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Advisor: Mariano Massimiliano Croce

Co-Advisor: Ina Simonovska

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Felix Gerding
ID number: 3143305
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Thesis abstract

This thesis consists of three papers on global finance and international asset pricing. The first paper develops a structural asset-pricing model in which a country's equity risk exposure depends on a world-volatility shock, a local-growth shock, and a time-varying trade-network centrality measure that scales exposure to global risk. The model rationalizes why USD betas rise with centrality in developed markets but decline or remain flat in emerging markets, and why local-currency betas in EMs fail to increase despite deeper trade integration. The second paper builds a five-decade panel of stock market returns for 37 developed and developing countries and documents that poorer countries have both higher and more volatile returns. It shows that these facts are consistent with a consumption-based long-run risk model, which predicts that countries with more growth-sensitive dividends earn higher risk premia. The third paper investigates whether the U.S. dollar's global dominance is eroding. Using new data on reserve composition, FX transactions, global debt issuance, and trade invoicing, it finds little evidence of de-dollarization: U.S. dollar dominance remains stable through late 2023 despite geopolitical and macroeconomic shocks. Together, the three chapters provide new theoretical and empirical insights into global equity risk, long-run return behavior, and the resilience of the dollar-based international monetary system.

Contents

Chapter 1

Trade Network Centrality, Risk Exposures, and Equity Returns Across Developed and Emerging Markets

Abstract

This paper develops and empirically tests a parsimonious global asset-pricing framework in which each country's equity risk exposure depends on (i) a common world-volatility shock, (ii) a local-growth shock, and (iii) a time-varying trade-network centrality measure that scales exposure to global volatility. We show that the model can rationalize both the positive slope of USD-betas in Developed Markets (DMs) and its negative (or flat) counterpart in EMs, as well as the puzzling fact that EM local-currency betas remain flat or even decline with centrality. The calibrated model also reproduces the distribution of cross-sectional betas, FX exposures, and global-index volatility, providing an explanation of why trade-network centrality predicts opposite risk exposures in developed versus emerging markets, and why EM LC-betas fail to rise with centrality.

JEL: G11, G12, G15, F36, F47.
Keywords: Trade network centrality; international CAPM; currency risk; emerging markets; GMM; risk exposures; segmentation.

1.1 Introduction

Why do some countries' equity markets load more heavily on global risk than others? In canonical international asset-pricing frameworks, deeper integration—through trade, finance, or production—should strengthen a country's exposure to common global risk factors. When goods, capital, and information flow freely, shocks that affect the global stochastic discount factor (SDF) are transmitted uniformly across countries, and measures of integration such as openness or trade-network centrality should map monotonically into global exposure. This prediction is deeply rooted in models such as the Global CAPM and International CAPM, where all investors price assets using a common SDF denominated in U.S. dollars (USD) Gerding et al. (2025); Colacito and Croce (2013a); Lustig et al. (2014). In these environments, any increase in integration—whether through broader trade linkages, deeper financial openness, or participation in global value chains—should raise both USD- and local-currency (LC) betas, as both measure the same underlying exposure to globally priced shocks.

Across DMs, greater trade-network centrality—a recursive measure of global embeddedness—coincides with higher USD betas and roughly flat LC betas. In these economies, greater trade centrality reinforces the transmission of common shocks through supply chains, multinational firms, and synchronized consumption cycles. The positive slope between centrality and USD betas therefore fits neatly within standard integrated-market theory.

In emerging markets (EMs), however, the relationship reverses. More central EMs—those with stronger participation in global trade networks—exhibit *lower* USD betas, suggesting that greater centrality is associated with reduced measured exposure to USD-priced global risk. Even more strikingly, the same negative or flat relationship appears for LC betas, which remove the direct influence of exchange-rate movements. This pattern violates the central prediction of integrated-market models: if exposure to global shocks were priced uniformly, LC betas should rise, not fall, with integration. The fact that both USD and LC betas decline with centrality points to a deeper structural asymmetry between blocs.

This paper argues that the key lies in the *composition of trade partners*. Centrality is not a monolithic measure of "integration": it captures how connected a country is, but not *to whom*. The composition of partners—specifically, whether they are financially integrated DMs or segmented EMs—determines how centrality translates into priced exposure. When a DM's centrality rises, it typically does so by strengthening links with other DMs, whose cash flows are already priced in USD and whose shocks are highly correlated with the global factor. This form of integration amplifies exposure to the USD-priced SDF, generating the positive centrality–beta slope predicted by theory. When an EM's centrality rises, by contrast, it usually reflects the formation of new trade links with other EMs—partners that are financially segmented, trade in LC, and whose shocks are only weakly correlated with USD-priced global risk. Such shifts reallocate exposure toward locally priced shocks, lowering the covariance with the USD factor and producing a negative or flat slope in both USD and LC terms.

This interpretation unites two strands of research. The first emphasizes that trade integration need not coincide with financial integration: institutional frictions, capital controls, and limited risk-sharing imply persistent segmentation across blocs Bekaert et al. (2007); Alfaro et al. (2008b); Aguiar and Gopinath (2007a); Naoussi and Tripier (2013). The second highlights that a country's position in the global trade network shapes the transmission of shocks (Richmond, 2019; Fang, 2021; Du et al., 2021). By focusing on the composition of partners rather than aggregate openness, I show that network structure determines whether centrality amplifies or dampens priced risk. The same rise in centrality can therefore increase global exposure in DMs but reduce it in EMs—a fundamental asymmetry that standard integrated models cannot reconcile.

Several examples illustrate the mechanism. During the 2000s, Indonesia's centrality rose sharply as it expanded trade within ASEAN and with China, another emerging economy at the time. These new links deepened exposure to region-specific shocks but did little to increase sensitivity to USD-priced global volatility. By contrast, Belgium's rise in centrality over the same period reflected deeper integration within the Eurozone and with other advanced economies—partners fully embedded in USD- and euro-denominated capital markets. As a result, Belgium's global beta increased, while Indonesia's declined. Such comparisons underscore that centrality's effect on global risk exposure depends not on the number of connections but on their pricing regime.

I formalize this mechanism in a model in which a single USD investor prices all assets, but countries' cash-flow exposures to USD-priced risk depend on their network position. Country-level dividend growth loads on both a global volatility shock and a local growth shock, with trade centrality scaling the covariance with the global component. The slope of this exposure is regime-specific: in DMs, centrality strengthens the loading on USD-priced shocks ; in EMs, it dampens. This parsimonious structure embeds the composition channel within a conventional SDF framework and yields closed-form expressions for price–dividend ratios, betas, and dividend volatility. The model matches all three empirical regularities: opposite-signed centrality–beta slopes across regimes, the similarity of USD and LC slopes within each regime, and the decline in dividend-growth volatility with centrality among EMs. I discipline the model using only moments that capture the link between trade centrality and risk exposure. The GMM estimation targets the within-year slopes of CAPM betas with respect to trade centrality and the first two moments of real dividend growth across countries. No realized return data or mean beta levels enter the estimation. The mean betas and expected return differentials that the model produces therefore represent out-of-sample implications rather than fitted moments.

Empirically, I combine bilateral export data from the IMF's *Direction of Trade Statistics* (DOTS) with MSCI total-return indices for 32 countries over 1970–2022. Trade-network centrality is computed recursively as a weighted eigenvector measure capturing both direct and indirect exposure to global trade partners. Equity betas are estimated in both USD and LC using rolling-window regressions of national excess returns on the world portfolio. The data reveal three robust patterns. First, among DMs, USD betas rise significantly with centrality while LC betas are flat, consistent with complete integration under a USD-priced kernel. Second, among EMs, both USD and LC betas decline with centrality, consistent with increasing intra-EM trade and segmented pricing. Third, among EMs, dividend-growth volatility falls with centrality, suggesting diversification of idiosyncratic cash-flow risk through intra-bloc trade. These regularities are stable across alternative centrality measures, time windows, and classifications. I then estimate a No-Arbitrage model using GMM, matching moments on average betas, dividend volatility, and the cross-sectional slopes of betas on centrality in both currencies. The fitted parameters jointly reproduce the observed slopes and levels of USD betas as well as the difference in dividend volatility across blocs.

The paper makes three contributions. First, it provides new empirical evidence on how trade-network structure shapes exposure to global risk. Second, it offers a unified reduced-form model linking trade composition to priced exposure under segmentation, generating testable predictions for USD betas. Third, it introduces a quantitative measure of the *composition channel*—the differential mapping from trade centrality to global exposure across regimes. By integrating network theory with segmented pricing, the paper reframes trade centrality as a measure of *who* a country trades with, not simply how much it trades.

Roadmap. Section 1.2 discusses the related literature. Section 1.3 descibes the data and measurements. Section 1.4 documents the empirical facts connecting centrality, composition, and risk exposure. Section 2.4 develops the No-Arbitrage model and derives its closed-form predictions for equity betas and dividend volatility. Section 2.5 concludes.

1.2 Literature Review

This paper builds on and bridges two complementary perspectives on global risk exposure. The first, exemplified by (Gerding et al., 2025), emphasizes cross-country differences in the *cash-flow* exposure of dividends to persistent global shocks. The second, developed by Richmond (2019), highlights how a country's position within the global trade network shapes the *transmission and pricing* of those shocks. Together, these strands imply that differences in global risk exposure arise not only from fundamentals such as productivity or trend growth but also from the structure of international linkages. The contribution of this paper is to integrate these views in an empirically tractable framework that connects trade-network centrality, partner composition, and segmented pricing. I show that the same increase in trade centrality can raise exposure to USD-priced risk in developed markets (DMs) but reduce it in emerging markets (EMs), resolving an apparent contradiction between standard integrated-asset-pricing predictions and the empirical evidence on betas and volatility.

Richmond (2019) provides a complementary perspective by linking global risk exposure to trade-network structure. Centrality, in this view, is a fundamental determinant of exposure to the priced global factor. My analysis generalizes this logic to equity markets and introduces an explicit composition channel. The data, however, reveal the opposite in EMs. I show that this reversal arises because trade centrality is not homogeneous: in EMs, increasing centrality primarily reflects denser intra-EM trade within a segmented financial bloc. Once segmentation (dividing countries into Developed/Emerging markets) is introduced, the sign of the centrality–exposure relationship depends on the composition of partners, not just on the degree of connectedness. Thus, this paper transforms the positive, universal relationship in Richmond's currency model into a regime-dependent one at the equity level: centrality amplifies priced risk in DMs but buffers it in EMs.

The paper relates to work on global value chains and external financial linkages, which highlights that international shock transmission depends on *who* countries trade and borrow with, not only on how much they trade. Research by Kalemli-Ozcan et al. (2003) and Kalemli-Ozcan et al. (2013) shows that risk sharing, synchronization, and crisis propagation hinge on the depth of financial integration and the composition of cross-border exposures. In parallel, studies of global liquidity and the risk-taking channel emphasize that EM vulnerabilities depend more on USD funding and intermediation capacity than on trade openness itself (Bruno and Shin, 2015, 2017). By introducing trade-partner composition as an explicit driver of priced exposure, this paper provides a bridge between those macro-financial channels and network-based models of risk transmission. The resulting empirical pattern—opposite centrality–beta slopes across blocs—offers a new lens on the local-currency beta puzzle and the asymmetric role of trade integration in global risk pricing.

1.3 Data and Measurement

This section introduces the data and explains how it is used to link a country's position in the global trade network to its exposure to global equity risk, with particular attention to how the composition of trading partners shapes this relationship. The first part of the section presents the data sources, while the second part constructs the measures of trade-network centrality, trade-partner composition, and equity betas in both U.S. dollars (USD) and local currency (LC), which are later used in the empirical analysis.

1.3.1 Sample and Data Sources

To construct the empirical measures, I combine financial and macroeconomic data from several established sources, as detailed below. Equity return data are obtained from MSCI. I use both the Total Return Gross Index and the Price Return Index, available in USD and in local currency. Annual returns are calculated as log differences. Following the conventional practice in the finance literature, I derive dividends as the difference between the two indices. To obtain real returns, I deflate nominal returns using U.S. CPI inflation from the Federal Reserve Bank of St. Louis (FRED). Excess returns are computed by subtracting the mean nominal yield on three-month U.S. Treasury bills, also obtained from FRED. This procedure closely follows the data used in Gerding et al. (2025) to ensure comparability across markets and currencies.

Macroeconomic variables are drawn from the International Monetary Fund's World Economic Outlook (WEO) database, which provides consistent cross-country series on GDP, inflation, trade openness, and other macro aggregates. In this paper, I employ nominal GDP expressed in U.S. dollars as a benchmark measure of economic size and as a scaling factor: specifically, I use GDP to normalize bilateral trade flows so that trade linkages can be compared relative to economic mass.

For bilateral trade, I rely on the IMF's Direction of Trade Statistics (DOTS), which report annual exports and imports between every reporting country pair over the sample period. The DOTS data are widely used in international-trade and network studies because they deliver a nearly complete panel of bilateral flows, enabling the construction of country-to-country trade networks.

1.3.2 Trade-Network Centrality

I measure trade-network centrality, which aggregates both direct and indirect trade linkages into a single index. For each year t, I define the bilateral trade intensity between countries i and j as

$$TI_{ij,t} = \frac{Exports_{ij,t} + Exports_{ji,t}}{GDP_{i,t} + GDP_{j,t}} \times ShareExp_{i,t}, \tag{1.1}$$

where $ShareExp_{i,t}$ denotes country i's share of world exports in year t. This construction assigns greater weight to trade relationships involving countries that are large global exporters, thereby emphasizing economically significant links within the network.

I collect these intensities into an adjacency matrix W_t, row-normalize it so that each country's total trade link sums to one, and apply the Bonacich operator to obtain

$$C_t = (I - W_t')^{-1}\mathbf{1}_N, \tag{1.2}$$

where $\mathbf{1}_N$ is an $N \times 1$ vector of ones, and ϕ is a scalar attenuation (or decay) parameter capturing how rapidly the influence of indirect connections decays along longer paths in the network. Following common practice in network-based empirical work, ϕ is chosen to be less than the reciprocal of the largest eigenvalue of W_t, which guarantees convergence of the inverse and ensures that the measure remains bounded. Intuitively, when ϕ is close to zero, the centrality measure approaches simple degree centrality, capturing only direct trade links, while larger values of ϕ assign increasing importance to indirect connections through multiple layers of trading partners.

This measure has several desirable features. First, it admits a direct economic interpretation as accumulated exposure to the global trade network through both direct and indirect channels. Second, it is comparable across countries and over time because of the normalization procedure and the bounded scaling implied by ϕ. Lastly, it is sensitive to changes in partner composition: for instance, establishing a new trade relationship with a highly central partner immediately raises $C_{i,t}$, even if the relative shares of existing partners decline. These properties make the Bonacich measure particularly well suited for capturing how countries' systemic positions within the global trade network evolve and relate to their exposure to global equity risk.

1.3.3 Trade-Partner Composition

To capture the composition of trading partners, I distinguish whether a country's centrality reflects integration with developed or emerging counterparts. For each country–year pair, I compute the share of total exports directed toward emerging market partners and the complementary share directed toward developed market partners:

$$EMShare_{i,t} = \frac{\sum_{j \in EM} Exports_{ij,t}}{\sum_j Exports_{ij,t}}, \qquad DMShare_{i,t} = 1 - EMShare_{i,t}. \tag{1.3}$$

This decomposition is essential for identifying the sources of rising centrality. In developed markets, increases in centrality typically indicate deeper integration with already USD-priced partners, thereby reinforcing exposure to globally priced risks. In contrast, among emerging markets, higher centrality often stems from growing intra-EM trade, where financial segmentation and limited capital market integration weaken the link between real trade integration and exposure to global equity shocks. By separating these components, I can assess whether changes in centrality are driven by integration into globally priced networks or by expansion within more financially segmented trading blocs.

1.3.4 Equity Betas

Global risk exposure is quantified by estimating equity betas in both USD and local currency (LC) terms. Concretely, for each country i I estimate two seperate regressions:

$$r^{USD}i,t = \alpha^{USD}i + \beta^{USD}ir^{USD}W,t + \varepsilon^{USD}i,t, \tag{1.4}$$

$$r^{LC}i,t = \alpha^{LC}i + \beta^{LC}ir^{LC}W,t + \varepsilon^{LC}i,t, \tag{1.5}$$

where $r^{LC}W,t$ and $r^{USD}W,t$ denote returns on a global benchmark portfolio expressed in local currency and in U.S. dollars, respectively. I define the world portfolio as the market-capitalization–weighted average of equity returns in the United States, the United Kingdom, France, Germany, and Japan, following Gerding et al. (2025).

USD betas capture how much a country's equity return (including the effect of currency fluctuations) co-moves with global equity risk from the perspective of a U.S. investor. In effect, β_i^{USD} reflects the combined sensitivity of the stock market and the exchange rate to global shocks. In contrast, LC betas abstract away currency fluctuations and thus isolate the pure covariance between local equity returns and global equity shocks. Comparing β_i^{USD} and β_i^{LC} allows me to assess how much of a country's global exposure is transmitted through exchange rate channels. Estimating both betas is informative for two reasons. First, in financially integrated markets with freely floating exchange rates, β_i^{USD} and β_i^{LC} are expected to be similar, as currency movements fully reflect global risk pricing. Significant divergence between the two suggests the presence of currency risk premia or partial segmentation. Second, in emerging or segmented markets, the difference $\beta_i^{USD} - \beta_i^{LC}$ serves as a diagnostic for the extent to which global equity risk is transmitted via exchange rate fluctuations rather than through direct equity market linkages. Taken together, these two beta estimates provide complementary perspectives on a country's global risk exposure and highlight the role of currency channels in shaping that exposure.

1.4 Empirical Findings

This section documents four robust empirical regularities linking trade-network centrality, trade-partner composition, and exposure to global risk. Across all results, I compare developed markets (DMs) and emerging markets (EMs) and interpret the differences through the lens of the trade-composition channel.

The evidence unfolds in four exhibits:

(1) Centrality has *opposite effects* on global equity betas in DMs and EMs.

(2) These opposite slopes arise from differences in trade-partner composition: DMs' centrality gains reinforce intra-DM trade, whereas EMs' gains reflect intra-EM trade expansion.

(3) The same composition differences translate into opposite patterns of *shock transmission*: centrality amplifies crisis exposure in DMs but mitigates it in EMs.

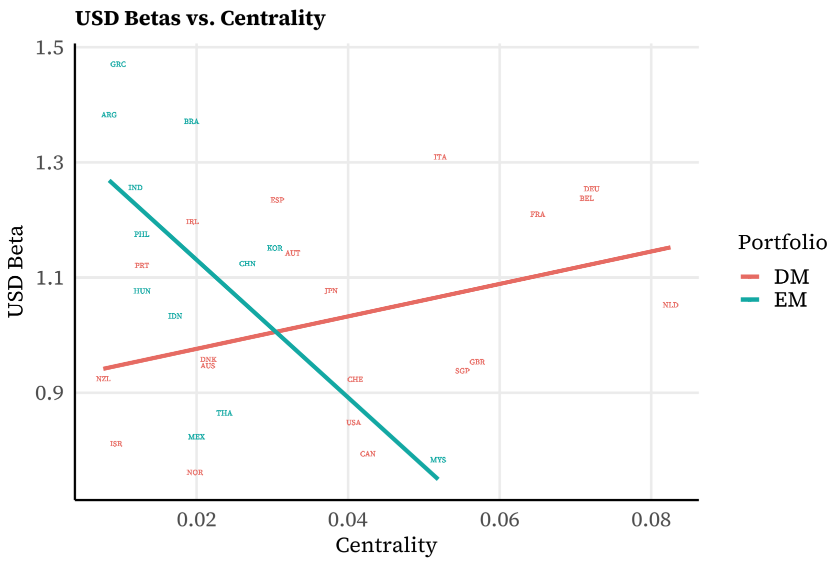
(4) Since 2001, rising intra-EM trade has made the β–centrality slope in EMs increasingly negative, linking trade reorientation to a structural

decline in global risk exposure.

1.4.1 Opposite Slopes in the Centrality–Beta Relationship (USD Denomination)

Figure 1.1 presents the baseline cross-sectional relationship between trade-network centrality and *USD-denominated* equity betas with respect to the world portfolio. Each point represents a country average over 1970–2022, with red markers denoting developed markets (DMs) and blue markers denoting emerging markets (EMs).

Figure 1.1: **Trade-network centrality and USD-denominated equity betas.**



Cross-sectional relation between a country's trade-network centrality and its equity beta with respect to the world portfolio (USD returns). Fitted lines show within-group OLS slopes.

The figure shows a striking asymmetry. In DMs, the slope is positive and statistically significant: more central economies exhibit higher global betas, indicating stronger co-movement with the USD-priced world factor. In EMs, the slope is negative and significant: more central economies exhibit

Table 1.1: **Cross-sectional regression: USD betas on trade-network centrality.**

	Dependent variable: β_i^{USD}
Centrality$_i$	3.02*
	(1.76)
$\mathbf{1}\{EM_i\}$	0.45***
	(0.14)
Centrality$_i$ $\times$ $\mathbf{1}\{EM_i\}$	-14.97^{***}
	(4.63)
Constant	0.92
	(0.62)
Observations	32
R^2	0.32

Standard Errors in parentheses. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

lower global betas, suggesting insulation from global risk. This opposite sign across regimes is the central empirical fact of the paper.

Regression estimates. Table 1.1 reports the regression

$$\beta_i^{USD} = \alpha + \gamma\,\text{Centrality}_i + \delta\,\mathbf{1}\{EM_i\} + \theta\left(\text{Centrality}_i \times \mathbf{1}\{EM_i\}\right) + \varepsilon_i, \tag{1.6}$$

where β_i^{USD} is the average USD beta and Centrality$_i$ the average trade-network centrality from 1970–2022. The interaction term allows the slope to differ systematically by regime.

The regression confirms the opposite-signed relationship: for DMs, $\hat{\gamma} > 0$ and significant; for EMs, $\hat{\gamma} + \hat{\theta} < 0$ and highly significant. The large negative interaction term ($\hat{\theta} = -14.97$) implies that the slope for EMs is significantly lower than for DMs, producing the clear reversal seen in Figure 1.1. The specification explains roughly one-third of the cross-country variation in global betas ($R^2 \approx 0.33$).

