Non-US global banks and dollar (co-)dependence: how housing markets became internationally synchronized

Torsten Ehlers

Mathias Hoffmann

Alexander Raabe*

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Abstract

Non-US global banks are an important driver of the international synchronization of house price growth. A loosening (tightening) of US dollar funding conditions leads non-US global banks to expand (contract) their foreign lending to borrowing countries. This induces a synchronization of lending across borrowing countries, which translates into an international synchronization of house price growth. This synchronization is driven by borrowing countries' indirect joint exposure to US dollar funding conditions via their non-US global creditor banks, not their common-lender exposures. We refer to this indirect joint exposure as "dollar codependence". We show theoretically and empirically that the exposure of non-US global banks to dollar funding conditions is captured by the bilateral treasury basis between the currency of the non-US global creditor banks' headquarters and the US dollar. Our results identify a novel international spillover channel of US dollar funding conditions: as these conditions vary over time, borrowing country pairs whose non-US global creditor banks are more exposed to US dollar funding variations exhibit higher house price synchronization.

Key words: house prices, synchronization, US dollar funding, global US dollar cycle, US treasury basis, convenience yield, global imbalances, capital flows, global banks, global banking network

JEL classification: F34, F36, G15, G21

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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 considerably over time and peaked in the run-up to the 2008 Global Financial Crisis, and again in the euro 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 it is the single most important collateralizable asset. Identifying the drivers of the international synchronization of house prices is therefore paramount to understand macro-financial linkages and financial stability at the global level.

This paper highlights a new international spillover channel of US dollar funding conditions. We show that their variation 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: global banks' sensitivity to dollar funding conditions, and borrowing countries' exposure to global banks as determined by the structure of the international bank lending network. If dollar funding conditions ease, global banks increase their foreign lending, which is mostly denominated in US dollars. We show that the additional international lending translates into higher domestic mortgage credit 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, and in turn the higher the global banks' sensitivity to US dollar funding conditions, the stronger is the effect on mortgage credit and thus house prices. The higher this indirect dependence on dollar funding conditions for a pair of any two borrowing countries—, leading to higher dollar co-dependence, the higher the co-movement of house prices. An important element of this spillover mechanism is non-US global banks' sensitivity to US dollar funding variations. 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 the issuance of US dollar-denominated loans. This dependence on US dollar funding is what makes non-US global banks susceptible 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 the 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 this interpretation. 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 wholesale funding markets . However, as opposed to synthetic funding, direct dollar funding ties up balance sheet capacity because it exposes non-US global banks to exchange rate risk. Notably, an appreciation of the US dollar will reduce non-US banks' risk-taking capacity, leading 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 harks back to 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)). Our empirical analysis documents that these non-US global banks play a central role in facilitating the international spillover of US dollar funding conditions to real estate markets worldwide. Our focus on non-US global banks is further motivated by the recent literature on the special 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

¹While the non-US global banks US dollar dependence is a key analytical feature, our analysis accounts 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 market-share weighted average of the bilateral US treasury bases.

their indirect US dollar funding exposure as measured by our conceptof 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 proceed along the same lines to establish the 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 banksC and country B from lender banksD, 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 banksD only affect country B. Therefore, uncorrelated lender bankspecific 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

²Instead 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. 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.

exposed 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 indirect impact of dollar funding shocks on house price synchronization via non-US global banks from the direct impact of common-lender specific shocks — including shocks to US banks. As we show, it is indeed the former, indirect, 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. For example, such feedbacks could arise 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 timevarying 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 (2020) and construct a granular instrumental variable (GIV) that purges lennder banks' dollar dependence of the potential feedback from common demand factors in borrowing countries. To our knowledge, ours is the first paper that applies the method of Gabaix and Koijen (2020) to the study of pairwise co-movement between macroeconomic variables. Thus, our approach extends the methodology of Landier et al. (2017) to settings where quasi-natural experiments are not readily available for identification.

Our analysis contributes to the literature on international capital flows and house prices Aizenman and Jinjarak (2009); Ferrero (2015); Hoffmann and Stewen (2020); Sá et al. (2014). With only a few exceptions (Alter et al. (2018), Milcheva and Zhu (2016)) this literature has not looked at international correlations in house prices nor has it 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); Boz et al. (2019); Cerutti et al. (2017); Habib and Venditti (2019); Miranda-Agrippino and Rey (2019); 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 that directly influences capital and trade flows (Avdjiev et al. (2018); Boz et al. (2018, 2017); Bruno and Shin (2019); Gopinath and Stein (2018a,b)) because it reflects the international shadow price of bank leverage. A cheaper US dollar relaxes financing conditions for global banks, directly affecting credit supply and investment in borrowing countries. Our focus in this paper is to show how US dollar funding conditions proliferate through non-US global banks rather than directly through internationally active US or ultimate borrowing country banks, and how 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) has more results on the transmission mechanism and additional robustness checks. Section (8) 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 indirect exposure of borrowing countries to US dollar funding conditions via their respective lender banks' exposure to US dollar funding conditions.

