the growth in domestic credit provided by borrowing country banks independent from the funding obtained through foreign borrowing. This allows us to abstract from domestic credit conditions, and to work with market shares based on foreign credit. Moreover, taking into account domestic credit would merely scale down the market shares. However, the cross-sectional distribution over lender banks would stand largely unaffected by this scaling, because the dominant lender banks have a large market share in every borrowing country, regardless of whether the share is computed in terms of foreign or total credit. Moreover, potential shifts in the cross-sectional distribution of the market shares due to scaling are negligible as the market shares only serve as weights in borrowing countries' dollar dependence as defined in equation (1). More relevant to the identification strategy is the lender banks' heterogenous exposure to US dollar funding shifts as measured by λ_t^b.

A.3 Locational versus consolidated banking statistics

The computation of the market shares is based on lender banks' foreign claims from the CBS on immediate counterparty basis, as opposed to the locational banking statistics (LBS). A practical reason for using the CBS is the availability of bilateral lending data, i.e. from a banking system of given nationality to a borrowing country, for the entire time period of our sample. This data has only started to be available in the LBS since 2012Q1 — a time period too short to analyze house price cycles. In addition to the availability of bilateral data, there are three economic reasons for using the CBS.

First, the nationality of the lender bank coincides with the decision making unit of the bank (Takáts and Temesvary (2016)). This is particularly relevant for global banks at the core of our analysis since policies on leverage and foreign currency funding — such as from the US dollar money market — are decided at a bank's global headquarters. Consequently, a global bank's lending — including the lending by foreign offices in the borrowing country — is driven by factors better captured by nationality. Therefore, a borrowing country's exposure vis-à-vis the global bank's lending should also be measured based on consolidated claims.Second, the CBS exclude interoffice positions by construction. Consider a British bank that extends a loan to a borrower in Chile. The exposure between the Chilean borrower and the British bank does not include any intermediate interoffice transactions, such as for instance between the British bank and its subsidiary in Mexico and from the Mexican subsidiary to the borrower in Chile. By virtue of consolidation, the CBS records only an exposure of the British bank vis-à-vis a borrower in Chile. This logic also applies to "looking through" financial centers through which a significant share of international transactions are routed. Suppose a German bank lends to a borrower in Finland through its German subsidiary in Luxembourg. The LBS would count two cross-border transactions, from the German bank to its subsidiary in Luxembourg and from the subsidiary to the borrower in Finland. The CBS, however, establish a direct link between the German bank and its borrower in Finland.

Third, the CBS take into account the two principal transaction forms of foreign credit provision. Foreign banks can provide credit either cross-border or through a local office in the borrowing country. As discussed by Kerl and Niepmann (2015), the choice depends on the "efficiencies of countries' banking sectors, differences in the return on loans across countries, and impediments to foreign bank operations". As the consolidated view does not differentiate between these two channels, it accounts for the entirety of foreign claims.

B A value-at-risk model of international dollar-lending by non-US banks

We consider the problem of a non-US bank that can raise funds in non-US home currency at interest rate r or US dollars at interest rate r^*. Our model focuses on international lending and we simplify the setup by assuming that the bank lends abroad only in US dollars (i.e. we do not model domestic lending and assume that it does not do any cross-border lending in its home currency). The non-US bank can raise direct funding, for instance through wholesale markets (debt securities, certificate of deposits, repos) or by issuing dollar deposits. It can raise further funds in the home currency, which then have to be converted into US dollars at the current spot exchange rate X^S(measured in non-US home currency per US dollar, meaning an increase in X^s is a US dollar appreciation). In line with regulatory requirements in most jurisdictions, we require that the amount of the bank's US dollar lending that is funded in non-US home currency (the non-US home currency amount of which we denote with S) has to be fully hedged in forward/futures markets at a forward premium ∇. The bank then uses the direct and indirect dollar funding to lend in the US dollar market at a lending rate r^l. The total amount of US dollars lent is A.

The bank's problem.