The positive slope for DMs aligns with the logic of integrated markets: more central DMs strengthen trade and financial ties with other USD-priced partners, increasing both cash-flow and discount-rate co-movement with global factors. The negative slope for EMs reveals segmentation: as EMs become more central, they connect primarily to other EMs, whose shocks are locally priced and less synchronized with the USD SDF. This compositional reweighting diversifies away globally priced risk even as gross openness increases. The positive β–centrality slope for DMs is consistent with the predictions of integrated-market asset-pricing models, in which all countries are priced by a common USD-denominated stochastic discount factor. As trade links among DMs expand, supply chains, corporate earnings, and consumption become increasingly synchronized, raising the covariance of both cash flows and discount rates with the global SDF. In these economies, a higher degree of trade-network centrality effectively embeds the country more deeply within the global pricing kernel: shocks to the USD world factor translate more strongly into local returns. This logic aligns with global CAPM and international CAPM formulations (e.g., Colacito and Croce, 2013a; Lustig et al., 2014) and with the general equilibrium mechanism emphasized by Obstfeld (1993), where trade openness and cross-border production links magnify the transmission of world shocks to domestic asset prices.

For EMs, the negative slope points to a segmentation margin that standard integrated models cannot accommodate. Rising centrality in EMs does not necessarily represent stronger connections to the USD-priced core but instead reflects expansion within the EM periphery. New or intensified links are typically with other EMs—partners whose exports, invoicing, and financial systems are denominated in local currency or partially insulated from USD-based pricing. As a result, greater centrality reweights trade toward locally priced shocks and reduces exposure to the global USD factor. This "composition channel" operates even when overall openness increases: the direction of integration, not its volume, determines whether global exposure rises or falls.

From a structural perspective, the asymmetry reflects the coexistence of an integrated core and a segmented periphery in the global trade and financial system. In the core, dominated by the United States and other advanced economies, trade and financial flows are denominated in USD, and prices adjust jointly to common global shocks. In the periphery, comprised of emerging markets, the same shocks are only partially transmitted through local-currency pricing, capital controls, or limited investor participation. Thus, the same rise in trade centrality can have opposite effects on measured USD betas depending on whether the additional links are with globally priced or locally priced partners.

The implications extend beyond asset pricing. They suggest that globalization is multidimensional: real integration through trade does not guarantee financial integration in pricing. For DMs, deeper trade embeddedness within the USD core amplifies exposure to world risk. For EMs, deeper embeddedness within an EM network attenuates it, creating a form of "network-based decoupling." These results provide direct, quantitative evidence that the global economy exhibits a dual structure—an integrated core that transmits global shocks and a segmented periphery that increasingly internalizes them.

Finally, it is worth emphasizing that the results are robust across alternative specifications and not driven by mechanical differences in size or income. Controlling for initial log income or GDP per capita yields virtually identical coefficients (Appendix 1.8). Replacing centrality with simpler measures of openness, trade concentration, or import–export breadth produces the same sign reversal (Appendix 1.7). Taken together, these findings confirm that the relationship between trade centrality and USD risk exposure reflects a structural network composition effect, not a spurious correlation with development or trade intensity.

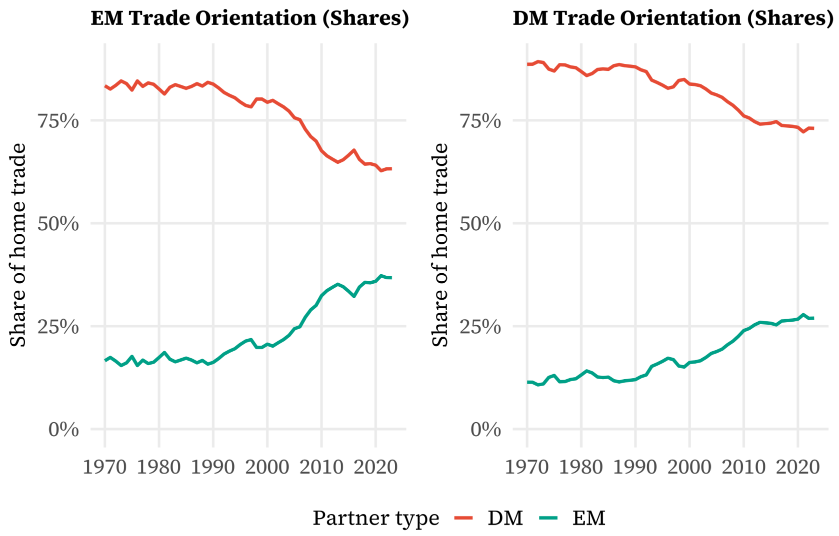
Robustness to currency denomination. The results are not specific to USD pricing. When betas are estimated in local-currency terms, the same qualitative pattern holds—positive slopes in DMs and negative slopes in EMs—with slightly stronger contrasts. These findings, reported in Appendix 1.9, confirm that the observed asymmetry reflects real cash-flow co-movement rather than exchange-rate valuation effects.

1.4.2 The Composition of Trade

Exhibit 1.4.1 shows that centrality maps into *opposite* beta slopes across regimes. Exhibit 1.4.2 documents the underlying *composition channel*: as countries become more central, *who* they trade with changes systematically, and this shift differs sharply for DMs and EMs.

Evidence from trade orientation. Figure 1.2 reports the evolution of export orientation by partner type. EMs reoriented dramatically toward other EMs: the EM→EM export share rises from about 20% in 1970 to more than 45% by 2020, with discrete accelerations during the 1990s (regional value chains, China's WTO accession) and the 2000s (South–South agreements). DMs, in contrast, remain predominantly oriented toward other DMs, with only a modest decline from roughly 90% to 70% over the sample.

Figure 1.2: **Trade Orientation by Market Type.**



Notes: Each panel shows the evolution of export orientation for emerging markets (EM, left) and developed markets (DM, right) from 1970–2022. The lines plot, for each bloc, the share of total exports directed to partners classified as EM or DM. For example, the EM→EM line represents the share of exports from EM countries that go to other EMs, relative to EMs' total exports in that year. Similarly, the DM→DM line shows the share of DMs' exports that remain within the DM group. Shares are calculated from bilateral export data in the IMF *Direction of Trade Statistics* (DOTS) by aggregating exports across partner types within each bloc and dividing by the bloc's total exports.

Two quantitative features matter for the finance results that follow. First,

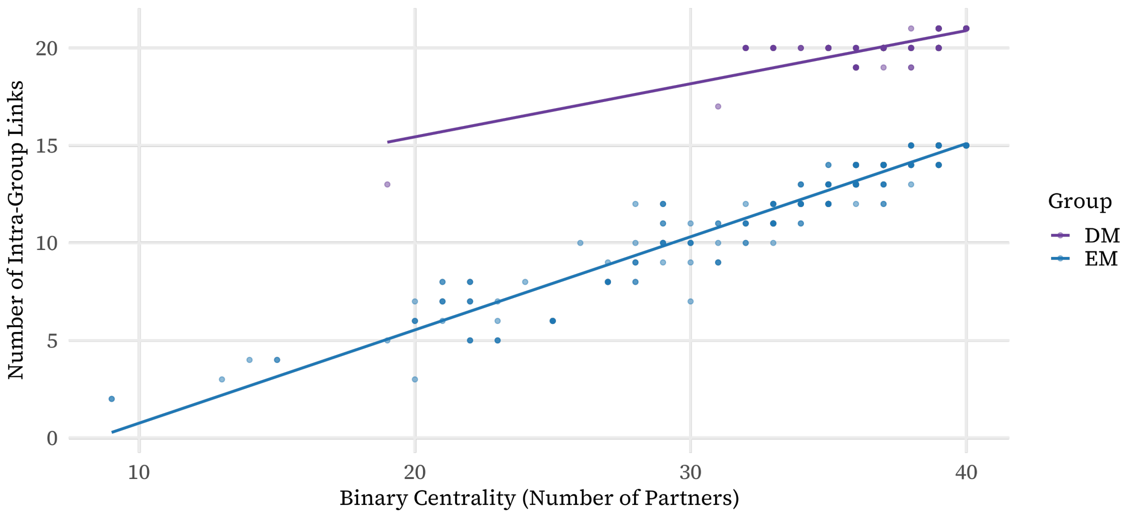


Figure 1.3: **Intra-group links vs. binary centrality.** Each point is a *country–year* observation. The horizontal axis ("binary centrality") is the number of distinct export partners with strictly positive bilateral exports in year t (partners_count). The vertical axis ("intra-group links") is the number of within-block partners in that same year (DM–DM for developed markets; EM–EM for emerging markets). Solid lines are OLS fits estimated separately by block.

for EMs the reorientation is large in levels (a 25 pp rise in the EM to EM share) and concentrated in years when measured global comovement falls. Second, for DMs the mild decline in DM to DM shares masks a powerful intensive-margin deepening of DM–DM trade volumes within global value chains, so that additional centrality largely intensifies ties inside the USD-priced core.

Extensive-margin evidence: who gets connected. The orientation trends aggregate two margins. Figure 1.3 focuses on the *extensive margin*—the number of within-group links—plotted against binary centrality (number of distinct partners). Each dot is a country–year; lines are within-group fits. The EM slope is steep: as EMs add partners, each new link is disproportionately another EM. DMs start from a much denser baseline and add within-DM links more slowly, consistent with an already-connected core.

Table 1.2: **Within-group connectivity and centrality.** This table reports OLS estimates of the regression.

$$\text{Intra-Group Links}_{it} = \alpha + \beta_1 \text{Centrality}_{it} + \beta_2 \text{EM}_i + \beta_3 (\text{Centrality}_{it} \times \text{EM}_i) + \varepsilon_{it}.$$

	Dependent variable: Number of intra-group links
Centrality	0.115^{***}
	(0.009)
EM indicator	-18.967^{***}
	(0.364)
Centrality $\times$ EM	0.343^{***}
	(0.010)
Constant	15.678^{***}
	(0.347)
Observations	1,936
R^2	0.990

Notes: The dependent variable is the number of within-group trade links (DM–DM or EM–EM) for country i in year t, based on positive bilateral export flows. Centrality_{it} is the country's trade-network centrality in year t. EM_i is an indicator for emerging-market countries. Sample covers 1,936 country–year observations, 1970–2022. Standard errors are robust to heteroskedasticity. Significance levels: $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

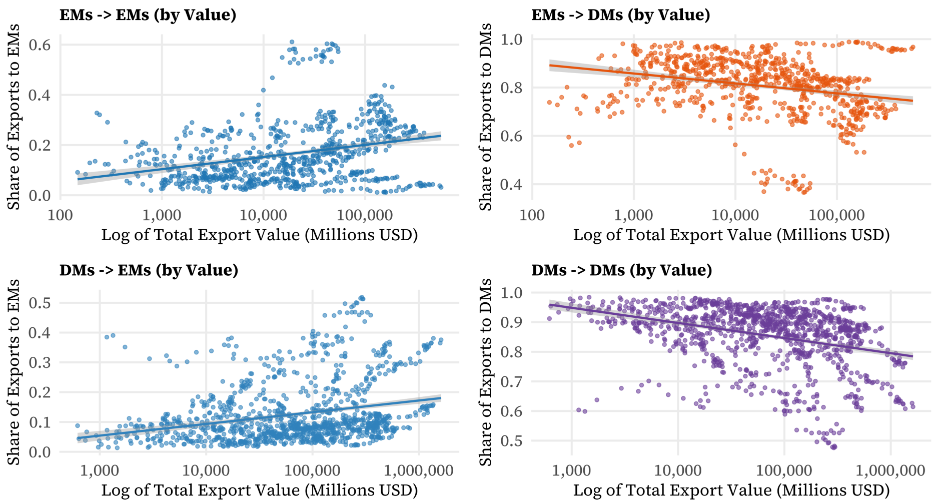
Regression quantification. Table 1.2 formalizes the patterns using:

$$\text{IntraGroupLinks}_{it} = \alpha + \beta_1 \text{Centrality}_{it} + \beta_2 \text{EM}_i + \beta_3 \left(\text{Centrality}_{it} \times \text{EM}_i\right) + \varepsilon_{it}. \tag{1.7}$$

The slope on centrality is strongly positive ($\hat{\beta}_1 = 0.115$) and the EM interaction is large and highly significant ($\hat{\beta}_3 = 0.343$). Evaluated at the interquartile range of centrality, moving from the 25th to the 75th percentile implies roughly *five* extra EM–EM links for EMs, versus *three* extra DM–DM links for DMs. The $R^2 \approx 0.99$ reflects the fact that the number of within-group links is mechanically tied to partner counts, but the *difference* in slopes across regimes identifies composition, not just density.

Intensive margin and scale. While EMs add *more* EM–EM links as they become central, DMs transact *much larger* DM–DM volumes on average (by an order of magnitude in levels by the 2010s). This resolves an appar-

Figure 1.4: **Export orientation by partner type**



Each panel plots the share of exports directed to EMs (left) and to DMs (right) against the exporter's total export value (log scale, millions of USD). The top row shows emerging-market (EM) exporters; the bottom row shows developed-market (DM) exporters. Each dot corresponds to a *country–year observation* constructed from bilateral export flows: for each exporting country i in year t, we compute the share of its total exports going to partners of type $j \in \{\mathrm{EM}, \mathrm{DM}\}$, averaged across all bilateral trade relationships in that year. Fitted lines show within-group OLS trends. The scatter thus summarizes shows all exporter–year observations in the sample from the IMF *Direction of Trade Statistics* (DOTS), 1970–2022.

ent paradox: DMs' centrality gains mostly intensify high-correlation trade inside the global core, whereas EMs' centrality gains expand the set of EM partners with lower correlation to the world portfolio. Consequently, the same increase in centrality has opposite implications for *covariance* with global returns.

Export orientation by partner type. To further illustrate how the composition of trade differs across blocs, Figure 1.4 plots for each country and year the share of exports directed to EMs (left panels) and to DMs (right panels) against the log of total export value. The top row shows EM exporters; the bottom row shows DM exporters.

Two regularities emerge clearly from Figure 1.4. First, among EMs, larger exporters devote a growing share of trade to other EMs (top-left panel) and a declining share to DMs (top-right panel). This pattern reflects the rise of South–South trade and the formation of regional supply chains after the late 1990s. As EMs expand their total exports, their marginal trade partners are increasingly other EMs rather than traditional DM destinations. This reorientation is a key empirical manifestation of the composition channel: greater trade centrality in EMs arises primarily through intra-EM connections.

Second, among DMs, the slope is flatter and less symmetric. DMs' export share to other DMs (bottom-right panel) remains dominant and only slightly declines with scale, while the share to EMs (bottom-left panel) rises modestly at the margin. This asymmetry underscores that DMs' centrality gains reinforce existing linkages within the integrated USD-priced core, whereas EMs' gains reflect new connections within a financially segmented periphery.

Quantitatively, the contrast is striking. For EMs, a one-log-point increase in total exports corresponds to roughly a 10–15 percentage-point increase in the EM-to-EM share and a comparable decline in the EM-to-DM share. For DMs, the same expansion raises the DM-to-EM share by only 2–3 points and reduces the DM-to-DM share by less than 5 points. These differences confirm that the evolution of trade composition is structurally different across regimes: DMs deepen integration within the global core, while EMs reallocate trade within the periphery.

Economically, this composition shift explains why centrality has opposite financial implications across blocs. In DMs, increased centrality tightens exposure to globally priced shocks, since additional trade occurs with other highly integrated economies. In EMs, rising centrality diversifies away from the global SDF, as new trade links primarily involve partners whose cash-flow and currency shocks are locally priced. The export-orientation evidence in Figure 1.4 therefore provides a micro-foundation for the negative centrality–beta slope documented in Exhibit 1 and the partner-composition mechanism formalized in the model that follows.Putt ing the pieces together: in DMs, incremental centrality chiefly *reinforces* integration with already USD-priced partners, so exposure to globally common shocks rises. In EMs, incremental centrality mainly *reorients* trade toward other EMs, diversifying away from the world factor and lowering covariance with it. This composition mechanism provides the concrete, micro-level foundation for the opposite beta–centrality slopes in Exhibit 1.4.1 and sets up the crisis evidence in Exhibit 1.4.3, where we show that central EMs experience *smaller* drawdowns in global stress episodes.

1.4.3 Global Shock Transmission and Trade Centrality

Exhibit 1.4.3 tests whether the composition differences documented in Exhibit 1.4.2 translate into distinct patterns of *shock transmission*. The analysis asks: when the global economy experiences a large adverse event, do more central countries transmit those shocks more strongly or, instead, absorb them more smoothly? The answer depends sharply on market type. For the 2008 Global Financial Crisis (GFC) I compute for each country:

$$\text{Response}_{i,t} = r_{i,t}^{\text{crisis}} - \overline{r}_{i,t-3:t-1},$$

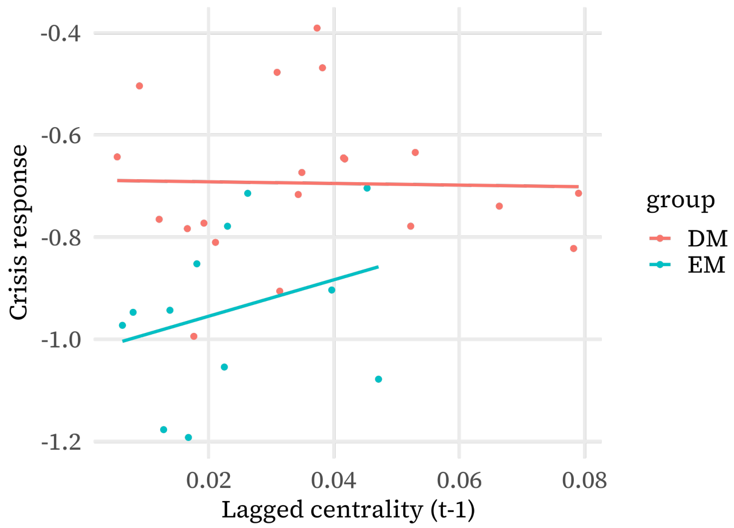
where $r_{i,t}^{\text{crisis}}$ is the annual real equity return in the crisis year and $\overline{r}_{i,t-3:t-1}$ is the mean over the preceding three years. Positive values imply that a country's equity market outperformed its own pre-crisis average; negative values imply deeper drawdowns. I then regress these responses on lagged trade centrality, separately for DMs and EMs:

$$\text{Response}_{i,t} = \alpha_t + \beta_t \,\text{Centrality}_{i,t-1} + \varepsilon_{i,t}.$$

The resulting $\hat{\beta}_t$ measures how much better (or worse) a one-unit rise in pre-crisis centrality predicts performance during global stress.

Figure 1.5 illustrates the 2008 episode. DMs show virtually no slope, confirming that deeper integration within the global core did not insulate them from the GFC. EMs, though hit hard in levels, display a positive cross-

Figure 1.5: **Global Shock Transmission and Trade Centrality.**



Estimated slopes from cross-sectional regressions of crisis responses on lagged centrality. DMs exhibit slopes close to zero across episodes, while EM slopes are positive and significant.

sectional relation: countries such as China, Korea, and Chile—those most central within the EM trade network— recorded smaller equity declines relative to their own pre-crisis performance.

1.4.4 Evolution of Trade–Risk Linkages and Trade Reorientation

The preceding exhibits establish three cross-sectional regularities: (i) centrality is associated with higher global betas in developed markets (DMs) but lower betas in emerging markets (EMs); (ii) EM dividend-growth volatility declines with centrality; and (iii) these effects are consistent across USD- and local-currency (LC) returns. A natural question is whether these patterns have been stable over time or whether they reflect a structural evolution in how trade centrality maps into priced risk. This subsection provides dynamic evidence showing that the β–centrality slope has become progressively more negative for EMs, and that this shift coincides with a systematic reorientation of trade toward other EMs. These patterns provide the empirical bridge between the static cross-sectional findings and the composition mechanism formalized in Section 2.4.

Time-varying global betas. For each country i, I estimate time-varying global risk exposures using ten-year rolling windows of monthly equity excess returns:

$$r_{i,t}^{USD} = \alpha_i + \beta_{i,t}\, r_t^W + \varepsilon_{i,t}, \qquad t = 1970, \ldots, 2022,$$

where r_t^W is the excess return on the world market portfolio. The resulting sequence $\{\beta_{i,t}\}$ provides a smoothed measure of how each country's covariance with global risk evolves. Figure 1.6 displays the resulting time series of average $\beta_{i,t}$ by bloc.

For DMs, average betas are relatively stable around 1, with modest cyclical fluctuations but no trend. For EMs, by contrast, global betas decline sharply after the mid-1990s. The downward drift coincides with the surge in South–South trade, China's WTO accession in 2001, and the creation of multiple regional trade agreements among emerging economies. By the 2010s, average EM betas fall below 0.8, even as EM openness and centrality continue to rise. These patterns suggest that financial exposure to global USD-priced risk has weakened over time, despite deeper trade integration—a first indication of the composition mechanism at work.

Rolling cross-sectional slopes. To examine how the relationship between trade centrality and global risk exposure evolves, I estimate for each year t the cross-sectional regression

$$\beta_{i,t} = \alpha_t + b_t\, \text{Centrality}_{i,t} + \varepsilon_{i,t}, \qquad i = 1, \ldots, N, \tag{1.8}$$

where $\beta_{i,t}$ is country i's equity beta with respect to the world portfolio, estimated from a rolling ten-year window of monthly excess returns. The

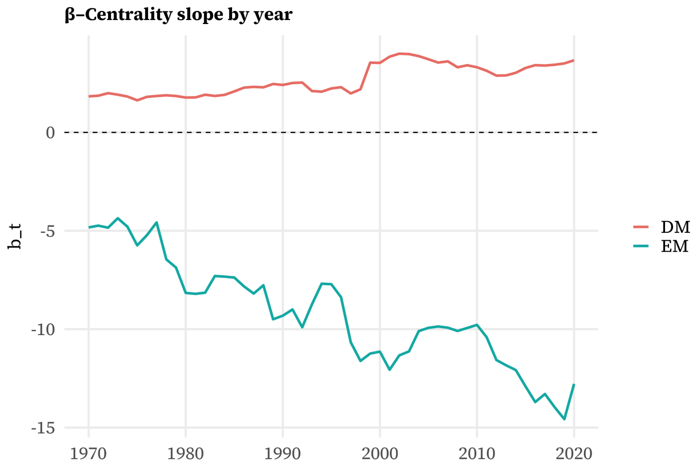
Figure 1.6: **Time-varying betas**

Average rolling-window equity betas with respect to the world portfolio. Developed-market betas (red); emerging-market betas (blue)

coefficient b_t captures the contemporaneous cross-sectional slope between centrality and global risk exposure in year t. Figure 1.7 plots the sequence of estimated slopes b_t separately for DMs and EMs.