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

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

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*. As further explained in section 2.1 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.

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 in wholesale markets which—unlike synthetic funding—ties up balance sheet capacity because it exposes non-US banks' (home currency denominated) balance sheets to exchange rate risk. Thus, the bilateral treasury basis is a suitable measure for $\lambda_{n,t}^b$, as it captures non-US bank's exposure 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, i.e. as the weighted average of the bilateral treasury bases of lender bank *b* of country *i*, with the market shares $\omega_t^{b,i}$ 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} ,

³Note 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 $-\lambda$.

⁴Note 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_t^i captures how dollar funding conditions affect borrowing country *i* through non-US lending banking systems.

we run a sequence of cross-sectional regressions on quarterly data from 2000 to 2020:

$$\Delta HP_t^i = \zeta_t \times DD_{t-1}^i + \text{constant}_t + \varepsilon_t^i$$

where ΔHP_t^i measures house price growth in country *i*.⁵ Figure 2 plots the sequence of estimated coefficients { ζ_t } against the four-quarter change in the effective US dollar exchange rate, an important measure of US dollar funding conditions (Avdjiev et al. (2018)). The correlation between the two time series, at -0.4, is striking suggesting that house prices rise as the US dollar depreciates, and vice-versa and that this link is stronger for countries with higher dollar dependence.

Our identification strategy relies on cross-country heterogeneity in DD_t^i . Figure D.1 plots DD_t^i (relative to its cross-county, time-t mean) for a selection of borrowing countries in our sample. Note that DD_t^i 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 $\omega_t^{b,i}$, as well as by the heterogenous exposure of lender banks *b* to variations in US dollar funding as given by λ_t^b .

The analytical framework that we propose in the next section, 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:

$$CoDD_t^{i,j} = DD_t^i \times DD_t^j$$
⁽²⁾

As we will show both theoretically and empirically, the synchronization of house price growth in two arbitrary borrowing countries *i* and *j* increases in $C_{ODD}_{t}^{i,j}$. Section (3) further expands on this intuition: 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

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

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. While the composition of country pairs in each portfolio varies over time, country pairs with the highest dollar codependence at any given point in time display, on average, the highest synchronization of house prices.

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, on average, highly exposed to variations in US dollar funding.

2.1 Measuring lender banks' exposure 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 US money market. 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. (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 $\lambda_{n,t}^b$ 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 money market. We refer to this option as direct US dollar funding. The non-US bank will still incur higher capital charges than the US bank for direct dollar funding, even when assuming equal rates for US and non-US banks borrowing in US dollar wholesale markets. This is because home country regulation requires the aggregate balance sheet of the non-US bank to be denominated in its non-US home currency. Thus, home currency value of the banks' global asset positions remains subject to exchange rate risk. This is true even if every unit of US dollar-denominated lending is financed by direct US dollar-denominated borrowing, avoiding any currency mismatch. 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.⁶

We would expect that 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,,translating into higher shadow costs of balance sheet capacity. This makes them particularly exposed 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. (2018)).

3 Analytical framework

We adapt and extend the methodological framework of Landier et al. (2017) for our analysis. These 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 lynchpin of 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

⁶Non-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))

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 head-quartered in a country instead of individual banks. Therefore, our framework is based on bilateral country-level exposures to *banking systems* rather than bilateral US state-level exposures to individual banks.⁷

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 banksthat 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' indirect exposure to dollar funding variations through borrowing countries' lender banks' exposure to these variations—independent of whether these lending banking systems are common to both countries in a pair. In addition, our framework leaves room for 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 do not support the common lender effect.

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

$$\frac{\Delta HP_t^i}{HP_{t-1}^i} = \alpha \frac{\Delta L_t^i}{L_{t-1}^i} + \varepsilon_t^i$$
(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

⁷As a reminder, we continue to refer to these banking systems as lender banks whenever this usage is unambiguous.