The non-US bank's problem is to maximize its future (expected) equity E_{t+1} *in non-US home currency*, taking as given today's equity (E) (all "today" variables have no time index) and non-US home currency and US refinancing rates r and r^*, respectively. With the assumptions above, E_{t+1} evolves according to

$$\begin{aligned} E_{t+1} &= \left[(A - S/X^S)(1+r^l) - (A - S/X^S)(1+r^*)\right] X^S_{t+1} \\ &\quad + S\left(1+r^l\right)\frac{X^F}{X^S} - (S-E)(1+r) \end{aligned} \tag{B.1}$$

where X^F is the forward rate at which the bank sells is synthetic US dollar position S (measured in home currency units per dollar), r^l is the interest rate on dollar lending, $r^*$is the dollar money market rate and r the domestic deposit rate of the bank. The first row of this expression is the profit — expressed in non-US home currency at tomorrow's spot exchange rate X^S_{t+1}— the bank makes on its directly funded dollar position, $A - S/X^S$. The second row is the profit in home currency the bank makes on its synthetic dollar position S.

We can rewrite this law of motion in terms of the forward premium ∇ as follows

$$\begin{aligned} E_{t+1} &= \left[(A - S/X^S)(1+r^l) - (A - S/X^S)(1+r^*)\right] X^S_{t+1} \\ &\quad + \left(1+r^l\right)\left(1 + \underbrace{\frac{X^F - X^S}{X^S}}_{=:\nabla}\right) S - (S-E)(1+r) \end{aligned} \tag{B.2}$$

Note that as the forward premium ∇ declines (and thus, for given home and dollar refinancing rates r and r^*, the treasury basis $r - r^* - \nabla$ increases), the higher will be the cost of hedging and the lower will be the return on synthetic lending.

We can expand with AX^S (using the approximation $r^l \times \nabla \approx 0$) to obtain

$$
\begin{aligned}
E_{t+1} &\approx \left[\left(\left[(1+r^l)-(1+r^*)\right]\left(1-\frac{S}{AX^S}\right)\frac{X^S_{t+1}}{X^S}\right)+(1+r^l)s\right]AX^S-(S-E)(1+r)+\nabla S \\
&= \left[\left(r^l-r^*\right)\frac{X^S_{t+1}}{X^S}\underbrace{(1-s)}_{\text{direct \$-funding share.}}+(1+r^l)\underbrace{s}_{\text{synthetic \$-funding share.}}\right]\underbrace{AX^S}_{\text{\$-assets in domestic currency}} \\
&\quad -S(1+r-\nabla)+E(1+r) \\
&= \underbrace{\left[(1+r^l)\frac{X^S_{t+1}}{X^S}-\underbrace{\left((1-s)(1+r^*)\frac{X^S_{t+1}}{X^S}+s(1+r-\nabla)\right)}_{\text{funding costs}}\right]}_{\Pi_{t+1}:=\text{excess return on bank portfolio in home currency}}AX^S+E(1+r)
\end{aligned}
$$

where $s=S/\left(AX^S\right)$ is the share of synthetic US dollar funding, and $1-s$ is the direct funding share.

Without additional constrains, the bank's problem is unbounded. For a given positive expected excess return, $\mathbb{E}_t\Pi_{t+1}>0$, it is always possible to increase expected equity by taking on more debt. Of course, the bank gets riskier as it leverages up. So, in order to bound the bank's problem, we impose that the bank maintains a fixed default probability, i.e. it faces a value-at-risk (VaR) constraint.

Default occurs when $E_{t+1}\leq 0$. Hence, setting $E_{t+1}=0$ and rearranging, we obtain the following lower bound on Π_{t+1}

$$
\Pi^{min}_{t+1}\leq\frac{-E(1+r)}{AX^S}
$$

If $\Pi_{t+1}<\Pi^{min}_{t+1}$ the bank will fail. Solvency therefore requires that

$$
\Pi_{t+1}\geq\Pi^{min}_{t+1}
$$

with a given default probability α so that

$$
\text{Prob}\left(\Pi_{t+1}\geq\Pi^{min}_{t+1}\right)=1-\alpha
$$

Then, given the variance $\sigma^2$of Π_{t+1} we can find an appropriate distance to default Ψ such that

$$\mathbb{E}_t\Pi_{t+1} - \Psi\sigma_{\Pi} = \Pi_{t+1}^{min}$$

Plugging in for Π_{t+1}^{min} from above we obtain

$$\mathbb{E}_t\Pi_{t+1} - \Psi\sigma \leq \frac{-E(1+r)}{AX^S}$$

Hence, the VaR constraint imposes the following upper bound on the leverage of the bank's lending portfolio:

$$\begin{aligned} Leverage := \frac{AX^S}{E} \leq \frac{1+r}{\Psi\sigma - \mathbb{E}_t\Pi_{t+1}} \\ = \frac{1+r}{\Psi\sigma - \mathbb{E}_t \underbrace{\left[(1+r^l)\frac{X_{t+1}^S}{X^S} - \underbrace{\left((1-s)(1+r^*)\frac{X_{t+1}^S}{X^S} + s(1+r-\nabla) \right)}_{\text{funding costs}} \right]}} \end{aligned} \tag{B.3}$$

and this condition will hold with equality, since expected future equity is monotonically increasing in leverage.

Hence, the VaR constraint pins down the amount of US dollar assets (expressed in non-US home currency) AX^S that the bank can hold for a given σ_Π and $\mathbb{E}_t\Pi_{t+1}$(and a given initial equity E). The bank can influence this upper bound by choosing s. Maximizing leverage therefore amounts to minimizing the denominator of the upper bound, i.e. $\Psi\sigma - \mathbb{E}_t\Pi_{t+1}$over s. This is a standard mean-variance problem.

Let the expected appreciation of the US dollar exchange rate be

$$\mu = \mathbb{E}_t\left(\frac{X_{t+1}^S}{X^S}\right)$$

Then we can write (assuming that exchange rate volatility is the only source of risk, i.e. r^l is predetermined and therefore non-stochastic):

$$\sigma_{\Pi} = \Big((1 + r^l) - (1 - s)(1 + r^*)\Big)\,\sigma_x$$

Hence the first-order condition for minimizing the denominator $\Psi\sigma_{\Pi} - \mathbb{E}_t\Pi_{t+1}$ and thereby maximizing leverage under the VaR constraint w.r.t. s is

$$\Psi(1 + r^*)\sigma_x - (1 + r^*)(\mu - 1) + (r - r^* - \nabla) - s\nabla'(s) = 0$$

As in Ivashina et al. (2015)), we assume that the supply of hedging is not fully elastic, so that an increase in the hedging demand leads to an increase in the cost of hedging. This amounts to assuming that $\nabla'(s) < 0$.[18]

Rearranging then yields

$$s = \frac{(1 + r^*)\,[(\mu - 1) - \Psi\sigma_x] - (r - r^* - \nabla)}{|\nabla'(s)|}$$

as the implicit solution for s.[19] Note how the right hand-side of this expression is directly related to the treasury-basis, $r - r^* - \nabla$! A higher basis means a lower synthetic funding share s, and thereby a higher direct funding share. Recall that we find that countries with higher bilateral basis $r - r^* - \nabla$ are more exposed to dollar re-financing conditions through direct funding as opposed to synthetic funding. Our model here explains this empirical regularity.

Using the model

We now use the model to see what happens when US dollar refinancing conditions change. In so doing, we assume the optimal choice of s by the respective lender bank as given. We then look at two ways in which US dollar refinancing conditions could change. First, a change in the US dollar exchange rate and secondly a drop in the US dollar interest rate (potentially caused by capital

[18]Note that a decline in the forward premium ∇ lowers the return of a hedged position in home-currency terms (see equation (B.2)), making hedging mor expensive.. Ceteris paribus, this is equivalent to an increase in the treasury basis $r - r^* - \nabla$ (as in Ivashina et al. (2015)). We also assume that the second derivative $\nabla''(s) \leq 0$. This ensures that $\nabla(s)$ is weakly concave so that the second derivative, which is given by $-2\nabla'(s) - s\nabla''(s)$ is positive. Hence, the first-order condition defines a minimum of $\Psi\sigma_{\pi} - \mathbb{E}_t\Pi_{t+1}$ and thus a maximum for the leverage.

[19]Note that we assume a representative but atomistic bank that takes the effective hedging supply function and thus$\nabla'(s)$ as given. Clearly, in equilibrium ∇ and s will be jointly determined. This would require to explicitly model the risk-taking capacity of the arbitrageur who is the counterparty in the hedging trade, as discussed in Ivashina et al. (2015).

inflows into the US dollar market, as in Hoffmann and Stewen (2020)).