The results reveal a striking divergence in dynamics across regimes. For DMs (red line), the β–centrality slope remains positive and stable from the early 1970s through 2022, consistent with a persistent mapping between trade embeddedness and exposure to the USD-priced global factor. For EMs (blue line), the slope is negative throughout and becomes steadily more negative over time. The downward trend accelerates after the mid-1990s, precisely when EMs experienced large trade liberalizations, regional trade agreements, and China's WTO entry. By the 2010s, the slope stabilizes around -10 to -15, implying that countries one standard deviation more central in the EM trade network exhibit roughly a ten-percentage-point lower global beta than otherwise comparable EMs. This dynamic evidence suggests that the cross-sectional asymmetry is not static but deepened endogenously as EMs integrated with one another.

Figure 1.7: **Evolution of the β–centrality slope by regime.**

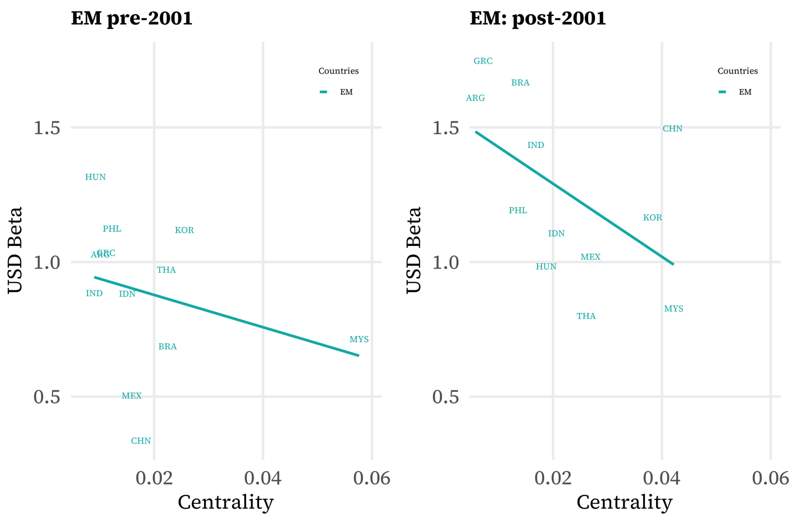


Estimated yearly slopes b_t from cross-sectional regressions of country betas on trade centrality (ten-year rolling windows). The DM slope (red); EM slope (blue)

Structural break and pre-/post-2001 comparison. The dynamic patterns in Figure 1.7 suggest that the negative β–centrality slope for EMs intensified after 2000, coinciding with the rise of South–South trade. To illustrate this structural shift more directly, Figure 1.8 compares the cross-sectional relation between USD betas and trade-network centrality for emerging markets before and after 2001. The left panel plots the relation using average values over the 1980–2000 period; the right panel uses averages from 2001–2020.

In the pre-2001 period, the β–centrality slope for EMs is only mildly negative and economically small. Most EMs were still relatively peripheral in the global trade network, and increases in centrality often reflected greater trade with developed-market partners. By contrast, in the post-2001 period, the slope steepens considerably. Countries that became more central—such as China, Korea, and Malaysia—show markedly lower global betas, while those whose centrality rose less—such as Argentina, Brazil, and Greece—retain higher betas. This divergence captures the essence of the composition channel: the nature of trade connections, not simply their

Figure 1.8: **Emerging-market betas and trade centrality before and after 2001.**



Left panel: 1980–2000; right panel: 2001–2020. The slope of USD betas on trade-network centrality becomes substantially more negative in the post-2001 period, coinciding with China's WTO entry.

volume, determines whether greater centrality amplifies or attenuates exposure to the USD-priced world factor.

The timing of this shift is economically meaningful. The early 2000s mark the acceleration of intra-EM integration, including China's WTO accession (2001), the creation of ASEAN+3 frameworks, and the proliferation of South–South trade agreements. These institutional changes reoriented EM trade from the DM core toward other EMs, lowering covariance with the global pricing kernel even as network centrality increased. The figure therefore provides visual evidence that the negative slope between betas and centrality in EMs is a post-liberalization phenomenon tied to the restructuring of global trade networks.

Trade reorientation and changes in betas. To connect this evolution to trade composition, I next relate changes in global risk exposure to changes in countries' trade orientation. The basic idea is that if EMs' declining betas reflect a structural shift toward trading with other EMs, then countries that reoriented their exports or imports more strongly toward EM partners should exhibit smaller increases—or larger declines—in global betas over time. Formally, I estimate

$$\Delta\beta_i = \alpha + \gamma\, \Delta\text{Share}_i^{EM\to EM} + \varepsilon_i, \qquad \Delta\beta_i = \beta_i^{post} - \beta_i^{pre}, \tag{1.9}$$

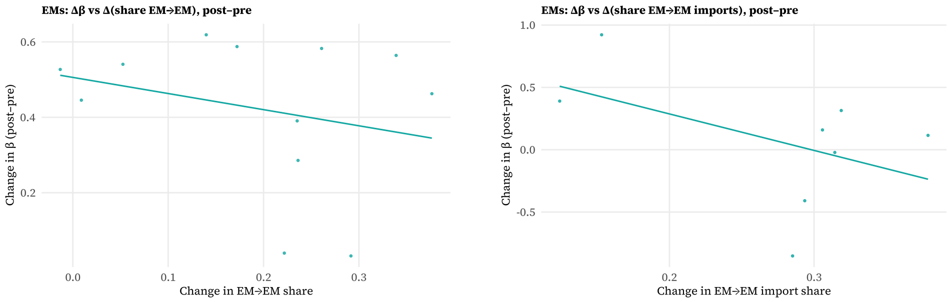
where $\Delta\beta_i$ is the change in country i's global beta between pre- and post-periods, and $\Delta\text{Share}_i^{EM\to EM}$ is the change in its export share directed toward other EMs. The post-period begins in 2001, capturing the phase of accelerated South–South trade following China's WTO accession; the pre-period covers 1980–2000.

Figure 1.9 plots the relationship between the change in EM–EM export share and the change in global beta for emerging markets. The slope is negative and significant: EMs that increased their export orientation toward other EMs experienced smaller increases—or larger declines—in global betas after 2001. For a median EM, a ten-percentage-point rise in EM–EM export share corresponds to a four- to five-percentage-point smaller change in global beta. The result implies that EM–EM export reallocation reduces exposure to the USD-priced world factor even as overall trade openness rises.

A symmetric pattern holds for imports. Figure 1.9 shows that EMs that reoriented imports toward other EMs also saw smaller changes in global betas. Together, these results indicate that both export and import reorientation dampen EM exposure to global risk. The pattern is consistent with the idea that intra-EM trade links transmit locally priced shocks that are imperfectly correlated with the global USD SDF. Hence, as EMs strengthen these intra-bloc linkages, their aggregate exposure to globally priced volatility declines.

Interpretation. The rolling-beta evidence and the trade-reorientation regressions together document a clear dynamic pattern: as emerging markets

Figure 1.9: **Trade reorientation and changes in EM global betas.**



(a) Change in EM betas vs. change in EM→EM **export** share.

(b) Change in EM betas vs. change in EM→EM **import** share.

Each panel plots the cross-sectional relation between the change in EM countries' global equity betas ($\Delta\beta_i$) and the change in their intra-EM trade shares, for exports (panel a) or imports (panel b). Each dot represents a single emerging market; the fitted lines show within-sample OLS slopes. $\Delta\beta_i$ denotes the change in the USD CAPM beta between the pre-2001 and post-2001 periods, and $\Delta\text{Share}_i^{EM\to EM}$ (or $\Delta\text{Share}_i^{EM\leftarrow EM}$) denotes the change in the share of exports (imports) directed to other EM partners over the same window.

have become more central in global trade, their *composition* of partners has shifted toward other EMs, and this shift has reduced their covariance with the global pricing kernel. Developed markets, by contrast, have retained trade partners within the USD-priced core, leaving their positive β–centrality slopes stable over time. This divergence in dynamics motivates the structural mechanism in Section 2.4: a single global investor with a USD-denominated SDF but bloc-specific mappings from trade centrality to priced exposure. The parameter γ_m in the model formalizes the composition channel uncovered here—positive in DMs, negative in EMs—capturing how the reorientation of trade links reverses the direction of global risk transmission.

1.5 A Reduced-Form Model of Trade Centrality and Risk Exposure

This section develops a reduced-form framework that rationalizes the empirical evidence documented in Section 1.4. The goal is to provide a tractable mechanism that links countries' network position to their exposure to global risk under segmented pricing. The model embeds trade-network centrality into an otherwise standard stochastic-discount-factor (SDF) environment and delivers closed-form expressions for equity betas, dividend volatility, and the opposite-signed centrality slopes observed across developed and emerging markets.

1.5.1 Economic Environment and Intuition

Consider a world with a representative USD investor who prices all assets. The USD is the global numéraire, and markets are complete in dollars. Differences across blocs arise not from distinct SDFs but from how each country's cash flows load on USD-priced shocks. Trade-network centrality, denoted $C_{i,t}$, determines the composition of these loadings: more central countries trade more intensely with the global core, altering the weight on globally priced versus locally priced shocks. The model builds directly on the empirical regularities documented above. It formalizes the observation that network position modifies how global shocks are transmitted through cash flows and asset prices. Conceptually, trade centrality $C_{i,t}$ plays the same role as an "integration index" in international asset-pricing models, but unlike traditional openness measures, it embeds information about *who* countries trade with. Because network centrality is determined by the structure of bilateral trade rather than its aggregate size, two countries with identical trade-to-GDP ratios can have very different exposures to globally priced risk, depending on whether they are DMs or EMs.

The central mechanism of the model can be understood as the propagation of a *global network shock*—for instance, an increase in tariffs, trade barriers, or input–output disruptions—through the international trade and production network. Such a shock affects both supply and demand conditions globally: higher input costs depress production and raise consumer prices, while retaliatory trade policies reduce external demand. Kalemli-Ozcan et al. (2025) shows these shocks transmit through the structure of the world input–output matrix, generating heterogeneous effects across countries depending on their position in the global network and on their policy responses. A country that is more central or upstream in global value chains is exposed to larger fluctuations in intermediate and final–goods demand, while peripheral countries experience muted but more idiosyncratic effects.

This mechanism implies that dividend and cash–flow dynamics are *heterogeneous across countries*. Even when all countries face the same underlying global disturbance, the resulting changes in profits, export revenues, and output differ systematically according to network position and the composition of trade partners. This heterogeneity in real cash flows provides the foundation for cross–country differences in risk exposures and betas. In the model, these differences are captured by allowing each country's dividend process to load on a global component that is scaled by trade–network centrality.

The idea that global shocks generate heterogeneous country exposures resonates with the broader literature on international risk transmission. Rey (2013); Miranda-Agrippino and Rey (2019) emphasize that in the presence of a *global financial cycle*, countries differ in their exposures and responses to common shocks because of heterogeneity in balance sheets, market structure, and policy regimes. Similarly, network–based analyses show that the propagation of global trade or production shocks is far from uniform: countries that are more central in the world input–output network experience larger "avalanche" effects when tariffs, supply disruptions, or demand shocks occur (Chang et al., 2022a). These findings motivate treating trade–network position as a structural state variable that links a country's cash–flow process to the global stochastic discount factor. Differences in economic development also shape how trade centrality translates into risk. Empirical work shows that poorer economies are typically more volatile because they specialize in riskier sectors and are more exposed to large idiosyncratic shocks, whereas volatility declines as GDP rises (Koren and Tenreyro, 2007; di Giovanni and Levchenko, 2012). The interaction between openness, trade centrality, and volatility is therefore state dependent. When global volatility or disruption risk is high, trade and production linkages transmit shocks strongly, but in tranquil periods spillovers are muted. Evidence from firm- and country-level data shows that large shocks propagate disproportionately through input–output networks and multinational trade links (Boehm et al., 2019; Barrot and Sauvagnat, 2016; di Giovanni et al., 2018). In such episodes, countries that are more central in the global trade network experience amplified co-movements in output, earnings, and asset prices, whereas in low-volatility states centrality has negligible effects. This asymmetry helps explain why the impact of trade centrality on systematic risk is positive in high-volatility regimes but small in stable times. From the perspective of dividends, the asymmetry arises because the sources of cash-flow risk differ systematically across blocs. In developed markets, firms are tightly integrated into global value chains and their earnings depend heavily on globally priced demand and financial conditions. A rise in trade centrality therefore increases the share of revenues exposed to the common global component, amplifying covariance with the USD pricing kernel and raising systematic risk (Herskovic, 2018; Chang et al., 2022b). In emerging markets, by contrast, rising centrality often reflects expanding trade with other segmented or regionally integrated economies that transact largely in local currencies or operate under idiosyncratic shocks (Kalemli-Ozcan et al., 2025; di Giovanni and Hale, 2021). As these linkages deepen, the composition of trade partners shifts toward peers whose demand and funding conditions are only weakly correlated with the global factor. The result is a form of network diversification: dividend growth becomes less synchronized with global volatility, and the covariance of cash flows with the world SDF declines. Rising centrality in emerging markets reallocates exposure from globally priced to locally priced risk, reducing the volatility of aggregate dividend growth and hence lowering measured betas. Empirically, this mechanism is consistent with the decline in dividend-growth volatility among central emerging markets documented in the data and with the composition channel formalized in the model, where the sign of the centrality coefficient is positive in developed markets and negative in emerging ones.

The structure of production and outsourcing further contributes to these differences. In developing economies, integration often occurs through the import of intermediate goods and the outsourcing of production stages. Reductions in input tariffs increase productivity and reallocate revenue exposure from domestic to foreign demand (Amiti and Konings, 2007). Consequently, firms with higher import content exhibit dividend streams that embed a larger foreign component and a smaller home component, altering their covariance with global shocks (di Giovanni and Hale, 2021; Chang et al., 2022b; Herskovic, 2018). At the aggregate level, this composition effect implies that trade centrality lowers exposure to the global factor when foreign partners are segmented and raises it when partners are globally priced.

When a country remains inward-oriented, its dividends load primarily on home demand and domestic productivity shocks. When it exports and imports extensively, its dividends load on a combination of home and foreign shocks determined by partner composition and input intensity. Greater export diversification reduces dependence on any single domestic shock and shifts risk toward foreign shocks. Because foreign shocks are a weighted average across markets, trade integration can reduce dividend volatility in emerging economies by diversifying away from purely domestic shocks toward less correlated foreign ones, while in developed markets it increases co-movement with the global factor through tighter integration with the core (di Giovanni and Levchenko, 2009; Koren and Tenreyro, 2007; di Giovanni et al., 2018). This logic highlights that the effect of centrality on risk is inherently compositional: it depends not only on how much a country trades, but with whom, and how those trade links map into the covariance structure of dividend growth and the global stochastic discount factor.

1.5.2 Global Pricing Kernel

The (log) USD SDF follows

$$-\log m_{US,t+1} = \alpha + \chi z_t^i + \tau z_t^w + \sqrt{\gamma z_t^i}\, u_{t+1}^i + \sqrt{\kappa z_t^i}\, u_{t+1}^g + \sqrt{\delta z_t^w}\, u_{t+1}^w + \sqrt{\omega z_t^w}\, u_{t+1}^c + \sqrt{\xi z_t^w}\, u_{t+1}^x. \tag{1.10}$$

The innovations $(u^i, u^g, u^w, u^c, u^x)$ are independent standard normally distributed random variables representing, respectively, local growth, global growth, global volatility, composition (network), and an additional orthogonal global shock.[1] The final term increases overall global variance but—because it is not scaled by $C_{i,t}$—does not modify the centrality mechanism. The USD pricing kernel in equation (1.10) captures how the representative global investor prices risk. The kernel includes both local and global sources of uncertainty and, crucially, a "composition" shock u_{t+1}^c whose price depends on trade centrality. This shock represents changes in the common component of globally priced cash-flow risk associated with shifts in the composition of trade partners. The final term, u_{t+1}^x, is an additional orthogonal world shock that raises the overall variance of the world factor but does not depend on $C_{i,t}$. It can be interpreted as a residual "global equity" shock common to all markets. Including u_{t+1}^x ensures the model remains rich enough to accommodate other globally correlated movements without affecting the key comparative statics with respect to centrality. Equations (1.18)–(1.20) can be interpreted as elasticities of valuation with respect to the underlying state variables. The coefficient B captures the sensitivity of valuation ratios to domestic conditions, C measures sensitivity to global volatility shocks, and D measures how valuation changes with a country's network position. The key comparative static is that D is proportional to γ_m, so the sign of the valuation–centrality relation flips across regimes. In DMs, higher centrality raises expected price–dividend ratios (greater exposure to priced global risk); in EMs, higher centrality lowers them (greater segmentation and lower correlation with the global kernel).

[1]The extra term u^x parallels the global "equity" shock introduced in Brusa et al. (2014a), which affects all countries symmetrically but does not alter network scaling.

1.5.3 State Variables

Three persistent state variables drive the economy:

$$z_{t+1}^{i} = (1-\phi_i)\theta_i + \phi_i z_t^i + \varepsilon_{t+1}^{i}, \quad \text{(local-growth state)}, \tag{1.11}$$

$$z_{t+1}^{w} = (1-\phi_w)\theta_w + \phi_w z_t^w + \varepsilon_{t+1}^{w}, \quad \text{(global-volatility state)}, \tag{1.12}$$

$$C_{i,t+1} = (1-\phi_c)\theta_c + \phi_c C_{i,t} + \varepsilon_{i,t+1}^{c}, \quad \text{(trade-network centrality)}. \tag{1.13}$$

All shocks are Gaussian and orthogonal. The three state variables summarize all relevant sources of persistence in the system. The local state z_t^i captures productivity and demand shocks specific to country i. The global state z_t^w captures the intensity of globally priced volatility or growth shocks that are common across countries. The centrality variable $C_{i,t}$ evolves much more slowly, reflecting gradual changes in trade architecture—such as the formation of regional value chains or entry into global trading blocs.

Empirically, centrality is extremely persistent, consistent with the slow reallocation of trade partners:

Table 1.3: **Persistence of trade-network centrality.**

	DM	**EM**
θ	0.0002^{***}	0.0001^{***}
	(0.0000)	(0.0000)
ϕ	0.9848^{***}	0.9875^{***}
	(0.0086)	(0.0115)
Observations	961	459

Notes: GMM estimates of $C_{i,t+1} = (1 - \phi_c)\theta_c + \phi_c C_{i,t} + \varepsilon_{i,t+1}^{c}$, estimated separately for DMs and EMs. Standard errors (in parentheses) are robust to heteroskedasticity.

Empirically, the very high persistence of $C_{i,t}$ is a crucial feature: it implies that differences in exposure across countries are structural rather than cyclical. In this sense, centrality acts more like a structural parameter of the international equilibrium than a short-run state variable. z_t^w controls the intensity of globally priced shocks, z_t^i captures local fundamentals, and $C_{i,t}$ encodes the country's position in the trade network. Centrality evolves slowly but governs how changes in z_t^w and z_t^i translate into exposure to USD-priced risk.

1.5.4 Dividend Dynamics and Centrality Scaling

The dividend process translates macroeconomic states into cash-flow risk faced by investors. Local and global shocks jointly determine dividend growth, but only the global component associated with u_{t+1}^c is scaled by centrality. This is the simplest functional form that reproduces the empirical finding that the strength of exposure to globally priced shocks varies systematically with network position. The square-root specification ensures homoscedasticity in logs and delivers a closed-form solution for price–dividend ratios under standard approximations. Country i's real dividend growth is given by

$$\begin{aligned}\Delta d_{i,t+1} = \mu_D + \psi z_t^i + \psi_w z_t^w + \sigma_D \sqrt{z_t^i}\, u_{t+1}^i + \sigma_D^g \sqrt{z_t^i}\, u_{t+1}^g + \sigma_D^w \sqrt{z_t^w}\, u_{t+1}^w \\ + \sqrt{\sigma_D^c + \gamma_m C_{i,t}}\, \sqrt{z_t^w}\, u_{t+1}^c + \sigma_D^x \sqrt{z_t^w}\, u_{t+1}^x . m \in \{EM, DM\}\end{aligned} \tag{1.14}$$

The final two terms represent global shocks. The key object is the square-root term in front of u_{t+1}^c: it embeds trade centrality in the variance of the global composition shock.

Data construction and reduced-form facts for dividends

Constructing dividends from prices and total returns. Let $R_{t+1}^{tr} \equiv \dfrac{P_{t+1} + D_{t+1}}{P_t}$ denote the gross total return and $R_{t+1}^p \equiv \dfrac{P_{t+1}}{P_t}$ the gross price return. Then the dividend yield at $t+1$ satisfies $\dfrac{D_{t+1}}{P_t} = R_{t+1}^{tr} - R_{t+1}^p$, and the dividend yield at t is $\dfrac{D_t}{P_{t-1}} = R_t^{tr} - R_t^p$. Hence the gross dividend growth rate is identified from observables by

$$\Delta d_{t+1} \equiv \frac{D_{t+1}}{D_t} - 1 = \left(\frac{R^{tr}_{t+1} - R^{p}_{t+1}}{R^{tr}_t - R^p_t}\right) R^p_t - 1, \tag{1.15}$$

which we compute in real terms (deflating P_t and D_t) and in both USD and LC units. Equation (1.15) underlies all dividend facts reported below and aligns the empirical objects with the model's $\Delta d_{i,t+1}$ in (1.14).