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

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 codependence on the synchronization of housing markets. The heterogenous exposure is given by λ_{t-1}^b . Section (2.1) provides the theoretical motivation why the bilateral treasury basis between the home currency of lender bank *b* and the US dollar constitutes an appropriate choice for 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

$$\frac{\Delta \mathrm{HP}_{t}^{i}}{\mathrm{HP}_{t-1}^{i}} = \alpha \gamma_{t} + \alpha \left(\sum_{b \in \mathcal{B}(i)}^{N} \omega_{t-1}^{b,i} \eta_{t}^{b} \right) + \alpha \underbrace{\left(\sum_{b \in \mathcal{B}(i)}^{N} \omega_{t-1}^{b,i} \lambda_{t-1}^{b} \right)}_{\mathrm{Dollar dependence}} \times \zeta_{t} + \varepsilon_{t}^{i} \tag{5}$$

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, v_t^i , the global factor γ_t and the factor ζ_t reflecting US dollar refinancing conditions 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*:

$$HP_{cov_{t-1}} = \alpha^{2} \sigma_{\gamma}^{2} + \alpha^{2} \sigma_{\eta}^{2} \underbrace{\left(\sum_{b \in \mathcal{B}(i) \cup \mathcal{B}(j)}^{N} \omega_{t-1}^{i,b} \omega_{t-1}^{j,b}\right)}_{co-Herfindahl} + \alpha^{2} \sigma_{\zeta}^{2} \underbrace{\left(\sum_{b \in \mathcal{B}(i)}^{N} \omega_{t-1}^{i,b} \lambda_{t-1}^{b}\right) \left(\sum_{b \in \mathcal{B}(j)}^{N} \omega_{t-1}^{j,b} \lambda_{t-1}^{b}\right)}_{dollar co-dependence}$$
(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 codependence 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 bankson the synchronization of house price growth .

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

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

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

$$HP_{corr_{t}^{i,j}} = \kappa + a \times CoHFI_{t-1}^{i,j} + b \times CoDD_{t-1}^{i,j} + n_{t-1}^{i,i} + n_{t-1}^{j,j}$$
(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

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 times series of credit to households and non-profit institutions serving households, provided by the BIS. The sample period for both data set 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 price, and analogously for the synchronization of mortgage credit growth. Because house price and mortgage growth are themselves measured four-quarters ahead, this effectively 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.

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.

⁸Our 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 please refer to appendix (A.1).

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

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 CoHFI^{*i*,*j*} and the dollar co-dependence 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 fundswithin a banking 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

In this section, we report our 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 mo-

¹⁰Foreign 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.

¹¹For 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.

ments 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:

$$HPgrowth_{t}^{i} = \boldsymbol{\beta}' \mathbf{DF}_{t} \times DD_{t-1}^{i} + \beta_{0} DD_{t-1}^{i} + \nu_{i} + \tau_{t} + \xi_{t}^{i}$$

$$\tag{9}$$

where HPgrowth^{*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), v_i and τ_t are country and time fixed effects respectively, and DF_t denotes a vector of variables driving US dollar funding conditions. The following variables enter the vector DF_t : i) the US Federal funds rate (including shadow rates for the period at the zero lower bound) to account for the effect of US monetary policy, ii) changes in the real effective exchange rate of the US dollar as shown by Avdjiev et al. (2018) to be an important driver of cross-border investment and also suggested by our model 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 shock that improves refinancing conditions and relaxes leverage constraints for banks borrowing in the US money market. . , We further include as broad 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 coefficient of interest, and should be signed such that an improvement in US dollar funding conditions loosens non-US banks' balance sheet capacity, increasing cross-border capital flows into foreign mortgage markets, and increasing house prices in borrowing countriesTo isolate the supply driven lending channel in response to a change in US dollar funding conditions, we include country and time fixed effects, ν_i and τ_t , to control for timeinvariant country characteristics and time-varying factors affecting all borrowing countries at the same point in time homogeneously, respectively. Note that the stand alone term of the vector DF_t is absorbed by the full set of time fixed effects. We cluster standard errors by both the country and time dimension to account for the correlation across borrowing countries at each point in time as well as within borrowing countries over time.

Table (1) shows the results for regression (9). We first report results for the individual factors in columns (1) to (5). In column (6), we then consider all factors jointly. The results for the individual factors suggest that an increase in capital inflows and broker-dealer leverage lead to higher house price growth in borrowing countries while a general appreciation of the dollar tightens funding constraints of non-US banks and thus leads to lower house price growth, consistent with our theory. Interestingly, the VIX is not individually significant while the federal funds rate is significant and positive. When we consider all factors jointly in column (6), only the federal funds rate, the dollar factor and treasury inflows retain their significance and the associated coefficients all remain stable relative to the specifications in the previous columns. By contrast, the coefficients on broker-dealer leverage and VIX are now both insignificant.

Hence, the three keyfactors that both the literature and ourtheory directly associate with US dollar funding conditions—the US Federal Funds rate, treasury flows and the US dollar exchange rate—are exactly the ones that the empirical analysis identifies as relevant. Conversely, indicators of the global financial cycle such as the VIX and broker-dealer leveragedo not primarily seem to affect house prices in borrowing countries through non-US global banks.