A dollar appreciation

Consider first what happens after a US dollar appreciation i.e. an increase in X^S. Note that in our solution for the banks' maximal leverage (B.3), changes in X^S matter only in as far as they affect the expected rate of appreciation $\mu = \mathbb{E}_t\left(\frac{X^S_{t+1}}{X^S}\right)$. For a given future exchange rate, an increase in X^Stherefore amounts to a decline μ. Note that

$$\begin{aligned}
\frac{d}{d\mu}\left(\frac{1+r}{\Psi\sigma - \mathbb{E}_t\Pi_{t+1}}\right) &= \frac{1+r}{\left(\Psi\sigma - \mathbb{E}_t\Pi_{t+1}\right)^2} \times \frac{d\mathbb{E}_t\left(\Pi_{t+1}\right)}{d\mu} \\
&= \frac{(1+r)\left(1+r^l-(1-s)(1+r^*)\right)}{\left(\Psi\sigma - \mathbb{E}_t\Pi_{t+1}\right)^2} \\
&= \frac{(1+r)\left(r^l - r^* + s(1+r^*)\right)}{\left(\Psi\sigma - \mathbb{E}_t\Pi_{t+1}\right)^2} > 0
\end{aligned}$$

If we assume that $r^l > r^*$, which is a necessary conditions for bank to make a profit on their dollar lending, then a decline in μ (e.g. due to a dollar appreciation), will lower banks' leverage. Given the bank's current equity, this is akin to a decline in lending.[20]

How does the treasury basis affect the response of leverage and thus international lending in US dollar? Note from above that the treasury basis and s are isomorphic: higher s implies a lower treasury basis and vice-versa. Hence, it is sufficient to show what happens to the response above when we change s. To this end, we first rewrite the above response as

$$\frac{d}{d\mu}\left(\frac{1+r}{\Psi\sigma - \mathbb{E}_t\Pi_{t+1}}\right) = Leverage \times \frac{\left(r^l - r^* + s(1+r^*)\right)}{\left(\Psi\sigma - \mathbb{E}_t\Pi_{t+1}\right)}$$

so that

$$\begin{aligned}
\frac{d^2}{d\mu ds}\left(\frac{1+r}{\Psi\sigma - \mathbb{E}_t\Pi_{t+1}}\right) = &\frac{dLeverage}{ds} \times \frac{\left(r^l - r^* + s(1+r^*)\right)}{\left(\Psi\sigma - \mathbb{E}_t\Pi_{t+1}\right)} \\
&- Leverage \times \frac{(1+r^*)\left(\Psi\sigma - \mathbb{E}_t\Pi_{t+1}\right) + \left(r^l - r^* + s(1+r^*)\right) \times \frac{d}{ds}\left(\Psi\sigma - \mathbb{E}_t\Pi_{t+1}\right)}{\left(\Psi\sigma - \mathbb{E}_t\Pi_{t+1}\right)^2}
\end{aligned}$$

[20]Because changes in X^S and $\mathbb{E}_t X^S_{t+1}$matter only in as far as they affect μ, the model also predicts that bank lending should only react to temporary exchange rate changes (i.e. changes today that leave future expected exchange rates unchanged or change them less than one to one). By contrast, changes in X^S that are expected to be permanent, (i.e. affect expected exchange rates to the same extent, so that $dX^S = d\mathbb{E}_t X^S_{t+1}$) should not affect bank lending.

This expression simplifies considerably once we realize that the bank has chosen s to maximize its leverage. So, the envelope theorem implies that

$$\frac{dLeverage}{ds} = d\left(\Psi\sigma - \mathbb{E}_t\Pi_{t+1}\right)/ds = 0$$

and we obtain

$$\frac{d^2}{d\mu ds}\left(\frac{1+r}{\Psi\sigma - \mathbb{E}_t\Pi_{t+1}}\right) = -Leverage \times \left[\frac{(1+r^*)}{(\Psi\sigma - \mathbb{E}_t\Pi_{t+1})}\right] < 0$$

which will always be negative. Because $d\mu/dX^S < 0$, this implies that a lower s (a higher Treasury basis) will be associated with a higher exposure to variations in the dollar exchange rate!