Define for each country i its unconditional volatility of real dividend growth $\sigma_i \equiv std(\Delta d_{i,t})$ over the sample. Figure 1.10 plots σ_i^2 against average trade-network centrality. Among EMs the slope is sharply negative (and statistically significant), while it is flat to weak in DMs: greater centrality is associated with *lower* dividend-growth volatility in EMs. This fact supports modeling dividend growth with a common global block and a centrality-dependent composition term as in (1.14), and it maps directly into the model's implication $\partial \operatorname{Var}_t(\Delta d_{i,t+1})/\partial C_{i,t} = \gamma_m z^w_t$ (negative in EMs, cf. (1.24)).

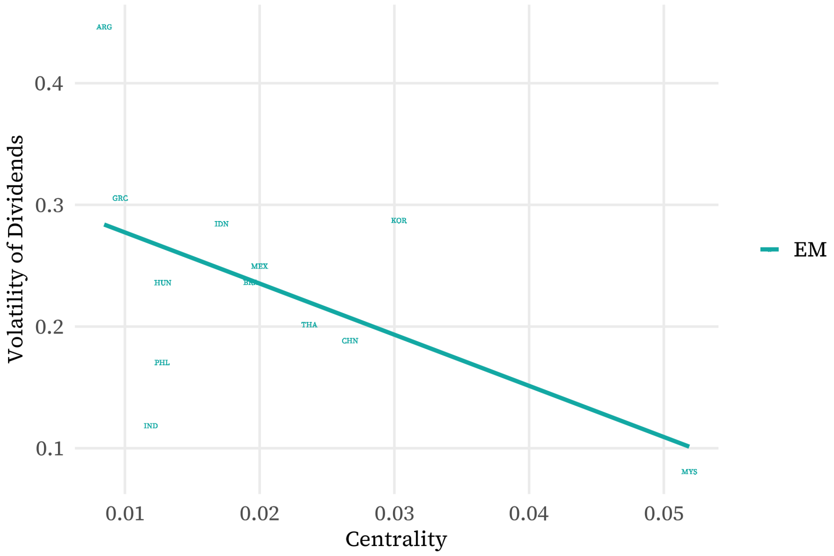
Dividend growth comovement with the world factor. Let Δd^W_t denote world dividend growth (market-value weighted over the large advanced economies). To quantify cross-bloc comovement, estimate

$$\Delta d_{i,t} = \alpha + \beta^{DM}\, \Delta d^W_t + \beta^{EM}\, \mathbf{1}\{i \in EM\} \cdot \Delta d^W_t + \gamma\, \mathbf{1}\{i \in EM\} + \varepsilon_{i,t}. \tag{1.16}$$

Consistent with long-run-risk evidence, the DM loading $\hat{\beta}^{DM}$ is positive: DM dividends co-move with the world dividend factor. The EM differential is negative: $\hat{\beta}^{EM} < 0$, implying a substantially weaker (and in many samples statistically lower) comovement of EM dividends with Δd^W_t relative to DMs. This bloc contrast motivates the model's bloc-specific composition parameter γ_m in (1.14) and the opposite-signed β–centrality slopes.

The construction in (1.15) ensures that the empirical $\Delta d_{i,t}$ is the object priced in the reduced-form model. The EM negative volatility–centrality slope (Figure 1.10) is the variance analogue of the model's composition channel: with $\gamma_{EM} < 0$, higher $C_{i,t}$ reduces the contribution of the priced

Figure 1.10: **Volatility of dividend growth vs. centrality (EM).**



Each point corresponds to a single emerging-market (EM) country, where the vertical axis reports the time-series standard deviation of real dividend growth in USD terms, and the horizontal axis reports the country's average trade-network centrality over the sample. The fitted line shows the within-group OLS relation between volatility and centrality.

Table 1.4: **Dividend growth on world dividend growth.**

	Δd_t^W (DM base)	EM $\times$ Δd_t^W	EM dummy
Coefficient	1.233***	-2.587***	0.044**
(Std. error)	(0.074)	(0.131)	(0.020)
Intercept	-0.008 (0.011)		
Observations	1,178		
R^2	0.273		

Notes: Dependent variable is country i's real dividend growth $\Delta d_{i,t}$ (USD). Regressors are world dividend growth Δd_t^W, the EM interaction, and an EM dummy. Standard errors in parentheses.

global block to dividend risk ((1.24)). The regression (1.16) pins down relative comovement of dividends with the global component across blocs; the positive DM loading and negative EM differential are the cash-flow counterpart of the opposite-signed exposure slopes for returns. Together these reduced-form dividend facts justify modeling dividend growth with a common global factor and centrality-scaled composition exposure and provide direct empirical discipline for the sign of γ_m used in the quantitative exercises.

1.5.5 Pricing and the Price–Dividend Ratio

No-arbitrage requires $E_t[m_{US,t+1}R_{i,t+1}] = 1$. Log-linearizing returns,

$$r_{i,t+1} \approx k_0 + k_1(pd_{i,t+1} - pd_{i,t}) + \Delta d_{i,t+1}, \qquad pd_{i,t} \equiv \log \frac{P_{i,t}}{D_{i,t}}.$$

Assume an affine pd function,

$$pd_{i,t} = A + Bz_t^i + Cz_t^w + DC_{i,t}. \tag{1.17}$$

Solving yields

$$B = \frac{-\chi - \frac{1}{2}\left[(\sqrt{\gamma} - \sigma_D)^2 + (\sqrt{\kappa} - \sigma_D^g)^2\right]}{k_1(\phi_i - 1)}, \tag{1.18}$$

$$C = \frac{-\tau - \frac{1}{2}\left[(\sqrt{\delta} - \sigma_D^w)^2 + (\sqrt{\omega} - \sigma_D^c)^2 + (\sqrt{\xi} - \sigma_D^x)^2\right]}{k_1(\phi_w - 1)}, \tag{1.19}$$

$$D = \frac{\gamma_m(\sqrt{\omega} - \sigma_D^c)}{k_1(\phi_c - 1)}. \tag{1.20}$$

The detailed solution is in Appendix 1.11.The new global shock u^x adds only to the quadratic term in C; D—the coefficient linking pd to centrality—is unaffected because u^x is not scaled by $C_{i,t}$. The coefficients B, C, and D summarize how valuation ratios respond to the three state variables—local growth z_t^i, global volatility z_t^w, and trade-network centrality $C_{i,t}$. Each term embeds a covariance between the SDF and the country's cash-flow shocks. Specifically, the quadratic differences $(\sqrt{\gamma} - \sigma_D)$, $(\sqrt{\kappa} - \sigma_D^g)$, $(\sqrt{\delta} - \sigma_D^w)$, and $(\sqrt{\omega} - \sigma_D^c)$ arise from the difference in exposure of the SDF and the dividend process to the same underlying shocks. These differences are the reduced-form representation of $\mathrm{Cov}(m_{US,t+1}, R_{i,t+1})$ in the log-linearized pricing equation: a larger gap between SDF and dividend loadings implies a stronger covariance with the pricing kernel and therefore a higher expected return.

The first two coefficients, B and C, capture how domestic and global shocks affect valuation across all countries. The new coefficient D links valuation directly to trade-network centrality and introduces the composition channel into pricing. Because D is proportional to $\gamma_m(\sqrt{\omega} - \sigma_D^c)$, its sign depends on the bloc-specific composition parameter γ_m: for developed markets ($\gamma_{DM} > 0$), higher centrality raises exposure to the USD-priced global factor, increasing the price–dividend ratio; for emerging markets ($\gamma_{EM} < 0$), higher centrality reallocates risk toward locally priced shocks, reducing global exposure and lowering the price–dividend ratio. Thus, equations (18)–(20) make explicit that the centrality term in the model operates through the covariance structure between the country's dividend growth and the global pricing kernel, not through an independent pricing factor.

1.5.6 Expected Returns and Innovations

Using (1.17) in the Campbell–Shiller decomposition,

$$E_t[r^e_{i,t+1}] = \bar{r}_i + \kappa_i z^i_t + \kappa_W z^w_t + \kappa_C C_{i,t}, \tag{1.21}$$
$$\kappa_i = k_1 B(\phi_i - 1) + \psi, \quad \kappa_W = k_1 C(\phi_w - 1) + \psi_w, \quad \kappa_C = k_1 D(\phi_c - 1).$$

Return innovations follow directly from (1.14) and (1.10); their composition term features $\sqrt{\sigma_D^c + \gamma_m C_{i,t}}$, generating centrality-dependent exposure to global risk.

1.5.7 Equity Betas and Centrality Slopes

Let the USD world return innovation be

$$r^W_{t+1} - E_t r^W_{t+1} = \sigma^W_w u^w_{t+1} + \sigma^W_c u^c_{t+1} + \sigma^W_x u^x_{t+1}.$$

Projecting country returns onto the world factor yields

$$\beta_i^{USD} = \frac{\sigma_D^w \sigma_w^W + \sqrt{\sigma_D^c + \gamma_m C_i}\,\sigma_c^W + \sigma_D^x \sigma_x^W}{(\sigma_w^W)^2 + (\sigma_c^W)^2 + (\sigma_x^W)^2}. \tag{1.22}$$

Differentiating with respect to centrality,

$$\frac{\partial \beta_i^{USD}}{\partial C_i} = \frac{1}{2} \frac{\gamma_m}{\sqrt{\sigma_D^c + \gamma_m C_i}} \cdot \frac{\sigma_c^W}{(\sigma_w^W)^2 + (\sigma_c^W)^2 + (\sigma_x^W)^2}. \tag{1.23}$$

The additional shock u^x rescales the denominator but does not affect the sign of the slope. An analogous result holds for LC betas, with $(\sigma_w^{W,LC}, \sigma_c^{W,LC}, \sigma_x^{W,LC})$ replacing the USD coefficients.

The model also delivers direct implications for the volatility of fundamentals. Differentiating the conditional variance of dividend growth with respect to centrality yields $\partial \operatorname{Var}_t(\Delta d_{i,t+1})/\partial C_{i,t} = \gamma_m z_t^w$, implying that volatility increases with centrality in DMs and decreases in EMs. This pattern mirrors the empirical results of Fact 3: dividend-growth volatility is essentially flat across central DMs but declines sharply across central EMs. Because both betas and dividend volatility depend on γ_m with the same sign, the model jointly explains the covariance and variance patterns observed in the data.

1.5.8 Dividend Volatility and Centrality

From (1.14),

$$\operatorname{Var}_t(\Delta d_{i,t+1}) = (\sigma_D^2 + \sigma_D^{g\,2}) z_t^i + [(\sigma_D^w)^2 + (\sigma_D^x)^2 + \sigma_D^c + \gamma_m C_{i,t}] z_t^w,$$

so that

$$\frac{\partial \operatorname{Var}_t(\Delta d_{i,t+1})}{\partial C_{i,t}} = \gamma_m z_t^w \begin{cases} > 0 & \text{if } \gamma_{DM} > 0, \\ < 0 & \text{if } \gamma_{EM} < 0, \end{cases} \tag{1.24}$$

consistent with flat volatility slopes in DMs and steeply negative ones in EMs (Fact 3).

1.5.9 Mapping Parameters to Data

Linearizing (1.22) around $(\bar{C}, \bar{z}^w)$ yields the empirical mapping

$$\hat{\gamma}_m \approx \frac{1}{2} \frac{\gamma_m}{\sqrt{\sigma_D^c + \gamma_m \bar{C}}} \cdot \frac{\sigma_c^W}{(\sigma_w^W)^2 + (\sigma_c^W)^2 + (\sigma_x^W)^2}, \tag{1.25}$$

and analogously for LC slopes. Equation (1.25) identifies the sign of γ_m and links its magnitude to observed β–centrality slopes. The fitted values $\gamma_{DM} = 0.52$ and $\gamma_{EM} = -0.85$ jointly match (i) mean β by bloc, (ii) mean and volatility of dividend growth, and (iii) the empirical $\beta \sim C$ gradients (Table 1.5).

Table 1.6 reports the calibrated parameters from the GMM estimation, which jointly targets the levels and slopes of the model-implied moments in USD. The estimation matches mean betas, their sensitivity to trade centrality, and the first two moments of dividend growth and returns across developed and emerging markets. The exposure of dividends to global volatility shocks is moderate (σ_D^W), while the baseline network variance σ_D^C captures residual local fluctuations in dividend growth. This divergence reflects the model's core mechanism—developed markets are integrated along the intensive margin of trade, whereas emerging markets remain segmented along the extensive margin. The global volatility risk price is π_W, and the price of network-composition risk is π_C. Substituting these values into the model's closed-form expressions yields mean USD betas of 1.20 for developed markets and 0.94 for emerging markets, closely reproducing the empirical pattern documented in Section **??**. Together, these estimates quantitatively anchor the model's interpretation: trade-network structure is a first-order determinant of how national equity markets load on global risk.

1.5.10 Comparative Statics and Dynamic Responses

A permanent increase in centrality ($\Delta C > 0$) raises (lowers) exposure to USD-priced risk in DMs (EMs). Because $C_{i,t}$ evolves as in (1.13), the

Table 1.5: **Model fit, empirical moments, and theoretical counterparts (USD).**

Moment	Data	GMM	Theory
Mean β (USD), DM	1.029	1.201	$\dfrac{\sigma_D^w \sigma_w^W + \sqrt{\sigma_D^c + \gamma_{DM} C}\, \sigma_c^W}{(\sigma_w^W)^2 + (\sigma_c^W)^2}$
Mean β (USD), EM	1.126	0.936	$\dfrac{\sigma_D^w \sigma_w^W + \sqrt{\sigma_D^c + \gamma_{EM} C}\, \sigma_c^W}{(\sigma_w^W)^2 + (\sigma_c^W)^2}$
Volatility of dividend growth, EM	0.345	0.340	$\sqrt{(\sigma_D^w)^2 + \sigma_D^c + \gamma_{EM} C}$
Volatility of dividend growth, DM	0.235	0.250	$\sqrt{(\sigma_D^w)^2 + \sigma_D^c + \gamma_{DM} C}$
Mean return, DM	0.088	0.221	$\sigma_D^w \pi_w + \sqrt{\sigma_D^c + \gamma_{DM} C}\, \pi_c$
Mean return, EM	0.111	0.167	$\sigma_D^w \pi_w + \sqrt{\sigma_D^c + \gamma_{EM} C}\, \pi_c$
Slope[$\beta \sim C$], DM	2.662	2.661	$\dfrac{1}{2} \dfrac{\gamma_{DM}}{\sqrt{\sigma_D^c + \gamma_{DM} C}} \dfrac{\sigma_c^W}{(\sigma_w^W)^2 + (\sigma_c^W)^2}$
Slope[$\beta \sim C$], EM	-9.193	-9.190	$\dfrac{1}{2} \dfrac{\gamma_{EM}}{\sqrt{\sigma_D^c + \gamma_{EM} C}} \dfrac{\sigma_c^W}{(\sigma_w^W)^2 + (\sigma_c^W)^2}$
γ_{DM}		0.516	
γ_{EM}		-0.854	

Notes: Data are country-group moments (USD). "GMM" reports moments implied by the estimated model. Theory column lists closed-form counterparts evaluated at group means where applicable. Parameters: σ_D^w and σ_D^c scale dividend exposure to the world (w) and composition (c) shocks; σ_w^W, σ_c^W are world-return loadings; π_w, π_c are corresponding prices of risk; C is average trade-network centrality by bloc; γ_{DM} and γ_{EM} are composition parameters.

impulse response of the covariance with the world return is

$$\Delta\, \mathrm{Cov}_h(r_i, r^W) = \mathrm{Var}(r^W) \left[\frac{1}{2} \frac{\gamma_m}{\sqrt{\sigma_D^c + \gamma_m \bar{C}}} \frac{\sigma_c^W}{(\sigma_w^W)^2 + (\sigma_c^W)^2 + (\sigma_x^W)^2} \right] \Delta C\, \phi_c^{h-1}. \tag{1.26}$$

Table 1.6: Full set of calibrated parameters

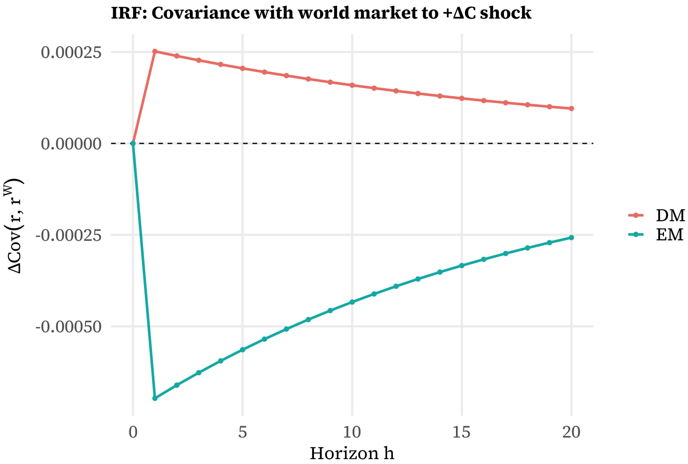
Parameter	Symbol	Estimate
Exposure to global volatility u^W	σ_D^W	0.150
Baseline network variance	σ_D^C	0.050
Global weight in world return	σ_W^W	0.95
Network weight in world return	σ_W^C	0.31
Price of global volatility risk	π_W	0.050
Price of network-composition risk	π_C	0.020

Notes: The parameters (σ_W^W, σ_W^C) are estimated jointly with the other structural parameters and normalized such that $(\sigma_W^W)^2 + (\sigma_W^C)^2 = 1$. They scale the relative contributions of the global volatility and network composition shocks to the world return and are required to reproduce the model-implied betas.

Figure 1.11 displays the implied impulse responses using $\gamma_{DM} = 0.52$, $\gamma_{EM} = -0.85$, $\phi_c = 0.99$, and a small centrality shock $\Delta C = 0.003$. The DM response (red) is positive and persistent, while the EM response (blue) is negative of similar magnitude, reflecting the composition channel.

Equation (1.26) and Figure 1.11 summarize the model's core mechanism. A positive centrality shock increases a DM's covariance with the world portfolio—reflecting deeper integration within the USD-priced core—but lowers an EM's covariance, reflecting risk reallocation within a segmented periphery. The impulse response decays slowly due to the high persistence of centrality ($\hat{\phi}_c \simeq 0.99$), matching the gradual evolution of trade networks in the data. The reduced-form model provides a compact description of how trade structure and financial segmentation interact. Under a single USD pricing kernel, the key object is the bloc-specific composition parameter γ_m, which governs the mapping between network position and priced exposure. A positive γ_m (DMs) links centrality to greater sensitivity to global shocks; a negative γ_m (EMs) links centrality to diversification and lower covariance with the world factor. This parsimonious mechanism reproduces

Figure 1.11: **Impulse response of covariance with the world factor to a centrality shock.**



The figure plots the model-implied impulse response of the covariance between a country's excess return and the world market, $\mathrm{Cov}(r_i, r^W)$, following a one-time positive shock to trade-network centrality ($\Delta C > 0$). The horizontal axis shows the horizon h (years after the shock), and the vertical axis shows the deviation of $\mathrm{Cov}(r_i, r^W)$ from its steady state. The red line with circular markers corresponds to developed markets (DM), and the teal line with triangular markers corresponds to emerging markets (EM). The impulse response is computed from Eq. (1.26) using the estimated parameters $\gamma_{DM} = 0.52$, $\gamma_{EM} = -0.85$, and $\phi_c = 0.99$. A one-time shock of $\Delta C = 0.003$ is applied at $h = 1$, with persistence $\phi_c = 0.99$. Units are deviations from baseline covariance.

all major empirical regularities—opposite β–centrality slopes in USD and LC terms, declining volatility with centrality in EMs, and the asymmetric crisis responses of Exhibit 3— while remaining analytically tractable and directly estimable from observable moments.

1.6 Conclusion

This paper reexamines the relationship between global trade integration and financial exposure through the lens of network structure and pricing segmentation. A conventional prediction of international asset-pricing theory is that greater integration should uniformly raise exposure to globally priced risk. I document that this prediction holds across developed markets but fails systematically in emerging markets. Whereas developed markets exhibit the expected positive relationship between trade centrality and global equity betas, emerging markets show the opposite: greater centrality coincides with lower exposure to the USD-priced world factor, in both USD and local-currency terms. These patterns are not anomalies or measurement artifacts but reflect a fundamental compositional asymmetry in how countries become more central in global trade networks. The empirical results reveal that trade centrality is not a uniform measure of "global integration." For developed markets, rising centrality primarily deepens ties with other advanced economies. For emerging markets, rising centrality is driven by the expansion of intra-EM trade—new connections within a financially segmented periphery. These composition differences explain why greater trade embeddedness amplifies exposure in the core but attenuates it in the periphery. The same increase in centrality can therefore have opposite financial consequences depending on who the partners are and in what currency their shocks are priced. By linking trade-network centrality to priced exposure through a composition channel, I reconcile the asymmetric patterns of global risk transmission across developed and emerging markets, offering a unified explanation for why globalization can simultaneously deepen integration in the core while preserving segmentation in the periphery.

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1.7 Alternative Measures of Integration and Robustness

This appendix demonstrates that the core results linking trade-network centrality to global risk exposure are robust to alternative measures of economic integration. The main analysis in Section 1.4 uses *network centrality* as a structural measure of embeddedness that captures both the intensity and the composition of trade connections. However, one might ask whether the opposite-signed β–centrality slopes simply reflect generic differences in openness or trade diversification, rather than network structure per se. To address this concern, I replicate the analysis using three standard integration metrics: (1) trade openness, measuring the *depth* of external exposure; (2) export concentration (HHI_X), measuring the *breadth* of export destinations; and (3) import concentration (HHI_M), measuring the *diversification of input sources*. In all cases, the results are consistent with the main findings: for developed markets (DMs), greater integration—whether by openness, breadth, or diversification—is associated with higher exposure to the global USD-priced factor; for emerging markets (EMs), the same measures are associated with lower or flat exposure, consistent with the segmentation mechanism described in the paper.