This could be explained by neither the VIX nor the broker-dealer leverage being specific to dollar funding, and thus not applicable to spillover of US dollar funding conditions.¹²

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

$$HPcorr_{t}^{i,j} = \beta \times CoDD_{t-1}^{i,j} + \delta \times CoHFI + CONTROLS_{t}^{i,j} + \theta_{i,j} + \mu_{t}^{i} + \delta_{t}^{j} + \epsilon_{t}^{i,j}$$
(10)

¹²For instance, the VIX measures the implied volatility of the U.S. stock market, and this is likely to imperfectly capture non-US banks' risks in managing their balance sheets.

where $\operatorname{HPcorr}_{t}^{i,j}$ denotes the conditional correlation of house price growth between borrowing countries *i* and *j*. We compute $\operatorname{HPcorr}_{t}^{i,j}$ using a forward rolling window of 16 quarters from period *t*. Our coefficient of interest is the one on the dollar co-dependence term, $\operatorname{CoDD}_{t-1}^{i,j}$. 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 the co-Herfindahl index $\operatorname{CoHFI}_{t-1}^{i,j}$ 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, and therefore helps us to considerably strengthen 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 timevarying country-*i* or country-*j* specific shocks — including any country-specific demand- or supply shocks for housing and foreign-funded credit— are controlled forby saturating the regression with a country-time 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 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 dol-

lar 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 CoHFI^{*i*,*j*} is an order of magnitude smaller than our estimate of β and insignificant throughout.

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 large global component in countries' exposure to dollar funding conditions and this component could affect house price synchronization in borrowing countries through other channels than through non-US banks.¹³ This could lead to reverse causality in the synchronization regression (10). To make things precise, assume that some global factor affects the bilateral treasury basis so that

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

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

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

Then $\epsilon_t^{i,j}$ will be correlated with $C_{0}DD_{t-s}^{i,j}$ and OLS estimates of $\alpha\sigma_{\zeta}^2$ would be biased.¹⁴ In this setting, our OLS estimations would suggest that global dollar funding shocks affect borrowing-

¹³In 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)Krishnamurthy and Lustig (2019)).

¹⁴Note that OLS is biased only if the loading $\psi_f^{i,j}$ is country-pair specific. If the loading was country-specific only, such that $\varepsilon_t^{i,j} = \psi^i f_t^n + \psi^j f_t^n + \psi_t^{i,j}$, the confounding effects of f_t^n would already be absorbed by the country *i*- time and country *j*-time effects in (10) above.

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 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 (2020) 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.¹⁵ 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:

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

where

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

is a set of granular weights with $\sum_{i} \Gamma_{t-1}^{i,b} = 0$ and $\#\mathcal{B}(i)$ is the number of lender banksactive in borrowing country *i*. 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), \mathcal{DD}_{t-1}^i can be constructed as the difference between the market share-weighted (defined as in 1) and the equally-weighted dollar dependence (defined as $DD_{t-1}^{E,i} = \sum_{b \in \mathcal{B}(i)}^{N} \frac{\lambda_{t-1}^b}{\#\mathcal{B}(i)}$) without need to estimate the individual u_{t-1}^b . To see this, note that we can write

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

where the first term is zero because of $f_{t-1} = \sum_{b \in \mathcal{B}(i)}^{N} \omega_{t-1}^{i,b} f_{t-1} = \frac{1}{N} \sum_{b \in \mathcal{B}(i)}^{N} f_{t-1}$.

¹⁵Landier et al. (2017) exploit the the quasi-natural experiment of state-level banking deregulation in the U.S. as an instrument.

We use 13) to construct \mathcal{DD}_{t-1}^{i} and $\mathcal{G}^{\text{CoDD}}$ which we then use as an instrument in the synchronization regression (10). Columns (1) to (3) of Table 3 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 $\lambda_t^{b.16}$ That is, factors with heterogenous 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

$$\lambda_t^b = \sum_{r=1..R} \phi_t^{,b,r} f_t^r + u_t^b \tag{14}$$

where the f_t^r for r = 1...R is a set of unobserved ("regional") factors that could affect the synchronization between borrowing countries. The loading coefficients $\phi_t^{,b,i}$ capture the spillbacks of these factors on the lender banks.

Differently from the factor structure (11) above, the loadings $\phi_t^{b,r}$ can differ across lender banksb, which implies that \mathcal{DD}_t^i cannot simply be constructed as the difference between DD_t^i and $DD_t^{E,i}$. Instead, we have to estimate the individual u_t^b directly.

To obtain these estimates of u_t^b , following Gabaix and Koijen (2020), we could jointly estimate the loadings $\phi_t^{b,r}$ and the factors f_t^r using some atheoretical factor-analytical technique and then extract the residuals u_t^b . Alternatively, we could employ some economic theory to proxy the loadings $\phi_t^{,b,r}$ in terms of observable lender bank-specific haracteristics. This would allows us to estimate the common factors by OLS as the series of coefficients on the interaction between $\phi_t^{,b,r}$ and a time*t*-country-group-*r* dummy. The residual of this regression would then provide us with estimates

¹⁶For 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.