A drop in the US interest rate (e.g. following a positive capital inflow shock)

$$\begin{aligned}\frac{d}{dr^*}\left(\frac{1+r}{\Psi\sigma - \mathbb{E}_t\Pi_{t+1}}\right) &= -\frac{1+r}{\left(\Psi\sigma - \mathbb{E}_t\Pi_{t+1}\right)^2} \times \frac{d\left(\Psi\sigma - \mathbb{E}_t\Pi_{t+1}\right)}{dr^*} \\ &= -\frac{1+r}{\left(\Psi\sigma - \mathbb{E}_t\Pi_{t+1}\right)^2}\left[(1-s)\mu - \Psi(1-s)\sigma_X\right]\end{aligned}$$

which will be negative whenever $\mu > \Psi\sigma_X$. This will usually be the case because μ is a gross change ($\mu = \mathbb{E}_t\left(\frac{X^S_{t+1}}{X^S}\right) = \mathbb{E}_t\left(1 + \Delta\log(X_{t+1})\right)$ while $\sigma_X = \sigma\left(1 + \Delta\log(X_{t+1})\right) = \sigma(\Delta\log(X_{t+1})$ is the volatility of a growth rate. Empirically, the variance of growth rates of the exchange rate are small compared to "1+growth rate", so we can conclude that a decrease of the interest rates will increase leverage, as found in our empirical specifications.

Again we can ask what happens if we vary s. Again, first rewrite

$$\frac{d}{dr^*}\left(\frac{1+r}{\Psi\sigma - \mathbb{E}_t\Pi_{t+1}}\right) = -Leverage \times \frac{\left[(1-s)\mu - \Psi(1-s)\sigma_X\right]}{\left(\Psi\sigma - \mathbb{E}_t\Pi_{t+1}\right)}$$

and then

$$\begin{aligned}\frac{d}{dr^*ds}\left(\frac{1+r}{\Psi\sigma - \mathbb{E}_t\Pi_{t+1}}\right) =& -\frac{dLeverage}{ds} \times \frac{\left[(1-s)\mu - \Psi(1-s)\sigma_X\right]}{\left(\Psi\sigma - \mathbb{E}_t\Pi_{t+1}\right)} \\ &- Leverage \times \frac{\left(\Psi\sigma_X - \mu\right)\left(\Psi\sigma - \mathbb{E}_t\Pi_{t+1}\right) - [...]\frac{d}{ds}\left(\Psi\sigma - \mathbb{E}_t\Pi_{t+1}\right)}{\left(\Psi\sigma - \mathbb{E}_t\Pi_{t+1}\right)^2}\end{aligned}$$

Using the envelope theorem again, we obtain

$$\frac{d}{dr^*ds}\left(\frac{1+r}{\Psi\sigma - \mathbb{E}_t\Pi_{t+1}}\right) = -Leverage \times \frac{(\Psi\sigma_X - \mu)}{(\Psi\sigma - \mathbb{E}_t\Pi_{t+1})} > 0$$

which is positive whenever $\mu > \Psi\sigma_X$. Hence, a lower synthetic funding share makes the positive response of lending to a decline in interest rate stronger, again as found in the data.

C Log-linearizing the expression for house price correlations

Using equations (6) and (7) we can write the correlation of house price growth rates between countries i and j as

$$\begin{aligned}
\text{HPcorr}^{i,j}_{t-1} &:= \frac{\text{HPcov}^{i,j}_{t-1}}{\sigma\left(\frac{\Delta\text{HP}^i_t}{\text{HP}^i_{t-1}}\right) \times \sigma\left(\frac{\Delta\text{HP}^j_t}{\text{HP}^j_{t-1}}\right)} \\
&= \frac{\alpha^2\sigma^2_\gamma + \alpha^2\sigma^{2\,i,j}_{\eta\,t-1} + \alpha^2\sigma^2_\zeta\text{CoDD}^{i,j}_{t-1}}{\left(\sigma^2_\varepsilon + \alpha^2\sigma^2_\gamma + \alpha^2\sigma^2_\eta\text{CoHFI}^{i,i}_{t-1} + \alpha^2\sigma^2_\zeta\text{CoDD}^{i,i}_{t-1}\right)^{1/2}\left(\sigma^2_\varepsilon + \alpha^2\sigma^2_\gamma + \alpha^2\sigma^2_\eta\text{CoHFI}^{j,j}_{t-1} + \alpha^2\sigma^2_\zeta\text{CoDD}^{j,j}_{t-1}\right)^{1/2}}
\end{aligned}$$