(1) Trade openness: depth of integration

The first metric is the standard measure of trade intensity,

$$\text{Openness}_{i,t} = \frac{Exports_{i,t} + Imports_{i,t}}{GDP_{i,t}},$$

which captures the *depth* of integration without regard to partner composition. High openness implies large trade volumes relative to domestic output (deep integration), while low openness reflects shallow external exposure. If the paper's results were driven simply by trade volume or openness, then replacing centrality with openness should yield similar slopes across DMs and EMs.

Figure 1.12 plots the average USD beta against average openness for each country. The pattern mirrors the main result: among DMs, openness is positively correlated with global betas; among EMs, the slope is weakly negative. This suggests that the core asymmetry—positive in DMs, negative in EMs—is not due to differences in overall trade intensity but rather to how integration is composed. DMs with higher openness trade predominantly within the USD-priced core, amplifying global exposure, whereas EMs with similar openness tend to trade with other EMs, dampening exposure. Thus, even the depth measure, which ignores network structure, produces results consistent with the composition mechanism once regimes are separated.

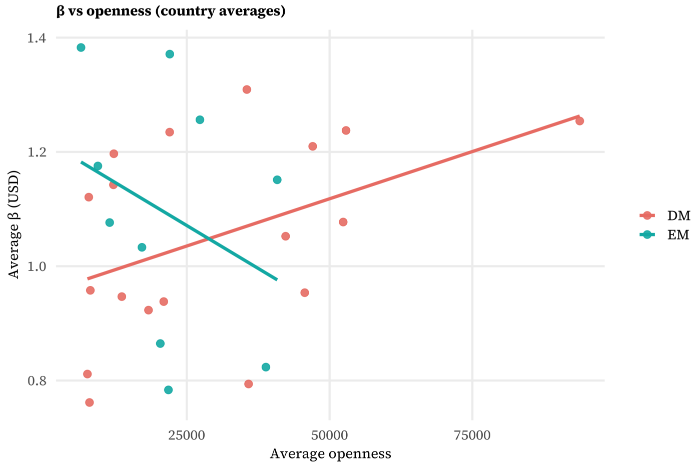

Figure 1.12: **Openness and global betas.** Average USD betas versus trade openness. DMs (red) show a positive slope; EMs (blue) show a flat or slightly negative slope. Integration depth alone cannot account for the cross-sectional divergence.

(2) Export concentration: breadth of destinations

The second metric captures the *breadth* of export portfolios using a Herfindahl–Hirschman index:

$$HHI_X(i,t) = \sum_j (s_{ij,t}^X)^2, \qquad s_{ij,t}^X = \frac{X_{ij,t}}{\sum_k X_{ik,t}},$$

where $X_{ij,t}$ is exports from country i to destination j. High HHI_X indicates that exports are concentrated in a few destinations (low diversification), while low HHI_X implies a broad export base. In an integrated setting, broader export diversification (low HHI_X) should correlate with higher exposure to common global shocks, since trade links are more evenly distributed across the global system. In a segmented world, however, the effect depends on who the additional partners are: if new destinations are mostly other EMs, diversification could lower covariance with the global pricing kernel.

Empirically, the β–HHI_X relation again differs by regime. As shown in Figure 1.13, DMs display a weakly positive slope—countries with more di-

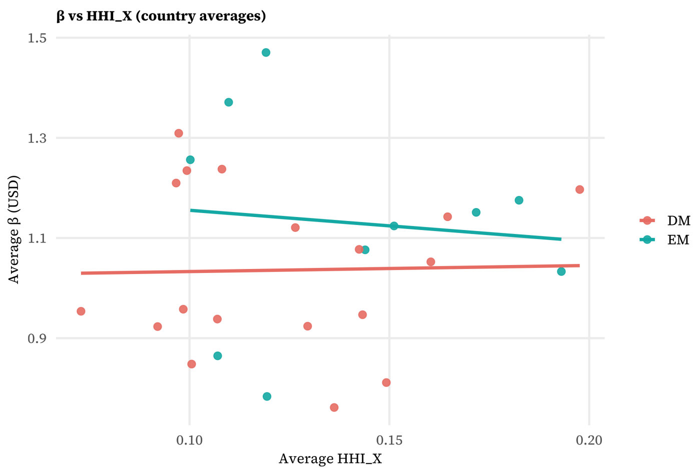

Figure 1.13: **Export concentration and global betas.** Average USD betas versus export concentration (HHI_X). Lower concentration (greater diversification) is associated with higher exposure for DMs but lower exposure for EMs.

versified export bases (lower HHI_X) exhibit higher USD betas—consistent with broad integration within the USD-priced core. EMs, by contrast, exhibit a negative slope: those with more diversified export destinations (lower HHI_X) show smaller global betas. This pattern arises because export diversification in EMs is overwhelmingly intra-EM: the marginal new partner is another EM whose shocks are not fully priced in USD. Hence, diversification increases the number of links but shifts exposure away from globally priced risk, reproducing the negative β–centrality slope through a different lens.

(3) Import concentration: diversification of sources

The third measure focuses on the sourcing side:

$$HHI_M(i,t) = \sum_j (s^M_{ij,t})^2, \qquad s^M_{ij,t} = \frac{M_{ij,t}}{\sum_k M_{ik,t}},$$

where $M_{ij,t}$ denotes imports by country i from source j. High HHI_M implies reliance on a small number of import partners, while low HHI_M indi-

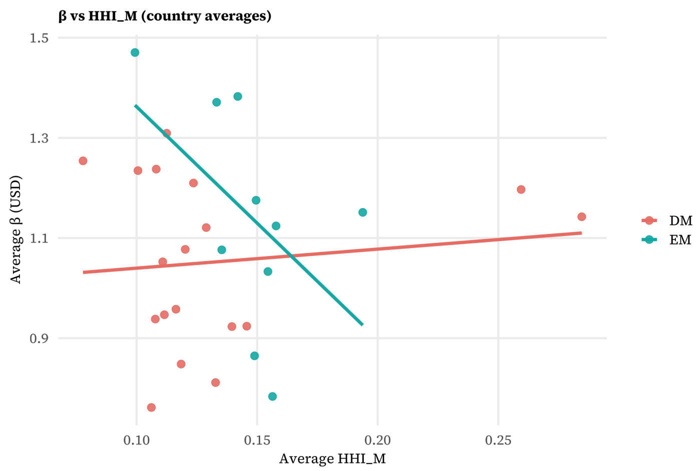

Figure 1.14: **Import concentration and global betas.** Average USD betas versus import concentration (HHI_M). For EMs, lower concentration (greater diversification) is associated with lower global exposure, consistent with intra-EM risk reallocation.

cates diversified sourcing and potential risk-sharing across multiple suppliers. If greater diversification increases exposure to common shocks, DMs should exhibit a positive β–diversification slope, whereas segmented EMs should not.

Figure 1.14 confirms this pattern. For DMs, the relationship between import diversification and global betas is flat to mildly positive. For EMs, the slope is negative and statistically significant: EMs that import from a wider set of partners—mostly other EMs—display lower global betas. This finding aligns closely with the composition channel: EMs that diversify import sources within the EM bloc insulate their cash flows from global USD shocks, even though they increase the number of trade relationships. Diversification within a segmented periphery therefore lowers priced exposure, whereas diversification within the integrated core raises it.

Discussion and interpretation

Across all three measures—trade openness, export diversification, and import diversification—the results are qualitatively identical to those obtained using network centrality. In DMs, greater integration (by any metric) increases exposure to globally priced risk. In EMs, greater integration reduces exposure, reflecting the composition of their marginal trade links. This robustness highlights two key points.

First, the asymmetry is not a statistical artifact of the centrality measure or of network construction. Even when integration is measured purely in terms of trade volume or partner diversification, the sign of the relationship between integration and global betas reverses across blocs. The negative slope in EMs is therefore not about how much they trade but about *with whom*. Second, the persistence of these results across metrics underscores that what matters for priced risk is the *composition of linkages* rather than their intensity. Centrality provides a sufficient statistic because it aggregates both the depth and the partner mix of trade integration; traditional openness and HHI measures capture only fragments of this structure but nonetheless point to the same underlying composition mechanism.

Taken together, these robustness checks reinforce the paper's central message: integration is multidimensional, and its financial implications depend on network topology and pricing regime. Regardless of how integration is measured—by depth, breadth, or partner concentration—the direction of the mapping between integration and global risk exposure flips across regimes. For DMs, deeper and broader integration within the USD-priced core amplifies exposure to global risk; for EMs, expanded integration within a segmented periphery attenuates it. This invariance across integration metrics provides strong empirical validation for the composition-based segmentation mechanism formalized in the model.

1.8 Controlling for Income and Development Level

A natural concern is that the positive centrality–beta relation for developed markets (DMs) and the negative relation for emerging markets (EMs) could reflect differences in income or development rather than trade structure. If higher-income countries are also more central, and if richer economies are more financially integrated, then the estimated coefficients on centrality might simply proxy for cross-country income variation rather than for network composition. To rule out this possibility, I re-estimate the main cross-sectional regression of equity betas on trade-network centrality while controlling for each country's initial income level.

Specifically, I augment equation (**??**) with the first available observation of log per-capita income, denoted $\ln y_i^{\text{first}}$, obtained from the World Development Indicators. The regression is estimated on the average country-level data used in the main analysis:

$$\beta_i = \alpha + \gamma\,\text{Centrality}_i + \delta\,\mathbf{1}\{EM_i\} + \theta\,\big(\text{Centrality}_i \times \mathbf{1}\{EM_i\}\big) + \psi\,\ln y_i^{\text{first}} + \varepsilon_i. \tag{1.27}$$

Table 1.7: **Betas on Centrality controlling for initial income.**

	Dependent variable: β_i
Intercept	3.022^{*}
	(1.756)
Centrality	0.454^{***}
	(0.141)
EM	-0.001
	(0.057)
Centrality × EM	-14.974^{***}
	(4.633)
Initial log income	0.924
	(0.621)
Observations	32
R^2	0.323
Adjusted R^2	0.223

Robust SEs in parentheses. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

The results, reported in Table 1.7, confirm that the main pattern is unchanged once initial income is taken into account. The coefficient on Centrality remains positive and highly significant for DMs, while the interaction term for EMs is large, negative, and highly significant. The estimated effect of initial income is small and statistically insignificant, indicating that early-stage development differences do not drive the relationship between network position and global risk exposure.

Economically, this finding strengthens the interpretation of trade-network centrality as a structural determinant of priced risk, distinct from income or growth potential. Countries' income levels capture the scale of their economies, whereas centrality measures their position within the global trade architecture. Even after conditioning on development, central DMs exhibit higher betas and central EMs lower betas, implying that the composition of trade linkages—not simply wealth or openness—drives the observed asymmetry. In other words, the β–centrality slopes are a pure *trade-network effect*: they reflect how embeddedness in differently priced trading blocs shapes exposure to the global USD-denominated stochastic discount factor rather than cross-sectional differences in growth or income per se.

This robustness is consistent with the model's interpretation of γ_m as a composition parameter that links trade-network structure to priced global exposure. It rules out alternative explanations based on income convergence or development stages, showing that the key mechanism operates through the *geometry* of international trade networks, not through generic macroeconomic development.

1.9 Local-Currency (LC) Betas and Trade Centrality

This appendix reports the corresponding results using local-currency (LC) returns. LC betas remove the exchange-rate valuation component of USD returns and isolate co-movement in local real equity returns.

The LC results mirror the USD results but with slightly stronger contrasts. The β–centrality slope remains positive for DMs and becomes more negative for EMs. Quantitatively, the coefficients differ by less than one standard error for DMs and by roughly 30% for EMs. Thus, the asymmetry is robust to currency denomination: it arises from real cash-flow co-movement rather than exchange-rate effects.

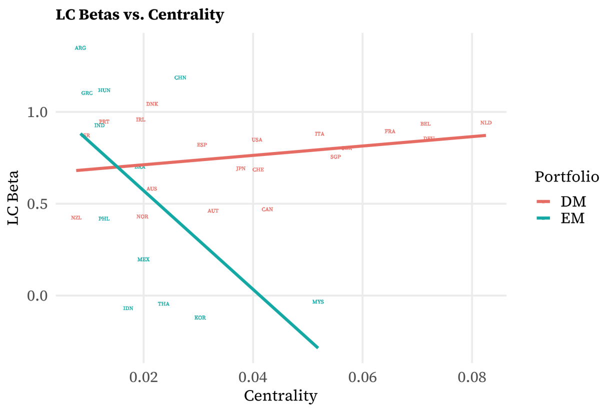

Figure 1.15: **Trade-network centrality and LC-denominated equity betas.** Cross-sectional relation between a country's trade-network centrality and its equity beta with respect to the world portfolio (local-currency returns).

Table 1.8: **Cross-sectional regression: LC betas on trade-network centrality.**

	Dependent variable: β_i^{LC}
Centrality$_i$	2.72
	(3.36)
$\mathbf{1}\{EM_i\}$	0.45*
	(0.25)
Centrality$_i$ $\times$ $\mathbf{1}\{EM_i\}$	-29.61^{***}
	(8.85)
Constant	0.66***
	(0.15)
Observations	32
R^2	0.34

Robust SEs in parentheses. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Because LC betas exclude currency revaluation, the persistence of opposite-signed slopes in both denominations rules out a purely nominal explanation. For DMs, higher centrality increases real exposure to global fundamentals; for EMs, higher centrality lowers it by shifting trade toward locally priced partners. This invariance across currency bases strengthens the interpretation that trade-network centrality is a structural determinant of real and financial integration rather than an artifact of exchange-rate movements.

Figure 1.16 reproduces the baseline analysis using local-currency (LC) returns instead of USD returns. The pattern is virtually identical to the USD benchmark: the slope of equity betas on trade-network centrality is positive and statistically significant for developed markets (DMs) and negative for emerging markets (EMs). The similarity of these relations across currency

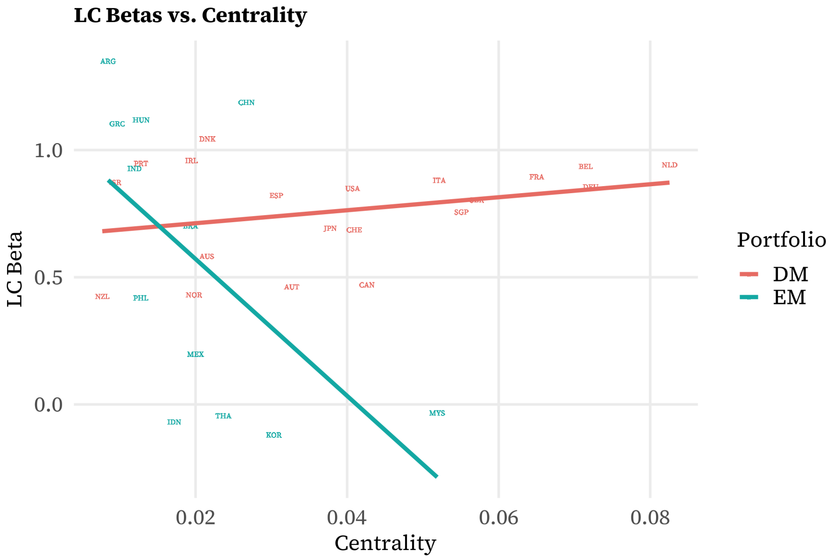

Figure 1.16: **Trade-network centrality and LC-denominated equity betas.** Cross-sectional relation between a country's trade-network centrality and its equity beta with respect to the world portfolio measured in local-currency (LC) returns. Red circles = developed markets (DMs); blue circles = emerging markets (EMs). Fitted lines show within-group OLS slopes.

denominations demonstrates that the centrality–beta asymmetry is not an artifact of exchange-rate valuation effects. For DMs, higher centrality raises co-movement with global fundamentals even after removing currency conversions, while for EMs, higher centrality continues to dampen exposure by shifting trade toward locally priced partners. The robustness of these findings across denominations reinforces the conclusion that trade-network centrality captures a structural, real channel linking trade composition to priced global risk rather than a purely nominal or financial one.

The LC results mirror the USD results: slopes are positive for DMs and negative for EMs. Hence, the opposite-signed centrality–beta relation is not driven by currency revaluation but by real cash-flow co-movement.

1.10 Summary Statistics

Table 1.9: Summary statistics by country

Group	Country	N	Avg. Centrality	$\hat{\beta}^{\mathrm{USD}}$	$\hat{\beta}^{\mathrm{LC}}$
DM	AUS	50	0.021	0.947	0.582
DM	AUT	49	0.033	1.142	0.461
DM	BEL	24	0.071	1.238	0.935
DM	CAN	50	0.043	0.794	0.470
DM	CHE	50	0.041	0.923	0.686
DM	DEU	50	0.072	1.254	0.855
DM	DNK	50	0.022	0.958	1.043
DM	ESP	50	0.031	1.235	0.821
DM	FRA	50	0.065	1.210	0.894
DM	GBR	50	0.057	0.954	0.803
DM	IRL	32	0.019	1.197	0.958
DM	ISR	26	0.009	0.811	0.872
DM	ITA	50	0.052	1.309	0.881
DM	JPN	50	0.038	1.077	0.692
DM	NLD	50	0.083	1.052	0.941
DM	NOR	50	0.020	0.762	0.430
DM	NZL	32	0.008	0.924	0.424
DM	PRT	32	0.013	1.121	0.947
DM	SGP	49	0.055	0.938	0.756
DM	USA	50	0.041	0.848	0.848
EM	ARG	21	0.008	1.383	1.349
EM	BRA	26	0.019	1.371	0.700
EM	CHN	26	0.027	1.124	1.187
EM	GRC	31	0.010	1.470	1.103
EM	HUN	24	0.013	1.076	1.118
EM	IDN	32	0.017	1.033	-0.069
EM	IND	27	0.012	1.256	0.927
EM	KOR	32	0.030	1.151	-0.120
EM	MEX	32	0.020	0.823	0.197
EM	MYS	32	0.052	0.784	-0.034
EM	PHL	30	0.013	1.175	0.418
EM	THA	30	0.024	0.865	-0.045

1.11 Proof: pricing, returns, and betas

Euler equation and log–linear identity. No-arbitrage implies

$$E_t[m_{US,t+1}R_{i,t+1}] = 1.$$

Let $r_{i,t+1} \equiv \log R_{i,t+1}$ and $pd_{i,t} \equiv \log \frac{P_{i,t}}{D_{i,t}}$. The standard Campbell–Shiller linearization gives

$$r_{i,t+1} \approx k_0 + k_1\big(pd_{i,t+1} - pd_{i,t}\big) + \Delta d_{i,t+1}, \qquad 0 < k_1 < 1. \tag{1.28}$$

Throughout, shocks are $u^i, u^g, u^w, u^c, u^x \sim \mathcal{N}(0,1)$ and orthogonal.

Affine pd conjecture. Postulate

$$pd_{i,t} = A + Bz_t^i + Cz_t^w + D\,C_{i,t}. \tag{1.29}$$

With the state laws $z_{t+1}^i = (1-\phi_i)\theta_i + \phi_i z_t^i + \varepsilon_{t+1}^i,\ z_{t+1}^w = (1-\phi_w)\theta_w + \phi_w z_t^w + \varepsilon_{t+1}^w,\ C_{i,t+1} = (1-\phi_c)\theta_c + \phi_c C_{i,t} + \varepsilon_{i,t+1}^c$, we have

$$pd_{i,t+1}-pd_{i,t} = B(\phi_i-1)z_t^i+C(\phi_w-1)z_t^w+D(\phi_c-1)C_{i,t}+\text{terms in } \varepsilon^i, \varepsilon^w, \varepsilon^c.$$

Substitute dividend and SDF processes. Dividends:

$$\begin{aligned}\Delta d_{i,t+1} = \mu_D + \psi z_t^i + \psi_w z_t^w + \sigma_D\sqrt{z_t^i}\,u_{t+1}^i + \sigma_D^g\sqrt{z_t^i}\,u_{t+1}^g + \sigma_D^w\sqrt{z_t^w}\,u_{t+1}^w \\ + \sqrt{\sigma_D^c + \gamma_m C_{i,t}}\ \sqrt{z_t^w}\,u_{t+1}^c + \sigma_D^x\sqrt{z_t^w}\,u_{t+1}^x.\end{aligned} \tag{1.30}$$

USD SDF:

$$-\log m_{US,t+1} = \alpha+\chi z_t^i+\tau z_t^w+\sqrt{\gamma z_t^i}\,u_{t+1}^i+\sqrt{\kappa z_t^i}\,u_{t+1}^g+\sqrt{\delta z_t^w}\,u_{t+1}^w+\sqrt{\omega z_t^w}\,u_{t+1}^c+\sqrt{\xi z_t^w}\,u_{t+1}^x. \tag{1.31}$$

Conditional log Euler approximation. Using a second-order log approximation of the Euler equation $0 = \log E_t[e^{\log m_{t+1}+r_{i,t+1}}] \approx E_t[\log m_{t+1}+$

$r_{i,t+1}] + \frac{1}{2}\,\mathrm{Var}_t(\log m_{t+1} + r_{i,t+1})$, and substituting (1.28), (1.29), (1.30), (1.31), collect coefficients on the states $(z_t^i, z_t^w, C_{i,t})$. matching the coefficients yields

$$B = \frac{-\chi - \frac{1}{2}\Big[(\sqrt{\gamma} - \sigma_D)^2 + (\sqrt{\kappa} - \sigma_D^g)^2\Big]}{k_1(\phi_i - 1)}, \tag{1.32}$$

$$C = \frac{-\tau - \frac{1}{2}\Big[(\sqrt{\delta} - \sigma_D^w)^2 + (\sqrt{\omega} - \sigma_D^c)^2 + (\sqrt{\xi} - \sigma_D^x)^2\Big]}{k_1(\phi_w - 1)}, \tag{1.33}$$

$$D = \frac{\gamma_m(\sqrt{\omega} - \sigma_D^c)}{k_1(\phi_c - 1)}. \tag{1.34}$$

Each square-root difference, e.g. $(\sqrt{\omega} - \sigma_D^c)$, is the gap between the SDF loading and the dividend loading on the same shock; those gaps are the reduced-form representation of the conditional covariance terms in the Euler equation.