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 inperpret the $\phi_t^{b,r}$ 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 banksb 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 $\phi_t^{b,r}$ for each lender bank.In turn, thisallows us to directly estimate the regional factors by OLS. We do so for a set of four geographical factors: besides the homogenous 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).

Columns (4) to (6) of Table 3 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 3 remain far above the usual critical value of around 10. All our previous conclusions remain intact. Though 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.

6 Transmission mechanism

Our results presented so far show that US dollar funding conditions affect house prices and their synchronization through the international lending network of non-US global banks. In this section, we explore the transmission mechanism. Specifically, our analytical framework assumes that

shocks to foreign lending affect house prices via their impact on domestic mortgage lending. This assumption rests on the fact that housing represents the most important collateralizable financial asset, making domestic banks in borrowing countries likely to channel additional funds from abroad into mortgage lending. In fact, mortgage lending constitutes about 90 percent of household lending in the borrowing countries of our sample.¹⁷ This would suggest that the patterns we have documented for house prices should be mirrored by the dynamics of mortgage lending. To explore this prediction, we therefore, re-run regressions (9) and (10) for mortgage credit growth and its synchronization across country pairs as the dependent variables. The construction of mortgage credit synchronization follows that of house price synchronization, applying a 16-quarter-ahead rolling-window correlation.

Tables (4) and (5) show the results. They reveal virtually the same patterns we documented for house prices. Turning first to the mortgage credit growth regressions , Table (4) reveals that the US dollar funding factorsnarrowly affect lender banks' ability to raise fundingin US dollars. The Federal Fundsrate, the US dollar exchange rate and US treasury inflowsare individually and also jointly significant. Broader indicators of the global financial cycle do not survive in the joint estimation. Similar to the analysis for house price growth, this suggests that our dollar dependence measure captures the effective exposure of a borrowing country's mortgage credit growth to US dollar funding shocks, and not to broader measures of the global financial cycle.

Turning to the synchronization regressions in Table (5), we confirm that the previously documented patterns for house prices synchronization are also visible for mortgage credit synchronization. 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.

7 Robustness

We provide additional robustness checks in Tables D.1 and D.1. First, we examine if our results hold up for alternative measures of synchronization. While our main results are based on cor-

¹⁷The 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)).

relations, 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.1 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.2. 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.

8 Conclusion

In this paper we have shown that non-US global banks are an important driver of the international synchronization of house prices. Non-US global banks borrow in US dollars, which they lend on to borrowing countries. Variations in US dollar funding conditions induce asynchronization of foreign lending. This variation in foreign lending supply then affects local real estate markets in the borrowing countries, leading to an international synchronization in mortgage credit growth and house price growth. As a result, borrowing country pairs with higher dollar co-dependence denoting an indirect joint exposure to US dollar funding conditions through their non-US global creditor banks —exhibit higher house price synchronization in response to shifts in US dollar funding conditions. Neither shocks to common lending banking systems nor direct borrowing from US banks is the main driver of synchronization. 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, as we show empirically and theoretically. Our findings highlight how the structure of international bank lending affects the synchronization of real outcomes, and in particular of real estate markets, illustrating that the "double decker structure of the global banking system" has first-order implications for the synchronization of real outcomes at the global level.

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

Note: This figure exhibits a box plot of the pairwise international house price synchronization over 595 country pairs at each point in time for the period 2000Q1-2015Q1, with the thick blue bars indicating the interquartile range. House price synchronization is computed as the 16-quarter correlations of four quarter ahead house price growth in countries *i* and *j* constituting a country pair. House price growth is calculated based on the country-wide residential real house price indices obtained for 36 countries from the OECD.





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

$$\Delta HP_t^i = \zeta_t \times DD_{t-1}^i + \text{constant}_t + \varepsilon_t^i$$

where ΔHP_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.