We expand this expression around the reference point of two countries that only borrow from the United States. It is useful to briefly consider what this means for our setting. First, our dollar co-dependence mechanism is present only for non-US lender banks(because the treasury basis of the US with itself is zero), so that we have

$$\text{CoDD}^{i,i} = \text{CoDD}^{j,j} = \text{CoDD}^{i,j} = 0$$

Furthermore, for countries that draw all their borrowing from one country, the Herfindahl and the co-Herfindahl indexes that measure the concentration of their borrowing, will all be unity:

$$\text{CoHFI}^{i,i}_{t-1} = \text{CoHFI}^{j,j}_{t-1} = \text{CoHFI}^{i,j}_{t-1} = 1$$

Then a first-order expansion yields

$$\begin{aligned}
\mathrm{HPcorr}_{t-1}^{i,j} &= \frac{\alpha^2\left[\sigma_\gamma^2+\sigma_\eta^2\right]}{\sigma_\varepsilon^2+\alpha^2\left[\sigma_\gamma^2+\sigma_\eta^2\right]} + \frac{\alpha^2\sigma_\eta^2}{\sigma_\varepsilon^2+\alpha^2\left[\sigma_\gamma^2+\sigma_\eta^2\right]} \times \left(\mathrm{CoHFI}_{t-1}^{i,j}-1\right) + \frac{\alpha^2\sigma_\zeta^2}{\sigma_\varepsilon^2+\alpha^2\left[\sigma_\gamma^2+\sigma_\eta^2\right]} \times \mathrm{CoDD}_{t-1}^{i,j} \\
&\quad - \frac{\alpha^2\sigma_\eta^2\left(\alpha^2\left[\sigma_\gamma^2+\sigma_\eta^2\right]\right)}{\left(\sigma_\varepsilon^2+\alpha^2\left[\sigma_\gamma^2+\sigma_\eta^2\right]\right)^{3/2}} \times \frac{\mathrm{CoHFI}_{t-1}^{i,i}+\mathrm{CoHFI}_{t-1}^{j,j}-2}{2} \\
&\quad - \frac{\alpha^2\sigma_\zeta^2\left(\alpha^2\left[\sigma_\gamma^2+\sigma_\eta^2\right]\right)}{\left(\sigma_\varepsilon^2+\alpha^2\left[\sigma_\gamma^2+\sigma_\eta^2\right]\right)^{3/2}} \times \frac{\mathrm{CoDD}_{t-1}^{i,i}+\mathrm{CoDD}_{t-1}^{j,j}}{2}
\end{aligned}$$