Expected excess returns. Insert (1.29) into (1.28) and take conditional expectations:

$$E_t[r_{i,t+1}^e] = \bar{r}_i + \kappa_i z_t^i + \kappa_W z_t^w + \kappa_C C_{i,t},$$

with

$$\kappa_i = k_1 B(\phi_i - 1) + \psi, \quad \kappa_W = k_1 C(\phi_w - 1) + \psi_w, \quad \kappa_C = k_1 D(\phi_c - 1). \tag{1.35}$$

Thus the *centrality* price loading is governed by D via κ_C.

Return innovations. Subtract the conditional mean to obtain

$$\begin{aligned} r_{i,t+1}^e - E_t[r_{i,t+1}^e] &= (\sqrt{\gamma} - \sigma_D)\sqrt{z_t^i}\,u_{t+1}^i + (\sqrt{\kappa} - \sigma_D^g)\sqrt{z_t^i}\,u_{t+1}^g \\ &\quad + (\sqrt{\delta} - \sigma_D^w)\sqrt{z_t^w}\,u_{t+1}^w + \Big(\sqrt{\omega} - \sqrt{\sigma_D^c + \gamma_m C_{i,t}}\Big)\sqrt{z_t^w}\,u_{t+1}^c + (\sqrt{\xi} - \sigma_D^x)\sqrt{z} \end{aligned} \tag{1.36}$$

Each coefficient is the difference between SDF and dividend loadings—i.e., the shock-by-shock covariance piece.

World beta (USD) and its centrality slope. Let world return innovations be $r^W_{t+1} - E_t r^W_{t+1} = \sigma^W_w u^w_{t+1} + \sigma^W_c u^c_{t+1} + \sigma^W_x u^x_{t+1}$. Project (1.36) on this space. Orthogonality eliminates u^i, u^g. The USD beta is

$$\beta^{USD}_i = \frac{\mathrm{Cov}(r_i, r^W)}{\mathrm{Var}(r^W)} = \frac{\sigma^w_D\, \sigma^W_w + \sqrt{\sigma^c_D + \gamma_m C_i}\, \sigma^W_c + \sigma^x_D\, \sigma^W_x}{(\sigma^W_w)^2 + (\sigma^W_c)^2 + (\sigma^W_x)^2}. \tag{1.37}$$

Differentiating w.r.t. centrality gives the comparative static

$$\frac{\partial \beta^{USD}_i}{\partial C_i} = \frac{1}{2} \frac{\gamma_m}{\sqrt{\sigma^c_D + \gamma_m C_i}} \frac{\sigma^W_c}{(\sigma^W_w)^2 + (\sigma^W_c)^2 + (\sigma^W_x)^2}. \tag{1.38}$$

Hence $\mathrm{sign}\left(\partial \beta^{USD}/\partial C\right) = \mathrm{sign}(\gamma_m)$: positive in DMs and negative in EMs. The extra global shock u^x rescales the denominator but does not affect the sign.

Equations (1.32)–(1.34) (valuation), (1.35) (expected returns), (1.36) (innovations), and (1.37)–(1.38) (betas) together show how the Euler equation and the shared shock structure pin down returns. Centrality affects pricing exactly through the covariance channel encoded by the gaps between SDF and dividend loadings; the bloc-specific parameter γ_m determines the sign of the centrality–beta slope.

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Chapter 2

The Risky Capital of Emerging Markets

with Espen Henriksen and Ina Simonovska

Abstract

We build a panel of stock market returns across 37 developed and developing countries spanning five decades. We document: (1) higher and more volatile returns in poorer over richer countries; (2) higher returns in countries with more sensitive dividends to changes in global predictable growth. We quantitatively explore whether consumption-based long-run risk can reconcile these patterns. When we estimate the parameters that govern the U.S. investor's consumption growth and each market's dividend growth process, the model generates higher risk premia in emerging over developed markets, and predicts levels and volatilities of stock market returns that are at par with data.

JEL Classification: F21, F37, F44, G12, G15, O4, O16, E22

Keywords: emerging markets, equity returns, equity-premium puzzle, long-run risk, global financial cycle, Lucas Paradox, returns to capital

2.1 Introduction

Since the late 1980s, following financial account liberalizations, emerging markets have been an appealing investment option for global investors. One common claim presented for investment in these markets is their potential to provide hedging opportunities for U.S. investors due to their higher returns and lower correlation with the U.S. market, in contrast to developed economies, whose business cycles are more closely aligned with that of the United States. Yet, capital flows to emerging markets remain systematically lower than those to developed markets—a phenomenon consistent with the "Lucas Paradox," which highlights the persistent empirical observation that required returns to capital are higher in less developed markets (Lucas, 1990).

In this paper, we propose that U.S. investors require significantly higher risk premia to invest in emerging markets, which results in lower capital intensities and capital inflows into these countries than what would otherwise have been the case. To explore this, we focus on equities, a crucial asset class that is directly comparable across countries and operates in relatively frictionless markets. Utilizing the MSCI database, we analyze USD-denominated annualized stock market returns from 22 developed markets dating back to 1970 and 15 emerging markets from 1988 to 2020.

Our analysis of 37 equity markets, which collectively represent 86% of global stock market capitalization and two-thirds of world GDP over the past five decades, reveals two key findings. First, stock market returns are higher and more volatile in emerging markets compared to developed ones. Doubling a country's income per worker is associated with a 3.1 percentage point reduction in mean stock market returns, while the correlation between return volatility and income is -0.6. Second, markets with higher stock returns exhibit stronger covariance between their equity dividend growth rate and global dividend growth, defined as the stock-market-capitalization-weighted mean of dividend growth among five major economies: U.S., U.K., France, Germany, and Japan. This relationship arises from differences in dividend growth rate volatilities across countries. Emerging markets experience substantially higher volatility in dividend growth rates, but exhibit weaker correlations with global dividend growth.

Motivated by these empirical findings, we investigate whether the risk-return trade-off predicted by asset pricing theory can explain the observed differences in risk premia and volatilities between developed and emerging markets. Specifically, we examine the role of long-run risks as defined by Bansal and Yaron (2004) who focus on the pricing of risks associated with persistent fluctuations in economic growth prospects. Our approach is guided by two key motivations. First, an expanding body of literature, initiated by Aguiar and Gopinath (2007b), highlights the significance of shocks to trend growth rates in explaining business cycle dynamics in poor and emerging markets, as well as in reconciling disparities in macroeconomic behavior between these markets and developed economies. Second, long-run risks have demonstrated profound implications for asset pricing, successfully addressing several longstanding puzzles in the literature, such as the equity premium puzzle (Bansal and Yaron 2004; Colacito and Croce 2011; Nakamura et al. 2017), the low correlation between consumption differentials and exchange rates (i.e., the Backus and Smith (1993) puzzle) and the forward-premium anomaly (Colacito and Croce, 2013b), bond return predictability and violations of uncovered interest parity (Bansal and Shaliastovich, 2013), and currency risk premia (Colacito et al., 2018b). In this context, we analyze whether the heterogeneity in risks stemming from volatile and uncertain growth prospects can account for the differences in international stock market returns, particularly between wealthy and poorer nations.

We consider an international endowment economy where a key distinguishing feature is the focus on a representative U.S. investor, endowed with a consumption stream, dividend payments from risky capital investments across different countries, and access to a risk-free asset. Beyond the inherent risks of stochastic dividend payments, the investor is also exposed to exchange rate risk, as foreign equity payouts are denominated in local currencies. Initially, we simplify the global economy by assuming the existence of a single, perfectly tradable good, which eliminates exchange rate volatility. In subsequent robustness exercises, we assess the impact of incorporating exchange rate risk. This approach allows us to concentrate on the core mechanism of our model: the pricing of global long-run risks in equity markets.

We assume that economic growth rates feature a small but persistent component, which manifests itself in both consumption growth and growth in dividend payments from invested capital. This persistent component comprises a global common piece and an idiosyncratic one unique to each country, with the latter being uncorrelated across markets. Countries vary in their sensitivity to the global piece. With recursive preferences as in Epstein and Zin (1989), a U.S. investor's valuation of foreign assets becomes particularly sensitive to persistent global shocks. Countries whose dividend growth is more responsive to these shocks are perceived as riskier investments by the U.S. investor and must offer higher expected returns to compensate for this added risk. Additionally, each country is subject to both common and idiosyncratic transitory shocks, which affect growth rates for only a single period. The common transitory shocks also contribute to risk premia differentials for a U.S. investor.

Under standard assumptions for preference parameters, quantifying the implications of long-run risks in our model presents three key challenges. First, we must identify global shocks. Second, we need to measure the exposure of individual countries' dividends to both global long-run growth prospects and purely transitory shocks. Third, we must estimate the parameters governing the U.S. investor's consumption growth process. Identifying global persistent shocks is particularly difficult due to the limited availability of historical macroeconomic and financial data, which exist for a few developed economies. To address this, we utilize the Macro-History Database described in Jordà et al. (2019) and Jordà et al. (2017), which includes data on consumption per capita and price-to-dividend ratios. Balancing the trade-off between cross-sectional and time-series coverage and data quality, we define our global variable using data from five major economies–U.S., U.K., France, Germany, and Japan–over the 1940-2020 period.

Following insights from Bansal et al. (2012) and Colacito et al. (2018b), we exploit the model's prediction that a country's logged price-to-dividend ratio is determined solely by the global persistent process. This relationship allows us to project future consumption growth onto lagged values of the price-to-dividend ratio, thereby recovering the time series of the persistent process. We implement this projection within a panel regression framework and define the persistent process as the equally-weighted mean of the predicted component of consumption growth across the five countries. The residual variation in global consumption growth, computed as the equally-weighted mean across the same five economies, is the transitory global component.

U.S. consumption growth exposures to global shocks follow from linear regressions of U.S. per-capita consumption growth series on the estimated global persistent process and the residual transitory component of global consumption growth. Further, we define a global dividend growth process, which incorporates the same global persistent and transitory shocks, in addition to an orthogonal transitory shock reflecting the higher volatility observed in financial over macroeconomic variables. Global dividend growth represents the stock-market-capitalization-weighted mean of dividend growth rates for the same five major economies over the 1975-2020 period, and the returns to this portfolio closely align with the returns on the MSCI-defined "World Index," validating our definition of a global variable.

We regress each country's dividend growth on the residual global dividend growth, excluding the persistent component, to recover exposures to transitory global shocks. With these estimates in hand, we recover exposures to the persistent global process from regression coefficients (i.e. covariances) of countries' dividend growth on total global dividend growth. These parameters directly link to excess returns in the model. Emerging markets whose dividends exhibit strong co-movement with global dividend growth have high inferred exposures to the persistent global component, prompting U.S. investors to demand higher risk premia to invest there.

Using this methodology across the 37 countries in our dataset, we demonstrate that long-run risks explain a substantial share of the observed return disparities. The model predicts a mean cross-country excess return of 9.8%, closely aligning with the observed mean of 9.3%. For the U.S., the model implies an excess return of 6.2%, consistent with the historical figure reported in Bansal and Yaron (2004), though slightly below the observed post-1970 mean of 7.6%. The model also predicts lower and less volatile excess returns in wealthier countries, with a doubling of income per worker leading to a 1.3 percentage point decline in risk premia. Notably, the correlation between model-implied and actual mean excess returns across countries is 0.54. Finally, the average standard deviation of returns predicted by the model is 0.365, closely matching the observed value of 0.336, with a cross-sectional correlation of 0.59 between model-implied and realized standard deviations.

To gain insights behind the mechanisms at play, we set all countries' exposures to the transitory global shock to match the level estimated for the U.S., while maintaining the heterogeneous exposure parameters to the global persistent process as in our baseline specification. The mean model-implied excess returns increase slightly to 10%, and the correlation between model-implied and observed excess returns remains virtually unchanged. These results reinforce the conclusion that cross-country differences in risk premia are largely driven by the sensitivity of dividends to the global persistent process. Yet, accounting for global transitory shocks in dividend growth remains crucial. To illustrate this, we set all countries' exposures to the transitory global shock equal to the level estimated for the U.S. We then infer new exposures to the persistent global process using the same dividend growth covariance moment between a country and the world as in the baseline. This effectively loads all cross-sectional variation in the covariance moment on the single parameter governing exposure to the global persistent process. While mean risk premia decrease slightly to 9.3%, the cross-sectional variation undergoes a significant shift. The correlation between mean model-implied and observed excess returns plummets to 0.2, indicating that the model struggles to capture the variation seen in the data.

The preceding analysis assumes away real exchange rate fluctuations, which can pose a risk to the U.S. investor when they covary with her stochastic discount factor. The literature offers a variety of theoretical perspectives on exchange rates, ranging from frictionless models where bilateral exchange rate movements mirror differences in consumers' stochastic discount factors, to frictional frameworks that introduce wedges and models where exchange rates are entirely decoupled from macroeconomic variables (Itskhoki, 2021). Rather than exploring various mechanisms of exchange rate determination, we perform an empirical robustness exercise to assess the potential impact of exchange rate risk on equity risk premia within our framework. When the real exchange rate deviates from unity, dividend growth rates in our analysis must be expressed in each country's local currency, as the U.S. investor prices these flows–including exchange rate fluctuations–through her no-arbitrage condition. Turning to MSCI, which provides consistent coverage of equity markets data in both USD and local currency, we re-estimate the key exposure parameters for dividend growth rates using local-currency data, while maintaining all other model parameters at their baseline values. The resulting model-implied mean risk premium declines by one percentage point to 8.2%, and the correlation between model-implied and observed risk premia across the 37 markets increases marginally to 0.56. Notably, model-implied risk premia remain closely aligned with those from our benchmark calibration, as evidenced by a cross-sectional correlation of 0.7 between the two sets of estimates.

The exercise demonstrates the robustness of our key moment–the comovement of countries' dividend growth with global dividend growth–in quantifying long-run risks in the presence of exchange rate volatility. This result aligns with the observation that dividend growth volatility is largely invariant across local- and foreign-currency measures for most countries. Similarly, Chernov et al. (2024) report that the volatility of equity returns is nearly identical whether measured in local currency or in USD. These findings reinforce the conclusion that different dividend loading on priced long-run risk is a dominant factor in explaining the observed disparities in risk premia across countries.

The remainder of the paper is organized as follows. In Section 2.2, we describe data sources and measurement of financial and macroeconomic variables across rich and poor countries. In Section 2.3, we document facts on international stock markets. In Section 2.4, we lay out our quantitative analysis of a risk-based explanation of these facts. In Section 2.5, we conclude and discuss directions for future research. Derivations, tables and figures are in the Appendix.

Related literature. Our modeling of international long-run risks is related to Colacito and Croce (2011) and Colacito and Croce (2013b), who examine macroeconomic and financial variables in the U.S. and the U.K., and Lewis and Liu (2015), who study consumption and equity correlations across the U.S., U.K. and Canada. All of these papers find a significant role for shared long-run risk across countries. Our emphasis on heterogeneous exposures to a global shock brings us closest methodologically to Colacito et al. (2018b), who examine a cross-section of FX risk premia in major industrialized countries. We build on these authors' insights and rely on predictive regressions to identify a global persistent process using historical consumption growth and price-to-dividend data. A key innovation in our analysis is to exploit our comprehensive equity dataset to analyze the implications for risk premia and their volatility in both developed and emerging markets for a single U.S.-based investor, and the identification of heterogeneous exposures of dividend growth to the global persistent process from the co-movement of countries' dividend growth with global dividend growth—a moment that has a high predictive power in reconciling observed risk premia across developed and emerging markets. These features of our analysis differentiate our work from the work by Nakamura et al. (2017), who find that heterogeneous exposures of consumption growth to global persistent shocks, alongside country-specific persistent shocks, both priced by local investors, account for a significant amount of the variation in risk premia among 16 developed economies. We find that systematic differences in dividend growth exposures to global shocks, rather than consumption growth exposures, account for risk premia differentials among developed and emerging markets.

Our study of emerging market equities relates our paper to **?**, who examine equity returns in 18 emerging markets during the 1987-2003 period using data from the S&P/IFC Global Equity Market Indices. Using empirical tools that are traditionally employed by the finance literature, the authors document an important role of a global factor–U.S. equity return–in explaining the time series of equity returns in emerging markets. This factor is particularly powerful in accounting for returns in internationally integrated emerging markets, while local liquidity shocks play an important role in driving returns in more closed markets. The authors' findings are one important reason why we focus on emerging markets that are categorized as 'investable' for international investors by MSCI. Brusa et al. (2014b) and Karolyi and Wu (2020) find that global currency factors can empirically reconcile asset returns across countries. Brusa et al. (2014b) further use international equity and currency returns data to quantify an asset pricing model that features a reduced-form stochastic discount factor. These papers complement our work and provide strong support for the role of global risk factors in driving equity risk premia around the world. Unlike these papers, we use consumption and dividend data to quantitatively account for first and second moments in equity returns in the cross-section of developed and emerging markets via the lens of a structural consumption-based long-run risk model. To our knowledge, the sensitivity of countries' dividend growth rates to the global dividend has not been previously studied by the international asset pricing literature.

A broader literature demonstrates the importance of global (often dominated by U.S.) shocks in driving currency returns (Lustig and Verdelhan, 2007; Lustig et al., 2011; Hassan, 2013; Kalemli-Özcan and Varela, 2021; Andrews et al., 2024), sovereign bond returns (Longstaff et al., 2011; Borri and Verdelhan, 2015), equity (Gourio et al., 2013) as well as corporate bond returns and capital flows (Rey, 2015; Miranda-Agrippino and Rey, 2020, 2022), and returns to capital (Hassan et al., 2016). A related strand investigates the failure of capital return equalization (i.e. the Lucas (1990) Paradox) and the implied lack of capital flows from low to high return countries (see Gourinchas and Rey (2013) for a survey of empirical and theoretical studies).

Finally, it is worth to point out that long-run risk is one approach to examine the Lucas Paradox through the lens of asset pricing theory. Two other leading approaches to address asset-pricing puzzles are habits in utility (Campbell and Cochrane, 1999) and rare disasters (Barro, 2006; Gabaix, 2008). Recently, Wang (2021) links currency risk premia to capital accumulation differences across developed countries within the context of a habit persistence model. Farhi and Gabaix (2016) link international asset prices to disaster risk, and Lewis and Liu (2017) show that global and idiosyncratic disasters can reconcile equity return differentials among 20 developed countries. None of these studies examine emerging markets equity returns, which is the focus of our paper. We choose to work with a long-run risk framework, and we add to the literature new evidence that differential dividend growth comovement with the world can reconcile first and second moments in equity returns across rich and poor countries.

2.2 Data Description

In this paper, we pool real and financial data from multiple sources as we describe below.

2.2.1 Measuring Returns Using Financial Data

The macroeconomic literature measures returns to capital in a country via the marginal product of capital (Caselli and Feyrer, 2007). Augmented by changes in the price of capital, this object equals the return to equity, if firms do not incur adjustment costs in capital investment (Gomme et al., 2011). Financial frictions, policy and institutional barriers create a wedge between documented and realized returns for investors (Banerjee and Duflo, 2005; **?**; Restuccia and Rogerson, 2008; Hsieh and Klenow, 2009; Song et al., 2011; Chari and Rhee, 2020), especially in emerging markets. While no measure of returns to capital is ideal, we study stock market returns as equity is an asset class that is more easily comparable across countries.

We obtain daily observations of the Total Return Gross Index by MSCI via Capital IQ, denominated in USD, for 37 developed and emerging markets that account for two-thirds of world GDP. We compute annualized returns, and we subtract annual total CPI inflation for the U.S., obtained from St. Louis FRED. We drop returns below -100% and above 200% to minimize measurement error. Our dataset includes stock market returns in 22 developed markets during the 1970-2020 period, and returns in 15 emerging markets dating back to 1988 until 2020.[1] To compute risk premia, we subtract the mean annual nominal interest rate on 3-month U.S. T-bills for the 1970-2020 period, obtained from St. Louis FRED, from nominal stock market returns for each country.

Table 2.8 in Appendix 2.7 contains descriptive statistics for equity returns in each country.[2] The 37 markets that we study account for 86% of world stock market capitalization and are considered investable by MSCI.[3] While equity is not the only way to access investment opportunities in these markets, it is a very important channel of capital inflow. Among the 37 markets, the stock market capitalization to GDP ratio amounts to a sizeable 63%, and the statistic is not systematically lower in emerging markets (see Figure 2.8 in Appendix 2.6). Continental European countries exhibit some of the lowest stock market capitalization ratios as firms in these markets predominantly rely on bank debt for financing. The majority of emerging markets enjoy higher stock market capitalization ratios, followed by Anglo-Saxon markets such as Canada, USA and Great Britain. Some financial centers such as South Africa (for neighboring African economies), Taiwan, Singapore and Switzerland enjoy stock market capitalization rates of over 150%.

[1]Emerging markets enter the database in the late 1980's and we include them in the analysis upon entry.

[2]We drop China as it was relatively closed to foreign investors for a substantial part of our period of analysis; and furthermore, the theoretical assumption of a small open economy is difficult to justify.

[3]International Finance Corporation (1986) documents that stock markets in developing countries are considered investable categories for international investors beginning in the late 1980's as they underwent significant financial liberalization episodes. MSCI revises the "investability" of different emerging markets for foreigners on a regular basis, and we focus on the markets that they deem investable in our study.

2.2.2 Country-Level Equity Dividends

Valuation theory equates a stock's price with the net present value of all future dividends, so equity dividends play a key role in our quantitative analysis. Specifically, we will derive moments on dividend growth rates to infer the key parameters of interest in the model.

To obtain dividend growth rates, we retrieve two daily series from MSCI, retaining the last date of each year: (i) Price Return Index (in USD), and (ii) Total Return Gross Index (in USD), and we apply a standard procedure in the finance literature (Jagannathan et al., 2000).[4,5] Real dividend growth rates follow by subtracting U.S. inflation rates. We drop dividend growth observations below -100% and above 200% to minimize measurement error. Table 2.9 in Appendix 2.7 contains descriptive statistics for dividend growth rates in each country. For robustness exercises that examine currency risk, we rely on real dividend growth rates denominated in local currency. We obtain the identical series as above from MSCI in local currency. The coverage is identical to the USD-denominated variables. To construct real local-currency variables, we use annual country-level inflation rates from the World Bank's World Development Indicators (WDI thereafter), supplemented by observations from the International Monetary Fund (IMF thereafter) for Taiwan.