Den en den (Wentelste				. <i>i</i>		
Dependent variable:	(1)	$\langle 0 \rangle$	HPgr	$owth_t^r$		(c)
	(1)	(2)	(3)	(4)	(5)	(6)
Variables						
$DD_{t-1}^{i} \times$ fed. fundsrate	0.062***					0.066**
<i>i</i> 1	(3.38)					(2.40)
$DD_{t-1}^{i} \times dollar factor$		-1.65***				-0.854**
<i>i</i> 1		(-3.36)				(-2.71)
$DD_{t-1}^{i} \times$ treasury inflows			4.08^{*}			4.78^{*}
			(1.89)			(1.71)
$DD_{t-1}^{i} \times broker$ dealer lev.				0.011***		-0.001
<i>i</i> 1				(2.99)		(-0.271)
$DD_{t-1}^i \times VIX$					-0.002	-0.002
<i>v</i> 1					(-0.689)	(-0.454)
DD_{t-1}^{i}	0.017	0.048	0.025	-0.169**	0.093	0.055
	(0.493)	(1.38)	(0.705)	(-2.73)	(1.23)	(0.682)
GDP growth	0.849***	0.894***	0.891***	0.868***	0.901***	0.841***
	(3.85)	(3.68)	(3.64)	(3.72)	(3.65)	(3.88)
Fixed-effects						
country	Yes	Yes	Yes	Yes	Yes	Yes
date	Yes	Yes	Yes	Yes	Yes	Yes
Fit statistics						
Observations	2,032	2,032	2,032	2,032	2,032	2,032
R ²	0.38626	0.37484	0.37384	0.37924	0.37296	0.38991
Within R ²	0.05650	0.03895	0.03742	0.04572	0.03605	0.06211

Table 1: House price growth and US dollar funding conditions

Note: This table reports the results from estimating equation (9) for our panel of borrowing countries for the period from 2000Q1 to 2020Q4. The dependent variable $HPgrowth_t^i$ is the growth rate of house prices in borrowing country *i* over four quarters ahead. The explanatory variables are country *i*'s US dollar dependence DD_{t-1}^i lagged by one quarter 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, 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.

Dependent Variable:	HPcorr $_{t}^{i,j}$			
	(1)	(2)	(3)	
Variables				
$\operatorname{CoDD}_t^{i,j}$	1.76***	1.76***	1.76***	
	(3.26)	(3.48)	(3.49)	
$CoHFI_t^{i,j}$	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	
\mathbb{R}^2	0.53560	0.54960	0.54962	
Within R ²	0.00150	0.00820	0.00823	

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

Note: This table reports the results from estimating equation (10) for the period from 2000Q1 to 2020Q4. The dependent variable $\operatorname{HPcorr}_{t}^{i,j}$ is the five-year ahead rolling window correlation of $\operatorname{HPgrowth}$ in countries *i* and *j*. The explanatory variables are US dollar co-dependence $\operatorname{CoDD}_{t}^{i,j}$ and the co-Herfindahl index $\operatorname{CoHFl}_{t-s}^{i,j}$. 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.

Dependent Variable:	$\operatorname{HPcorr}_{t}^{i,j}$						
GIV constructed using:	sinį	single global factor			multiple regional factors		
	(1)	(2)	(3)	(4)	(5)	(6)	
Variables							
$\operatorname{CoDD}_t^{i,j}$	1.76***	1.72***	1.72***	2.29*	2.04^{*}	2.04*	
·	(4.20)	(4.59)	(4.61)	(1.81)	(1.77)	(1.77)	
$\operatorname{CoHFI}_t^{ij}$	0.175	0.174	0.177	0.168	0.170	0.173	
	(0.329)	(0.313)	(0.318)	(0.313)	(0.303)	(0.308)	
GDP growth corr.		0.154***	0.154***		0.155***	0.155***	
		(3.14)	(3.14)		(3.13)	(3.12)	
trade integration			-5.32			-5.28	
			(-0.238)			(-0.236)	
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	1,553.2	1,533.0	1,532.7	

Table 3: Instrumental variables regressions for house price synchronization

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.

Dependent Variable:	mortgage growth					
1	(1)	(2)	(3)	(4)	(5)	(6)
Variables						
$DD_{t-1}^{i} \times$ fed. fundsrate	0.070**					0.072**
	(2.57)					(2.23)
$DD_{t-1}^{i} \times dollar factor$		-1.12				0.156
		(-1.66)				(0.179)
$DD_{t-1}^{i} \times$ treasury inflows			6.92**			8.57*
			(2.18)			(1.85)
$DD_{t-1}^{i} \times broker$ dealer lev.				0.015**		$1.42 imes 10^{-5}$
· -				(2.30)		(0.002)
$DD_{t-1}^i \times VIX$					-0.001	-0.002
v 1					(-0.260)	(-0.412)
DD_{t-1}^{i}	0.054	0.085	0.050	-0.207**	0.104	0.056
v 1	(1.14)	(1.19)	(0.666)	(-2.21)	(0.800)	(0.485)
GDP growth	1.73**	1.77**	1.76**	1.75**	1.78^{**}	1.71**
U U	(2.10)	(2.12)	(2.11)	(2.10)	(2.12)	(2.09)
Fixed-effects						
country	Yes	Yes	Yes	Yes	Yes	Yes
date	Yes	Yes	Yes	Yes	Yes	Yes
Fit statistics						
Observations	1,983	1,983	1,983	1,983	1,983	1,983
R ²	0.42470	0.41904	0.42018	0.42287	0.41870	0.42684
Within R ²	0.04809	0.03873	0.04062	0.04507	0.03816	0.05163

Table 4: Mortgage credit growth and US dollar funding conditions

Note: This table reports the results from estimating equation (9) for our panel of borrowing countries for the period from 2000Q1 to 2020Q4. The dependent variable is the growth rate of mortgage credit in borrowing country *i*, four quarters ahead. The explanatory variables are country *i*'s US dollar dependence DD_{t-1}^{i} lagged by one quarter 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, 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.