which we can rearrange to obtain

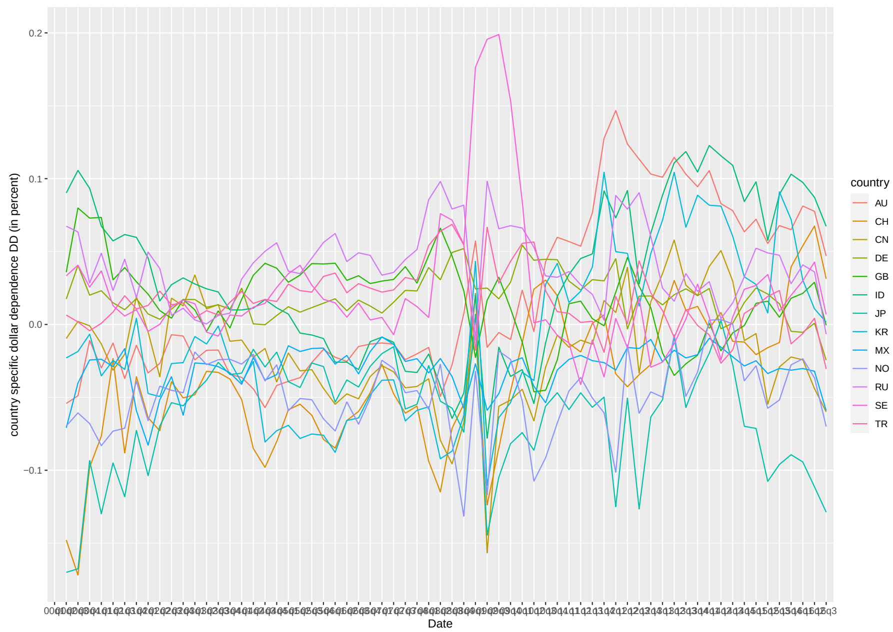
$$\begin{aligned}
\kappa &= \frac{\alpha^2\sigma_\gamma^2}{\sigma_\varepsilon^2+\alpha^2\left[\sigma_\gamma^2+\sigma_\eta^2\right]} + \frac{\alpha^2\sigma_\eta^2\left(\alpha^2\left[\sigma_\gamma^2+\sigma_\eta^2\right]\right)}{\left(\sigma_\varepsilon^2+\alpha^2\left[\sigma_\gamma^2+\sigma_\eta^2\right]\right)^{3/2}} \\
a &= \frac{\alpha^2\sigma_\eta^2}{\sigma_\varepsilon^2+\alpha^2\left[\sigma_\gamma^2+\sigma_\eta^2\right]} \\
b &= \frac{\alpha^2\sigma_\zeta^2}{\sigma_\varepsilon^2+\alpha^2\left[\sigma_\gamma^2+\sigma_\eta^2\right]}
\end{aligned}$$

and

$$n_{t-1}^{ii} = -\frac{\alpha^2\sigma_\eta^2\left(\alpha^2\left[\sigma_\gamma^2+\sigma_\eta^2\right]\right)}{\left(\sigma_\varepsilon^2+\alpha^2\left[\sigma_\gamma^2+\sigma_\eta^2\right]\right)^{3/2}} \times \frac{\mathrm{CoHFI}_{t-1}^{i,i}}{2} - \frac{\alpha^2\sigma_\zeta^2\left(\alpha^2\left[\sigma_\gamma^2+\sigma_\eta^2\right]\right)}{\left(\sigma_\varepsilon^2+\alpha^2\left[\sigma_\gamma^2+\sigma_\eta^2\right]\right)^{3/2}} \times \frac{\mathrm{CoDD}_{t-1}^{i,i}}{2}$$

D Supplementary Figures and Tables

Figure D.1: Time variation in dollar dependence across borrower countries



Note: This figure plots the country-specific component of dollar dependence $\mathrm{DD}_t^i = \sum_{b \in \mathcal{B}(i)}^{N} \omega_t^{b,i} \lambda_t^b$, for a selection of borrower countries in our sample. The country-specific component is DD_t^c minus the cross-sectional (time t) mean DD_t^c across all countries.

Table D.2: **Instrumental variables regressions for mortgage growth synchronization**

Dependent Variable:	correlation of mortgage growth					
GIV constructed using:	single global factor			multiple regional factors		
	(1)	(2)	(3)	(4)	(5)	(6)
Variables						
$\text{CoDD}_t^{i,j}$	0.017	0.133***	0.133	4.18*	3.03*	3.02*
	(0.223)	(2.93)	(1.16)	(1.89)	(1.93)	(1.92)
CoHFI_t^{ij}	0.017	-0.183	-0.179	-0.038	-0.224	-0.220
	(0.035)	(-0.399)	(-0.396)	(-0.074)	(-0.461)	(-0.462)
GDP growth corr.		0.197**	0.197**		0.199***	0.199***
		(2.52)	(2.56)		(2.84)	(2.84)
trade integration			-4.90			-4.73
			(-0.230)			(-0.226)
Fixed-effects						
CountryPair	Yes	Yes	Yes	Yes	Yes	Yes
country1-date	Yes	Yes	Yes	Yes	Yes	Yes
country2-date	Yes	Yes	Yes	Yes	Yes	Yes
Fit statistics						
F-test (1st stage), CoDD	242,932.4	228,191.9	228,190.4	1,843.4	1,810.0	1,809.7