[4]Let R_t^p be the annual growth rate of the Price Return Index in year t, and let R_t^{tr} be the growth rate of the Total Return Gross Index in year t. To back out the dividend growth rate, notice that:

$$R_{t+1}^{tr} \equiv \frac{P_{t+1} + D_{t+1}}{P_t} = R_{t+1}^p + \frac{D_{t+1}}{P_t} \Rightarrow \Delta d_{t+1} \equiv \frac{D_{t+1}}{D_t} - 1 = \frac{\left(R_{t+1}^{tr} - R_{t+1}^p\right)}{\left(R_t^{tr} - R_t^p\right)} \frac{P_t}{P_{t-1}} - 1 = \frac{\left(R_{t+1}^{tr} - R_{t+1}^p\right)}{\left(R_t^{tr} - R_t^p\right)} R_t^p - 1$$

[5]These dividend data are implied by and derived from indices that investors actively trade, and are therefore subject to continuous market validation. **?** survey the literature on corporate payout policies and show that share repurchases have become an increasingly important component of payouts in the U.S. As noted by **?**, repurchases remain primarily a U.S.-centric phenomenon. Moreover, repurchase data are typically reported and, unlike dividend data, cannot be isolated or inferred from the prices of indices or other instruments actively traded by investors. As documented by **?** and **?**, dividends are generally funded out of permanent earnings. Measured dividend growth rates likely reflect lasting innovations to the growth rate of firms' underlying profitability, and are relevant for the long-run investor.

2.2.3 Global Macroeconomic and Financial Variables

In order to quantify long-run risks priced by a U.S. investor, we need to identify global persistent and transitory shocks as well as the sensitivity of U.S. consumption growth to these shocks. To estimate the global persistent component of consumption, we will rely on methods used by Bansal et al. (2012) and Colacito et al. (2018b), which require consumption growth and equity price-to-dividend data. Recovering global persistent shocks is particularly challenging because we need to balance time-series and cross-sectional data coverage of these variables. Ideally, we would like to use historical series for as many countries as possible in order to identify a global persistent component in consumption, but price-to-dividend observations are very limited. We obtain historical consumption, population and price-to-dividend observations from the MacroHistory Database, described in Jordà et al. (2019) and Jordà et al. (2017). To balance off time series and cross sectional coverage against measurement error, we define 'global' or 'world' per-capita consumption growth to be the mean of the following five economies: U.S., U.K., France, Germany and Japan during the 1940-2020 period,

$$\Delta c_{Wt} = \frac{1}{5}\sum_{k=1}^{5}\Delta c_{kt}, \tag{2.1}$$

where k indexes each country. These countries account for 61% of stock market capitalization of the 37 countries in our sample—a significant portion of global financial market activity.[6]

[6]GDP-weighted mean of consumption growth across the five countries yields a global process that has a correlation of 0.96 with the equally-weighted mean process. In Appendix 2.10, we show that our measure of a 'global' persistent process is robust to including 6 other developed economies with historical data coverage, which account for an additional 7% of global stock market capitalization. As we describe in Appendix 2.6, stock market capitalization data coverage begins in 1975. This data limitation is the reason why we cannot use stock-market-capitalization weighted mean of consumption data in our definition of a global variable. Since adding countries changes the share of stock market capitalization only marginally, and since we cannot weigh historical series by market cap, we opt to limit our definition of 'global' to the five economies with the largest contribution to global financial activity. The reported share of 61% is a lower bound for our entire study period, which dates back to 1940, when the five major economies constituted a greater share of

Finally, in order to quantify differences in risk premia across countries, we need to identify the sensitivity of each country's dividend growth rates to global persistent and transitory shocks. To do so, we rely on dividend growth rates for each country, as described in Section 2.2.2 above, as well as on a 'global' dividend growth rate. We define global dividend growth to be the stock-market-capitalization weighted mean of dividend growth rates for the same five economies as above during the 1975-2020 period.

$$\Delta d_{Wt} = \sum_{k=1}^{5} \omega_{kt} \Delta d_{kt}, \tag{2.2}$$

where Δd_{Wt} (Δd_{kt}) denotes global or 'world' (country-level) dividend growth rate. Our definition of the 'world' equity market corresponds closely to MSCI's. Figure 2.10 in Appendix 2.10 plots returns to equity for our world portfolio (stock-market-capitalization weighted average of five countries), and the returns from the MSCI series labeled as 'World Index'. The two series are very closely linked, which reflects the dominance of the five countries of our choice in world equity markets. Among these countries, the key role of the U.S. is apparent from the high correlation of U.S. equity returns with both indices. Since we aim to understand cross-country patterns of stock markets in this paper, given historical data limitations discussed above, it is reasonable to approximate 'global' variables with our set of five countries.

Finally, to measure the level of development of each country, we turn to Version 10.0 of the Penn World Tables (PWT thereafter) described in Feenstra et al. (2013). We use this dataset to compute income per worker from annual series of real GDP and employment for each country dating back to the year in which the country enters the MSCI dataset until 2019, which is the terminal year for the PWT dataset. We supplement with data from WDI for the year 2020.

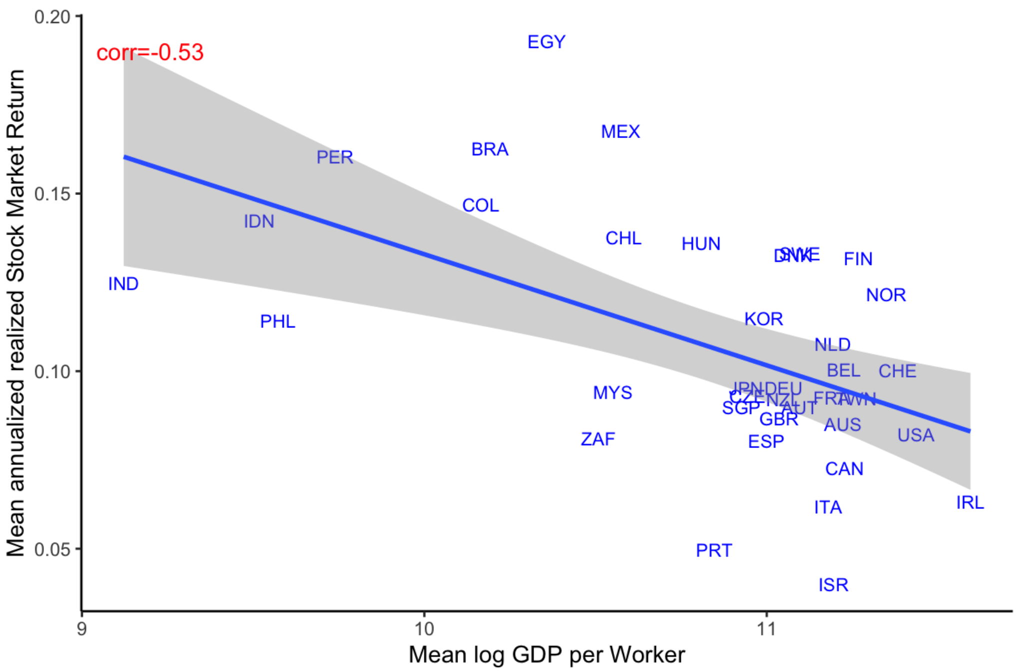
global economic activity.

2.3 Stock Markets Across Countries

In this section, we describe a number of empirical properties of the returns to equities—most notably, a systematic negative link between the level of development and the first and second moments of stock market returns across countries, as well as a positive relationship between returns and the sensitivity of dividends to global fluctuations.

2.3.1 Stock Market Returns

Figure 2.1: The Cross-Section of Stock Market Returns



Notes: The figure plots mean time-series USD-denominated stock market returns against mean time-series log income per worker for 37 countries for 1970-2020. Data Sources: MSCI, PWT and WDI.

Figure 2.1 plots mean realized stock market returns, r_i, against the mean (log) income per worker, y_i, for each country during the entire period for which data are available via MSCI, as well as the correlation between the two variables. Equity returns are systematically higher in poorer countries—

doubling a country's income per worker results in a 3.1 percentage point decline in returns. The top panel of Table 2.1 reports summary statistics for the (mean) realized stock market returns across countries as well the results of a linear regression of returns on income per worker. Stock market returns amount to 10.7% on average, but there is a great deal of heterogeneity across countries. Returns are as low as 4% in Israel and as high as 19.3% in Egypt. The U.S. return to equity is approximately 8% over this period.

Table 2.1: Summary Statistics for Equity Returns and Risk Premia

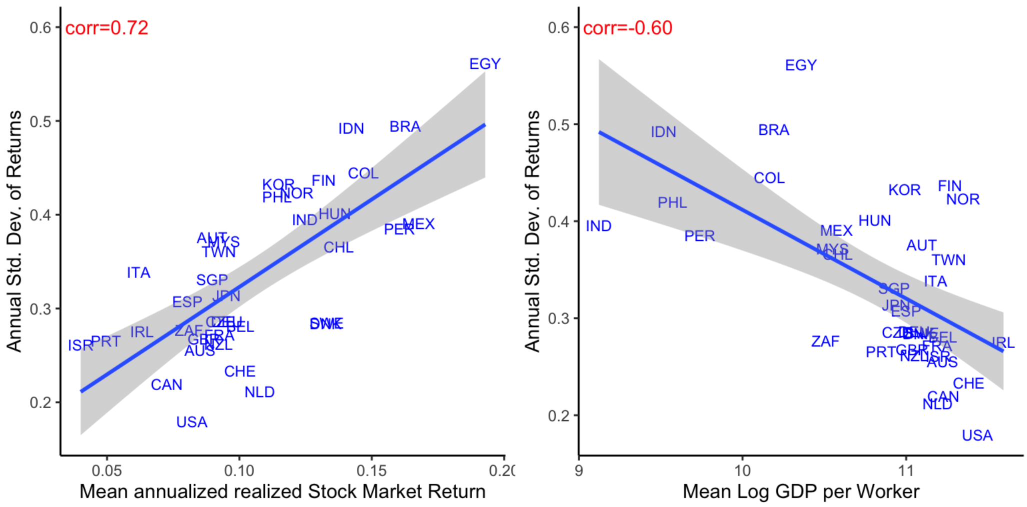
	N	Mean	Median	Std. Dev	Q_{10}	Q_{90}	Constant	y_i	R^2
r_i	37	0.107	0.095	0.034	0.069	0.152	0.446***	-0.031***	0.282
							(0.091)	(0.008)	
r_i^e	37	0.093	0.089	0.032	0.057	0.131	0.349***	-0.024***	0.180
							(0.093)	(0.009)	
σ_{r_i}	37	0.336	0.314	0.088	0.247	0.440			

Notes: Table reports summary statistics for 37 countries, in decimal points, of mean annual USD-denominated real stock market returns r_i, mean excess returns r_i^e computed as the difference between annual USD-denominated nominal stock market returns and the mean nominal interest rate on 3-month U.S. T-bill, and the results of a linear regression of (mean) r_i and r_i^e on (mean) income per worker, y_i. σ_{r_i} denotes the time-series standard deviation of USD-denominated real stock market returns from 1970-2020. Data Sources: MSCI, PWT, WDI and St. Louis Fred. Standard errors statistics in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

In Appendix 2.11, we analyze the time series of the country-level returns. Table 2.17 reports the results of a linear regression of stock market returns for country i in time period t on contemporaneous income per worker, y_{it}. The coefficient estimate on income is -0.041 and highly statistically significant, and it remains negative and precisely estimated when we incorporate year fixed effects to control for global business cycles, and country fixed effects that closely capture very persistent characteristics such as institutional quality levels, which are tied to cross-country capital flow differentials (Alfaro et al., 2008a). In order to eliminate look ahead bias, we repeat the analysis with lagged income per worker in Table 2.18, and we obtain very similar results. Moreover, in Figure 2.9 in Appendix 2.7, we plot the cross-section of mean equity returns against income per worker for three sub-periods of equal duration. The first spans 1971-1987 when only developed markets are in the MSCI sample, while the second two span 1988-2004 and

2005-2021. It is clear from these figures that returns vary significantly over the entire period of study, and that there is no convergence in returns across rich and poor countries. On the contrary, the most recent sub-period shows some of the largest return differentials between developed and emerging markets.

Figure 2.2: Volatility of Stock Market Returns



Notes: Figure plots cross-sectional standard deviation of time-series mean annualized USD-denominated stock market returns in each country against mean annualized realized USD-denominated stock market returns (left) and time-series mean log income (right) for 37 countries for 1970-2020. Data Sources: MSCI, PWT and WDI.

While equities in emerging markets yield higher returns, they are also more volatile. The left panel of Figure 2.2 plots the standard deviation of stock market returns against the mean level of returns for the 37 countries in our sample, as well as the correlation between the two variables, which amounts to 0.72. Not surprisingly, countries that enjoy higher returns also display higher volatilities of returns. Moreover, emerging markets have more volatile returns, as can be seen from the right panel of Figure 2.2, which plots the standard deviation of returns against countries' income levels. The bottom panel of Table 2.1 reports summary statistics for the standard deviation of equity returns across countries. Volatilities differ substantially across countries, with the U.S. being the least and Egypt being the most volatile market.

For completeness, the middle panel of Table 2.1 reports summary statistics for excess returns or risk premia for each country, r_i^e. The mean excess return across countries is 9.3%, and the mean for the U.S. is 7.6% over this period, which aligns with the value of 7.1% in Nakamura et al. (2017), but is higher than the historical value of 6.2% reported by Mehra and Prescott (1985). Risk premia are systematically higher in poorer countries—doubling a country's income per worker results in a 2.4 percentage point decline in risk premia. It is precisely these cross-country differences in risk premia that we will quantify via the lens of an asset pricing model.

2.3.2 Equity Dividend Growth Rates

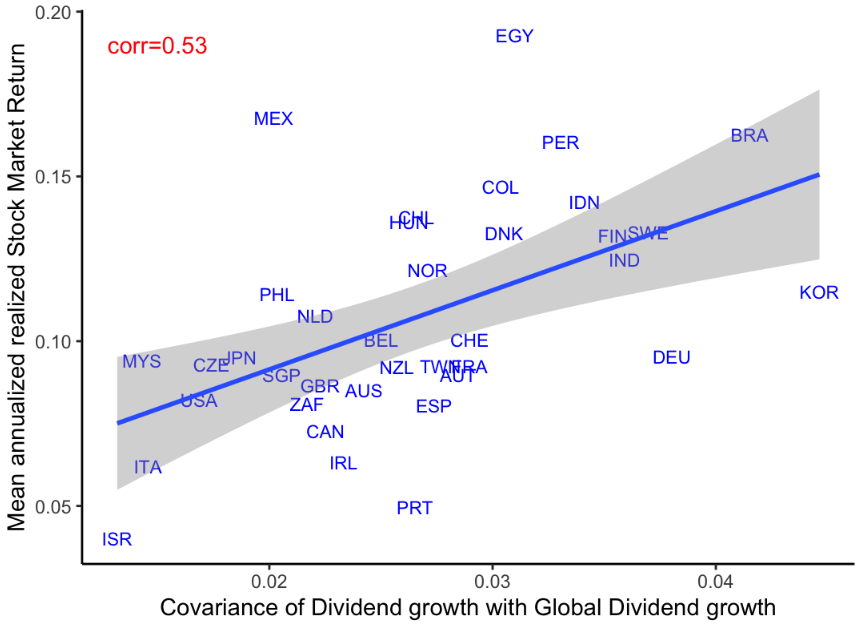
The question that we want to answer is what drives the systematic relationship between returns and income. Figure 2.3 offers the first clue. It plots the mean stock market returns for each country for the entire period of study against $cov(\Delta d_i, \Delta d_W)$, which is the covariance of the country's dividend growth rate with the growth rate of the 'world' dividend defined in expression (2.2) in Section 2.2.3. Countries with higher stock market returns are characterized by higher co-movement of dividend growth rates with the world. In an economy where investors' consumption co-moves with global shocks, countries whose dividends co-move more strongly with the world would be considered more risky, so an investor would demand higher returns to invest there.

In order to analyze the statistical properties of dividend co-movement, recall that one can recover the covariance for each country from a linear regression of the country's dividend growth rate on the global dividend growth rate,

$$\Delta d_{it} = \gamma_i^d \Delta d_{Wt} + \epsilon_{it}. \tag{2.3}$$

The covariance follows from the coefficient estimate of this regression, which we denote by $\gamma_i^d \equiv \frac{cov(\Delta d_i, \Delta d_W)}{var(\Delta d_W)}$. γ_i^d has a natural interpretation:

Figure 2.3: Stock Market Returns and Dividend Comovement



Notes: Figure plots time-series mean annualized USD-denominated stock market returns against the covariance of the country-level annualized dividend growth rate with the world annualized dividend growth rate, defined as stock-market-cap weighted average of dividend growth rates in the U.S., U.K., France, Germany and Japan for 1970-2020. Data Sources: MSCI, PWT and WDI.

it measures the sensitivity of a country's fundamentals (i.e. dividends) to global fluctuations in dividends. In Table 2.2, we report the coefficient estimate for each country, followed by the standard error. The average country has a coefficient estimate of 1.24, and the cross-country standard deviation is 0.34. The coefficients are precisely estimated for almost all the countries, suggesting that this statistic is highly informative.[7] Specifically, the covariances in dividend growth rates with the world extracted from the estimated γ_i^d are strongly related to mean returns to equity as seen in Figure 2.3 above.

When we take a step further and we decompose the covariance of dividend

[7]Our specification assumes that dividend growth is a random walk. To evaluate this assumption, we conduct the Augmented Diskey-Fuller test of the null hypothesis that a unit root is present in each country's dividend growth rate series. We reject the null for 18 countries, and fail to reject it for 19. We re-estimate our coefficients of interest controlling for lags of Δd. The mean estimate is 1.21 with cross-country standard deviation of 0.32, and the correlation between the new estimates and the ones reported in Table 2.2 is 0.91.

Table 2.2: Dividend Comovement and Risk Premia, by Country

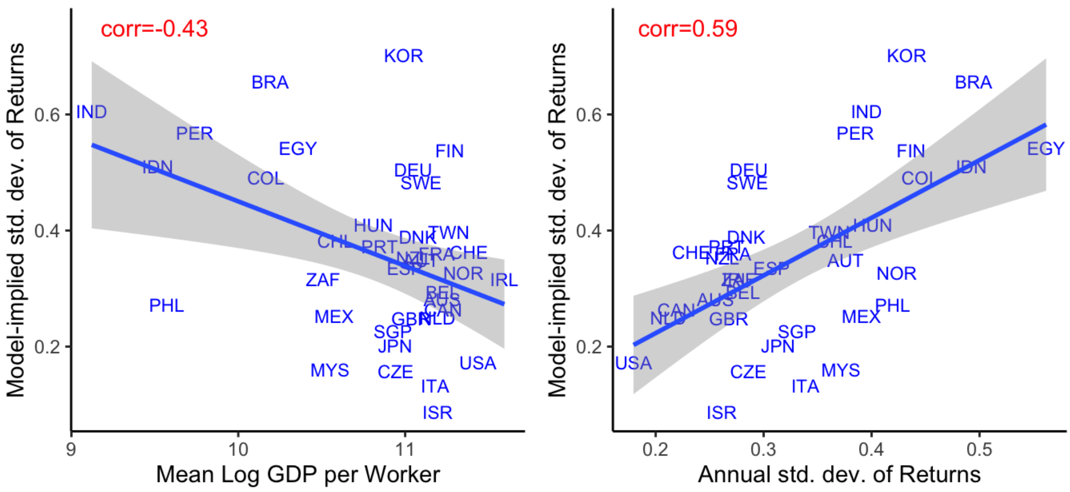
	AUS	AUT	BEL	BRA	CAN	CHE	CHL	COL	CZE	DEU	DNK	EGY	ESP
γ_i^d	1.24***	1.47***	1.28***	1.77***	1.15***	1.48***	1.19***	1.19**	0.97	1.94***	1.56***	1.13*	1.40**
s.e.(γ_i^d)	0.19	0.41	0.18	0.46	0.16	0.20	0.37	0.53	0.82	0.27	0.21	0.61	0.28

	FIN	FRA	GBR	HUN	IDN	IND	IRL	ISR	ITA	JPN	KOR	MEX	MYS
γ_i^d	1.58***	1.48***	1.14***	0.94	1.52**	1.41***	1.04***	0.76	0.74**	0.95***	1.99***	0.90	0.64*
s.e.(γ_i^d)	0.47	0.17	0.16	0.59	0.59	0.33	0.25	0.59	0.30	0.15	0.54	0.59	0.33

	NLD	NOR	NZL	PER	PHL	PRT	SGP	SWE	TWN	USA	ZAF
γ_i^d	1.13***	1.39***	1.15***	1.25*	0.89	1.18***	1.05***	1.89***	1.23***	0.86***	0.85***
s.e.(γ_i^d)	0.15	0.29	0.28	0.72	0.63	0.39	0.19	0.25	0.44	0.06	0.23

Notes: Table reports country-level estimated coefficients γ_i^d and standard errors from a regression of Δd_{it} on Δd_{Wt} from 1970-2020. Data Sources: MSCI.