Dependent Variable:	correlation (1)	n of mortga (2)	ge growth (3)
Variables			
$\operatorname{CoDD}_t^{i,j}$	0.433***	0.644***	0.643***
r	(7.39)	(17.4)	(4.64)
$\operatorname{CoHFI}_t^{i,j}$	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	$6.69 imes 10^{-5}$	0.00899	0.00901

Table 5: Mortgage credit growth synchronization and dollar co-dependence

Note: This table reports the results from estimating equation (10) for the period from 2000Q1 to 2020Q4, but with the correlation of mort-gage growth as the dependent variable. The correlation is computed as the four-year ahead rolling window correlation of 4 quarter-ahead mort-gage growth in countries *i* and *j*. The explanatory variables are US dollar co-dependence $C_{ODD}_{t}^{i,j}$ and the co-Herfindahl index $C_{OHFI}_{t-s}^{i,j}$. 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.

A Appendix

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 t 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 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^* . To focus on international lending, we simplify the setup by assuming that the bank lends abroad only in US dollars (i.e. it has no domestic lending nor does it do any cross-border lending in its home currency). Any funds raised in non-US home currency therefore have to be converted into US dollars at the current exchange rate X^S (measured in non-US home currency per US dollar, meaning an increase in X^S is a US dollar appreciation). We assume 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 discount ∇ . The bank then lends in the US dollar market at a riskless lending rate r^I . 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

$$E_{t+1} = ((A - S/X^S)(1 + r^l) - (A - S/X^S)(1 + r^*))X_{t+1}^S + (1 + r^l)\frac{X^F}{X^S}S - (S - E)(1 + r)$$

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 exchange rate X_{t+1}^S — 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

$$E_{t+1} = ((A - S/X^{S})(1 + r^{l}) - (A - S/X^{S})(1 + r^{*}))X_{t+1}^{S} + (1 + r^{l})\left(1 + \underbrace{\frac{X^{F} - X^{S}}{\sum}}_{=:\nabla}\right)S - (S - E)(1 + r)$$

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

$$E_{t+} \approx \left[\left(\left[(1+r^l) - (1+r^*) \right] \underbrace{\left(1 - \frac{S}{AX^S} \right)}_{\text{direct $\$-funding share}} \frac{X_{t+1}^S}{X^S} \right) + (1+r^l)s \right] AX^S - (S-E_t)(1+r) + \nabla S$$

$$= \left[\underbrace{\left[r^l - r^* \right] \frac{X_{t+1}^S}{X^S}}_{\text{direct $\$-funding share}} \underbrace{(1-s)}_{\text{direct $\$-funding share}} + (1+r^l)s \right]_{\text{assets in domestic currency}} -S(1+r-\nabla) + E_t(1+r)$$

$$= \underbrace{\left[(1+r^l) \frac{X_{t+1}^S}{X^S} - \underbrace{\left((1-s)(1+r^*) \frac{X_{t+1}^S}{X^S} + s(1+r-\nabla) \right)}_{\text{funding costs}} \right] AX^S + E_t(1+r)$$

 $\Pi_{t+1}:=$ excess return on bank portfolio in home currency

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

Without additional constrains, the bank's problem is unbounded. For a given $\mathbb{E}_t \Pi_{t+1} > 0$, 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 is required to maintain a constant default probability. This is 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}

$$\Pi_{t+1}^{\min} \le \frac{-E_t(1+r)}{AX^S}$$

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

$$\Pi_{t+1} \ge \Pi_{t+1}^{min}$$

with a given default probability α so that

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

Then, given the variance σ^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

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

Hence, the VaR constraint imposes the following upper bound on the leverage of the bank:

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

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 non-US home currency assets AX^S that

the bank can hold for a given σ_{Π} and $\mathbb{E}_t \Pi_{t+1}$ (and a given *E*, of course). But the bank can further 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 simple mean-variance problem.