Note: This table reports IV results equation (10) but with mortgage growth as the dependent variable for the period from 2000Q1 to 2020Q4 using the granular instrument $\mathcal{G}^{\text{CoDD}}$ defined in (12). The dependent variable $\text{HPcorr}_t^{i,j}$ is the five-year ahead rolling window correlation of HPgrowth in countries i and j. The explanatory variables are US dollar co-dependence $\text{CoDD}_t^{i,j}$ and the co-Herfindahl index $\text{CoHFI}_{t-s}^{i,j}$. Two versions of $\mathcal{G}^{\text{CoDD}}$ are used: a version taking account of a single, homogenous global factor, constructed as the difference between the market share-weighted and the equally weighted US dollar dependence. Second, a version in which the granular residuals u_{t-1}^{b} are estimated as the residuals of a model with several regional factors as discussed in section 5.3.
Standard errors are clustered two-way, by country i and country j, t-statistics are shown in parentheses. *, ** and *** indicate statistical significance at the 10%, 5% and 1% level, respectively. First-stage F-statistics are reported in the last row of the table.

Table D.3: **Robustness to alternative synchronization measures**

Dependent Variable:	synchronization of ..			
	house price growth		mortgage growth	
	(1)	(2)	(3)	(4)
sync. measure:	covariance	avg. beta	covariance	avg. beta
Variables				
$\text{CoDD}_t^{i,j}$	0.0022^{***}	3.028^{***}	0.0014^{**}	0.4695
	(3.111)	(3.196)	(2.661)	(0.5995)
$\text{CoHFI}_t^{i,j}$	0.0012	0.6889	0.0018	-0.4173
	(1.624)	(0.5430)	(1.195)	(-0.1817)
GDP growth corr.	2.404^{*}	-0.0069	1.460	0.1317
	(1.741)	(-0.3432)	(0.8423)	(1.562)
trade integration	-0.0284	6.926	-0.0070	-51.50
	(-0.9151)	(0.1476)	(-0.2331)	(-0.9798)
Fixed-effects				
CountryPair	Yes	Yes	Yes	Yes
country1-date	Yes	Yes	Yes	Yes
country2-date	Yes	Yes	Yes	Yes
Fit statistics				
Observations	26,894	26,894	25,691	25,691
R^2	0.71415	0.48707	0.51605	0.43964
Within R^2	0.00718	0.00095	0.00079	0.00220

Note: This table reports the results from estimating equation (10) for house price and mortgage growth for different synchronization measures. Columns (1) and (3) report results for pairwise covariances. Columns (2) and (4) show results for average pairwise betas computed as $0.5(\beta_{ij} + \beta_{ji})$ where β_{ij} (β_{ji})is the regression coefficientof house price or mortgage growth in country i (j) on the same variable in country j (i). All synchronization measures are computed over rolling windows of 16 quarters. *, ** and *** indicate statistical significance at the 10%, 5% and 1% level, respectively.

Table D.4: **Robustness across maturities and horizons**

Dependent Variable:	correlation of house price growth		
	(1)	(2)	(3)
treasury basis at tenor	1yr	3yr	10 yr
Variables	(4qtr)	(8qtr)	(20qtr)
$\text{CoDD}_t^{i,j}$	0.8157	1.207	0.6363**
	(1.092)	(1.401)	(2.341)
$\text{CoHFI}_t^{i,j}$	0.1915	-0.0108	0.0522
	(1.478)	(-0.0398)	(0.0951)
GDP growth corr.	-0.0012	0.0367	0.1813***
	(-0.1535)	(1.437)	(3.013)
trade integration	-14.97**	-11.18	0.6033
	(-2.467)	(-0.7798)	(0.0280)
Fixed-effects			
CountryPair	Yes	Yes	Yes
country1-date	Yes	Yes	Yes
country2-date	Yes	Yes	Yes
Fit statistics			
Observations	26,894	26,894	26,894
R^2	0.33142	0.41387	0.59636
Within R^2	0.00022	0.00114	0.01002

Note: This table reports the results for equation (10) with dollar co-dependence computed based on treasury bases at maturities of 1-year, 3-years, and 10-years respectively. To align the correlation horizon with maturities without loosing too many observations, we set the rolling window width to 4, 8 and 20 quarters for the 1-, 3- and 10-year maturity, respectively. *, ** and *** indicate statistical significance at the 10%, 5% and 1% level, respectively.

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