Let

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

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

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

Hence the first-order condition (for minimizing the denominator $\Psi \sigma_{\pi} - \mathbb{E}_t \Pi_{t+1}$. i.e maximizing leverage under the VaR constraint w.r.t. *s*) is

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

Here, we have assumed (as in Ivashina et al. (2015)) 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. Specifically, we assume $\nabla = \nabla(s)$ to be weakly convex.¹⁸

Rearranging then yields

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

as the implicit "solution" for *s*. Note how the right hand-side of this expression is negatively related to the treasury-basis, $r^* - r + \nabla$! Hence, a more negative basis is directly related to a lower synthetic funding share *s* and a higher direct funding share. Note that in our empirical implementation, we look at the bilateral basis the other way around, i.e. at $r - r^* - \nabla$. 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

¹⁸Note that the second derivative is given by $\nabla'(s) + s\nabla''(s)$ which is positive if $\nabla(s)$ is weakly convex, so that $\nabla''(s) \ge 0$. 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.

empirical regularity.

Example Assume a simple linear function for the forward-spread, i.e.

$$\nabla(s) = \kappa s$$

Then, we obtain

$$s = \frac{-(r^* - r + \kappa s) + (1 + r^*) \left[(\mu - 1) - \Psi \sigma_x\right]}{\kappa}$$

and therefore

$$s = rac{-(r^* - r) + (1 + r^*) \left[(\mu - 1) - \Psi \sigma_x\right]}{2\kappa}$$

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

A dollar depreciation

Consider first what happens after a US dollar depreciation. Hence, we can ask what happens to leverage after an increase in μ :

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

if we assume that $r^l > r^*$, then this is clearly positive: a US dollar depreciation ($d\mu > 0$) will lead to an increase in leverage and, thus, more US dollar lending.

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. So, if we can show what happens to the response above when we change *s*, we are done. To this end, first rewrite the 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

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

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. Hence, a lower *s* (higher $r - r^* - \nabla$) 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)

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

which will be negative whenever $\mu > \Psi \sigma_X$. This will usually be the case because μ is a gross

change $(\mu = \mathbb{E}_t \left(\frac{X_{t+1}^S}{X^S}\right) = \mathbb{E}_t (1 + \Delta \log(X_{t+1}))$ while $\sigma_X = \sigma (1 + \Delta \log(X_{t+1})) = \sigma(\Delta \log(X_{t+1}))$ is the volatility of a growth rate. Empirically, the 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

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

and then

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

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 (higher) syntehtic (direct) 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

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

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

$$C_{ODD}^{i,i} = C_{ODD}^{j,j} = C_{ODD}^{i,j} = 0$$

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:

$$Cohfl_{t-1}^{i,i} = Cohfl_{t-1}^{j,j} = Cohfl_{t-1}^{i,j} = 1$$

Then a first-order expansion yields

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

which we can rearrange to obtain

$$\kappa = \frac{\alpha^2 \sigma_{\gamma}^2}{\sigma_{\varepsilon}^2 + \alpha^2 \left[\sigma_{\gamma}^2 + \sigma_{\eta}^2\right]} + \frac{\alpha^2 \sigma_{\eta}^2 \left(\alpha^2 \left[\sigma_{\gamma}^2 + \sigma_{\eta}^2\right]\right)}{\left(\sigma_{\varepsilon}^2 + \alpha^2 \left[\sigma_{\gamma}^2 + \sigma_{\eta}^2\right]\right)^{3/2}}$$

$$a = \frac{\alpha^2 \sigma_{\eta}^2}{\sigma_{\varepsilon}^2 + \alpha^2 \left[\sigma_{\gamma}^2 + \sigma_{\eta}^2\right]}$$

$$b = \frac{\alpha^2 \sigma_{\zeta}^2}{\sigma_{\varepsilon}^2 + \alpha^2 \left[\sigma_{\gamma}^2 + \sigma_{\eta}^2\right]}$$

and

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

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 $DD_t^i = \sum_{b \in \mathcal{B}(i)}^N \omega_t^{b,i} \lambda_t^b$, 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.

Dependent Variable:	synchronization of				
	house price growth		mortgage	growth	
	(1)	(2)	(3)	(4)	
sync. measure:	covariance	avg. beta	covariance	avg. beta	
Variables					
$\operatorname{CoDD}_t^{i,j}$	0.0022***	3.028***	0.0014**	0.4695	
·	(3.111)	(3.196)	(2.661)	(0.5995)	
$\operatorname{CoHFl}_t^{i,j}$	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	
\mathbb{R}^2	0.71415	0.48707	0.51605	0.43964	
Within R ²	0.00718	0.00095	0.00079	0.00220	

Table D.1: Robustness to alternative synchronization measures

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 coefficient of 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.

Dependent Variable:	correlation of house price growth				
	(1)	(2)	(3)		
treasury basis at tenor	1yr	3yr	10 yr		
Variables	(4qtr)	(8qtr)	(20qtr)		
$\operatorname{CoDD}_t^{i,j}$	0.8157	1.207	0.6363**		
	(1.092)	(1.401)	(2.341)		
$\operatorname{CoHFI}_t^{i,j}$	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		

Table D.2: Robustness across maturities and horizons

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.