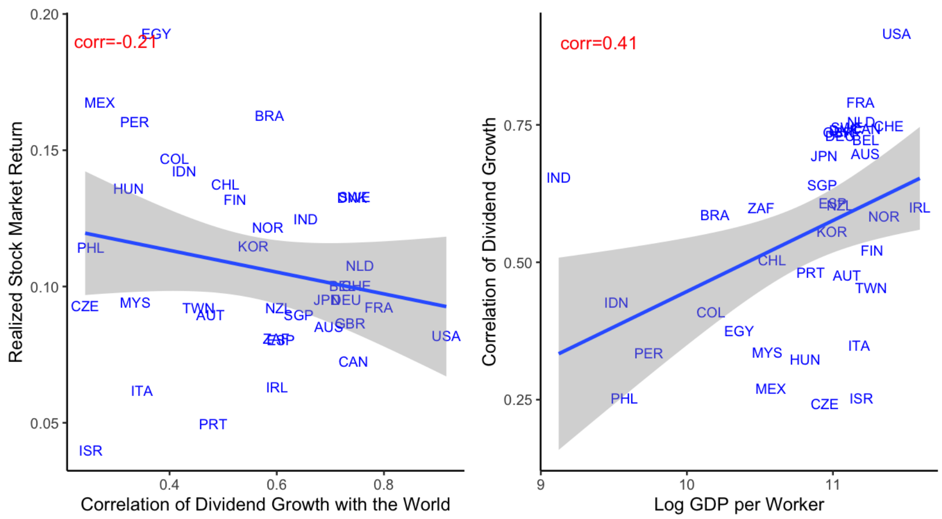
Figure 2.4: Volatility of Dividend Growth



Notes: The figure plots time-series mean annualized USD-denominated stock market return against the annualized time-series standard deviation of dividend growth rates (left) and the time-series standard deviation of annualized dividend growth rates against time-series mean log income per worker (right) for 37 countries for 1970-2020. Data Sources: MSCI, PWT and WDI.

growth rates into each country's standard deviation of dividend growth and the correlation between the dividend growth and the global growth rate, it becomes apparent that the systematic relationship between returns and covariances is driven by countries' volatility levels. In fact, the left panel of Figure 2.4 plots the country-level mean stock market returns against the standard deviation of dividend growth rates as well as the cross-country correlation between the two variables, which amounts to 0.57. Countries that enjoy high returns are those that exhibit high underlying fundamental volatility. Moreover, the right panel of the same figure demonstrates that it is the less developed economies that experience more volatile dividend growth rates.

Figure 2.5: Dividend Growth Correlation



Notes: Left panel plots time-series mean annualized USD-denominated stock market return against the correlation of the time-series country-level annualized dividend growth rates with the world annualized dividend growth rates (left), defined as a stock-market-cap weighted average of the dividend growth rates of U.S., U.K., France, Germany, and Japan. Right panel plots the correlation of the country-level annualized dividend growth rates with the world annualized dividend growth rate against time-series mean log income per worker for 37 countries for 1970-2020. Data Sources: MSCI, PWT and WDI.

In contrast, countries whose dividend growth rates are more correlated with the world do not exhibit systematically different returns as is apparent in the left panel of Figure 2.5, where the correlation is only -0.2. Not surprisingly, it is the poorer countries that are less correlated with the world, which can be seen in the right panel of the same figure.

In Table 2.3, we include summary statistics for the covariance of countries'

Table 2.3: Summary Statistics for Equity Dividends

Variable	N	Mean	Median	Std. Dev	Q_{10}	Q_{90}	corr$(\cdot, r_i)$	corr$(\cdot, y_i)$
cov$(\Delta d_i, \Delta d_W)$	37	0.027	0.027	0.008	0.017	0.036	0.530	-0.266
s.d.(Δd_i)	37	0.353	0.341	0.121	0.213	0.531	0.572	-0.545
corr$(\Delta d_i, \Delta d_W)$	37	0.553	0.586	0.181	0.302	0.747	-0.212	0.414

Notes: The table displays the covariance of the time-series country-level annualized dividend growth rates with the world annualized dividend growth rate, the time-series standard deviation of country-level annualized dividend growth rates, and the correlation of the time-series country-level annualized dividend growth rates with the world annualized dividend growth rate, defined as a stock-market-cap weighted average of the dividend growth rates of the U.S., U.K., France, Germany, and Japan for 1970-2020. Data Source: MSCI.

dividend growth rates with the global dividend growth rate, $cov(\Delta d_i, \Delta d_W)$, the corresponding correlation, $corr(\Delta d_i, \Delta d_W)$, and the standard deviation of each country's dividend growth rate, $s.d.(\Delta d_i)$.[8] For reference, in the last two columns of the table we report how each of these variables correlates with mean returns across countries as well as with mean income levels. Standard deviations range from as low as 0.13 for the U.S. to a three-fold value of 0.61 for an emerging market like Peru, and they are generally decreasing in countries' level of development. Meanwhile, the U.S. enjoys the highest correlation with the world of 0.92, which reflects the predominant role that the U.S. plays in world financial markets, followed by developed European markets. Some of the least correlated countries with the world include Czech Republic, Philippines and Israel, all of which are characterized by a γ_i^d that is not precisely estimated.

The strong relationship between returns and covariances motivates a theory of stock markets in which global shocks take center stage. Nonetheless, given the large variation in stock market returns across countries and over time, it is important to account for both global and idiosyncratic shocks

[8]The cross-country correlation between returns and covariances of dividend growth rates with the world reported in Figure 2.3 is not mechanically driven by the five major economies that constitute the 'world'. Dropping these countries from the sample yields a correlation of 0.56. Furthermore, the covariance moment is robust to the frequency of study. We compute covariances by applying a band-pass filter to isolate fluctuations in the 20-30 year periodicity range, yielding covariances of Δd_i and Δd_W at long-run frequencies. The mean covariance across countries is 0.029, with cross-country standard deviation of 0.0103. The correlation of this low-frequency moment with the contemporaneous covariance reported in Table 2.3 is 0.94.

when modeling the behavior of macro and financial variables across countries. In the following section, we formalize both global and idiosyncratic shock processes and we derive predictions about risk premia via the lens of an asset pricing model.

2.4 A Long-Run Risk Explanation

In this section, we quantitatively explore a novel explanation for the observed cross-sectional variation in returns on the basis of country income levels—namely, the risk-return trade-off implied by asset pricing theory, and specifically, the role of global long-run risks due to uncertainty regarding future economic growth prospects.

2.4.1 The Model

We consider an endowment economy following the international long-run risk literature.[9] We view each market as a small open economy, and we focus on asset valuations from the perspective of a U.S. investor. Consumption of the investor and payments to equity in each country experience shocks to expected future growth rates. Each country is exposed to global and idiosyncratic components of these shocks. Countries differ in their exposure to the global shock process and in the characteristics of the idiosyncratic one. Heterogeneity in exposure to global shocks will play a crucial role in leading to expected return differences across countries.

Preferences. The representative U.S. investor has recursive preferences à la Epstein and Zin (1989). The investor seeks to maximize lifetime utility

$$V_t = \left[(1-\beta) C_t^{\frac{\psi-1}{\psi}} + \beta \nu_t \left(V_{t+1}\right)^{\frac{\psi-1}{\psi}} \right]^{\frac{\psi}{\psi-1}}, \quad \nu_t\left(V_{t+1}\right) = \left(\mathbb{E}_t\left[V_{t+1}^{1-\gamma}\right]\right)^{\frac{1}{1-\gamma}}$$

[9] An important exception is Colacito et al. (2018a), who analyze capital flows in an international production economy featuring long-run risk.

where ψ denotes the intertemporal elasticity of substitution, γ is risk aversion, β is the rate of time discount, and $\nu_t(V_{t+1})$ is the certainty equivalent of period $t+1$ utility. The Euler equations for the risk-free asset, the U.S. risky asset and the foreign risky asset are:

$$\begin{aligned} 1 &= \mathbb{E}_t\left[M_{USt+1}R_{ft+1}\right] \\ 1 &= \mathbb{E}_t\left[M_{USt+1}R_{USt+1}\right] \\ 1 &= \mathbb{E}_t\left[M_{USt+1}\frac{q_{it+1}}{q_{it}}R_{it+1}\right]\ \forall\, i \neq US, \end{aligned} \tag{2.4}$$

where R_{ft} is the return on a risk-free bond, R_{USt} is the (gross) return to equity in the U.S., R_{it} is the (gross) return to equity in country i, denominated in local consumption units, and $q_{it} = \frac{P^c_{it}}{P^c_{USt}}$ is the real exchange rate between the U.S. and country i, where P^c denotes the price of consumption. Furthermore, M_{USt+1} is the U.S. investor's stochastic discount factor (SDF thereafter) whose log, denoted by m_{USt+1}, is given by:

$$m_{USt+1} = \theta \log\beta - \frac{\theta}{\psi}\Delta c_{USt+1} + (\theta - 1)r^c_{USt+1}, \tag{2.5}$$

where $\theta = \frac{1-\gamma}{1-\frac{1}{\psi}}$, and r^c_{USt+1} denotes the return on an asset that pays aggregate U.S. consumption as its dividend, or equivalently, the return to aggregate wealth.

Dynamics of Consumption and Dividends. The following system lays out the joint dynamics of consumption and dividends for any country i:

$$\begin{aligned} \Delta c_{it+1} &= \mu_i + \phi_i x_t + x_{it} + \pi_i \eta_{t+1} + \eta_{it+1} \\ x_{t+1} &= \rho x_t + e_{t+1} \\ x_{it+1} &= \rho_i x_{it} + e_{it+1} \\ \Delta d_{it+1} &= \mu^d_i + \phi^d_i x_t + \tilde{\phi}^d_i x_{it} + \pi^d_i \eta_{t+1} + \tilde{\pi}^d_i \eta_{it+1} + \eta^d_{it+1} \end{aligned} \tag{2.6}$$

In the consumption process, μ_i is the unconditional mean, and x_t and x_{it} are, respectively, the common (i.e. world) and i-specific (i.e. idiosyncratic) time-varying, small but persistent components of the growth rate, so that the conditional mean at t of consumption growth in $t+1$ is $\mu_i + x_t + x_{it}$. The world and local persistent components evolve according to AR(1) processes with persistence parameters ρ and ρ_i, and variances in the innovations σ_e^2 and σ_{ei}^2. ϕ_i governs the exposure of i's consumption to the global persistent component, while η_{t+1} and η_{it+1} are the transitory global and idiosyncratic shocks, respectively, with variances σ_η^2 and $\sigma_{\eta_i}^2$.

Dividend growth has unconditional mean μ_i^d and levered exposures to the persistent components of consumption growth, x_t and x_{it}, captured by ϕ_i^d and $\tilde{\phi}_i^d$. The transitory consumption shocks η_{t+1} and η_{it+1} influence the dividend process with exposures π_i^d and $\tilde{\pi}_i^d$. There is a residual transitory shock that governs dividends denoted by η_{it+1}^d with variance $\sigma_{\eta_i^d}^2$. We assume that all shocks are independent and normally distributed, with respective variances as defined above. We do not assume any particular pattern regarding the exposure parameters to shocks, and we recover the parameter values from macro and financial data in our quantitative exercise.

Risk Premia. To emphasize the importance of countries' heterogeneous exposures to global persistent shocks in driving equity return differences, we begin by solving the model in the absence of exchange rate risk. We assume no exchange rate volatility across countries (driven by the shocks in our model), which is a commonly-employed benchmark by the macroeconomics literature consistent with the assumption of a single traded good, where the real exchange rate is unity. In Section 2.4.4, we relax this assumption and we demonstrate that our main quantitative results are robust in an environment where real exchange rate risk is explicitly priced.

Furthermore, we assume that the return to any asset (including the asset that pays aggregate U.S. consumption as a dividend) that the U.S. investor requires reflects global shocks only. Since we use U.S. consumption growth data in our quantitative exercise, we recognize that idiosyncratic shocks may be present. For these reasons, we allow for idiosyncratic shocks in the empirical processes for consumption and dividend growth, and we separately identify the global shocks in the quantitative exercises. In Section 2.4.3, we quantify the contribution of global and idiosyncratic shocks to countries' dividend growth rates and we discuss the roles that these shocks play in driving equity returns and their volatilities.

To derive risk premia, we solve the model using a log-linear approximation around the balanced growth path as described in detail in Appendix 2.8. Under the assumption of log-normality, the log-linear approximations to the Euler equations in expression (2.4) yield the following risk premia (or excess returns, $\mathbb{E}\left[\hat{r}_i^e\right]$) for a risky asset from country i:

$$\begin{aligned}\mathbb{E}\left[\widehat{r_i^e}\right] &\equiv \log \mathbb{E}\left[R_i\right] - \log \mathbb{E}\left[R_f\right] + \mathbb{E}\left[\Delta q_i\right] \\ &= -\operatorname{cov}\left(m_{US}, r_i\right) - \frac{1}{2}\operatorname{var}\left(r_i\right),\end{aligned} \tag{2.7}$$

where r_i is the logged real return to the risky asset from country i and $\mathbb{E}\left[\Delta q_i\right] = 0$ by assumption. Under the assumption that returns reflect only global shocks and following the methodology outlined in Appendix 2.8, the risk premia can be written as:

$$\begin{aligned}\mathbb{E}\left[\widehat{r_i^e}\right] =& \gamma \pi_{US} \pi_i^d \sigma_\eta^2 + (1-\theta)\kappa^2 \left(\frac{\phi_{US} - \frac{\phi_{US}}{\psi}}{1-\kappa\rho}\right)\left(\frac{\phi_i^d - \frac{\phi_{US}}{\psi}}{1-\kappa\rho}\right)\sigma_e^2 \\ &- \frac{1}{2}\left[\kappa^2\left(\frac{\phi_i^d - \frac{\phi_{US}}{\psi}}{1-\kappa\rho}\right)^2 \sigma_e^2 + \left(\pi_i^d\right)^2 \sigma_\eta^2\right],\end{aligned} \tag{2.8}$$

where κ is a constant defined in Appendix 2.8 that is a function of the mean growth rate of consumption, μ_{US}. The risk premium features a fundamental trade off between the covariance of the SDF and returns (first line of expression (2.8)), and the variance of returns (second line), and it reflects the variance in both temporary and persistent global shocks, σ_η and σ_e, respectively, as well as the exposures of the countries' dividend growth to these shocks, π_i^d and ϕ_i^d. The U.S.-specific consumption exposure parameters, π_{US} and ϕ_{US}, reflect the assumption that the U.S. agent is pricing the assets. With preference parameter values commonly employed in the literature, risk premia are rising in countries' exposures to growth shocks, π_i^d and ϕ_i^d, and differences in these parameters drive cross-country return differentials.

2.4.2 Identification of Parameters

To derive the model's risk premia implications and to assess its ability to account for the cross-section of stock market returns in the data, we must assign values to the parameters governing the preferences as well as the consumption and dividend processes laid out in expression (2.6). From expression (2.8), note that risk premia only reflect parameters that govern the global shocks and countries' exposures to these shocks. Additionally, observe that country-specific exposures of consumption growth to global shocks (ϕ_i and π_i) for foreign countries (vis-à-vis the U.S.) do not drive risk premia since the U.S. investor is pricing the assets in our model.

We use standard preference parameters from the literature (Colacito and Croce, 2011). Furthermore, we implement techniques used by Bansal and Yaron (2004), Bansal et al. (2012), and Colacito et al. (2018b) to estimate the parameters that govern the global persistent process and global transitory shocks using consumption growth and price-to-dividend ratio data for five major economies (U.S., U.K., France, Germany and Japan) during the 1940-2020 period provided by Jordà et al. (2019) and Jordà et al. (2017). U.S. consumption growth exposures to global shocks follow from linear regressions of U.S. consumption growth series on the estimated global persistent process and the residual transitory component of global consumption growth. Since these procedures are standard in the literature, we relegate the details to Appendix 2.9, and we focus on the new moments on dividends that we introduce to the literature, which ultimately drive cross-country return differentials.

Global dividend growth parameters. To identify parameters pertaining to dividend growth rates, we specify a global dividend growth process (measured by expression (2.2) in the data):

$$\Delta d_{Wt+1} = \mu_W^d + \phi_W^d x_t + \pi_W^d \eta_{t+1} + \eta_{Wt+1}^d, \tag{2.9}$$

where $\eta_W^d \sim N(0, \sigma^2_{\eta_W^d})$ is independent of all country-specific and global shocks defined above, and idiosyncratic components have been averaged out. As was the case for each individual country, the global dividend growth process features an additional transitory shock, η^d_W, which reflects the possibility of sources of variation in equity dividends that are not related to sources of variation in real variables such as consumption.

As we describe in Section 2.2.3, we define the 'world' portfolio to be the stock-market-capitalization weighted mean of five major economies during the 1975-2020 period. Due to the relatively short time coverage on dividend growth data in MSCI, compared to consumption growth data, we are not able to identify π^d_W and, especially, ϕ^d_W from linear regressions as in the case for consumption.[10] In order to relate our results to the existing literature, we set the leverage parameter to the global persistent process, ϕ^d_W, to 3, which is the value that Nakamura et al. (2017) use for 12 developed economies. We set π^d_W to 4.9, which implies that the transitory and persistent global shocks that the U.S. agent prices in our model account for 89% of the observed variation in global dividend growth in MSCI data, with the residual shock, η^d_W, accounting for only 11% of variation. Furthermore, as we show in column (i) of Table 2.5 below, our choice of a value of 4.9 implies that the leverage parameter of U.S. dividend growth on the global persistent process is 3.5, which compares favorably to the value for this parameter of 3 used by Bansal and Yaron (2004) and Colacito and Croce (2011) for the U.S.[11]

We summarize the values for the parameters that govern preferences, global processes and the U.S. consumption growth process in Table **??**, alongside the moments. As is clear from this table, the variance of the global persistent shock is lower than the variance of the global transitory shock, and the residual variance of the shock that governs world dividends is rather high,

[10] Missing historical stock market capitalization data, used to weigh countries, prevents us from defining global dividend growth prior to 1975 and estimating π^d_W and ϕ^d_W following the standard procedure used to estimate the U.S. consumption growth parameters outlined in Appendix 2.9. While consumption series across countries are relatively smooth and cross-country means are robust to different weighting schemes (see footnote 8), the same is not true for financial series, so country weights affect the results, and the unweighted mean is quite different from the 'world' equity portfolio as defined by MSCI.

[11] In Figure 2.10 in Appendix 2.10, we show that the 'world' portfolio returns are driven predominantly by U.S. returns, so it is reasonable that the U.S. and the world dividend growth processes in the model have similar leverage parameters to the persistent process.

which reflects the fact that dividend growth is several orders of magnitude more volatile than consumption growth.

Table 2.4: Moments and Parameters for Preferences, Global Processes, and U.S. Consumption

Parameter	Value	Moment
γ	4	Literature (Colacito and Croce, 2011)
ψ	1.5	Literature (Colacito and Croce, 2011)
β	0.99	Literature (Colacito and Croce, 2011)
ρ	0.758	AR1 reg. coeff. est. in eq. (2.6) for x_t estimated in eq. (10), 5 countries
σ_e	0.017	Autoreg. coeff. est. in eq. (2.18) for Δc_W defined in eq. (2.1), 5 countries
σ_η	0.021	Variance of Δc_W (net of x_t) defined in eq. (2.1), 5 countries
μ_{US}	0.021	Mean of Δc_{US}, USA
ϕ_{US}	2.133	Reg. coeff. est. in eq. (2.19) for Δc_{US}, USA
π_{US}	0.425	Reg. coeff. est. in eq. (2.20) for Δc_{US}, USA
ϕ^d_W	3.000	Literature (Bansal and Yaron, 2004; Nakamura et al., 2017)
π^d_W	4.900	$\phi^d_W \sigma^2_x + \pi^d_W \sigma^2_\eta = 89\%$ of var(Δd_W) defined in eq. (2.2), 5 countries
$\sigma_{\eta^d_W}$	0.047	11% of var(Δd_W) defined in eq. (2.2), 5 countries

Notes: Table reports parameter values and moments used in benchmark calibration.

Idiosyncratic dividend growth parameters. We rely on the same MSCI data to assign values to the idiosyncratic parameters that govern the process in the fourth line of expression (2.6) for all countries. To recover π^d_i, we run the following regression for each country:

$$\Delta d_{it+1} = \pi^d_i (\Delta d_{Wt+1} - \phi^d_W x_t)/\pi^d_W + \epsilon_{it+1}, \qquad (2.10)$$

which identifies the parameter of interest under the independence assumption among all idiosyncratic and global shocks. The resulting regression coefficient estimates for each country are reported in Table 2.5. For the majority of countries, the parameters are precisely estimated; the mean centers at 5.97 and the parameter value for the U.S. is 4.21.[12] These parameters do not display any systematic pattern across rich and poor countries.

Finally, given all other parameters, to recover the key dividend exposure parameters to the persistent global process, ϕ^d_i, we rely on the covariance

[12] Standard errors omitted due to space constraints and available in Replication Package.

of a country's dividend growth rate with the world, which is given by $\text{cov}(\Delta d_{it+1}, \Delta d_{Wt+1}) = \phi_i^d \phi_W^d \sigma_x^2 + \pi_i^d \pi_W^d \sigma_\eta^2$. We recover the covariance from the coefficient estimates of country-level regressions of dividend growth rates on the global dividend growth, γ_i^d, as reported in Table 2.2 in Section 2.3.2. Recall that this moment is very informative about stock market returns in the data (see Figure 2.3 in Section 2.3.2), and it is the most important moment in the identification procedure as it directly dictates the risk premia differentials that we document below. Crucially, notice that we identify all country-specific parameters using dividend growth data only—we do not rely on cross-country equity prices in our identification procedure.

We report the resulting parameter values for ϕ_i^d for each country in Table 2.5, along with the correlation of this variable with income per worker. The mean value for the leverage parameter across countries amounts to 6.21 and the value for the U.S. is 3.5.[13] As is evident from Table 2.5, emerging markets display higher exposures to the persistent global process than developed ones–the correlation between income per worker and ϕ_i^d is -0.45–and are characterized by higher model-implied excess returns, as we demonstrate in the next section.

Given the importance of these parameters in the quantitative analysis, in Appendix 2.10, we re-estimate them using four different definitions of the 'global' portfolio that range from the G-10 countries to the entire set of 37 countries in our study. The correlation of the newly-estimated parameters and our benchmark estimates is nearly 1 in all the specifications, even though the mean levels of the parameters change somewhat depending on the specification. These findings imply that approximating the world with our five countries of choice is a reasonable assumption.

[13] It is clear from the regressions used to identify π_i^d and ϕ_i^d that π_W^d scales all country-specific dividend growth leverage parameters. This is why we use the estimate of π_{US}^d of 3.5 to cross check our choice of 4.9 for π_W^d.

Table 2.5: Dividend Growth Exposure Parameters and Resulting Risk Premia, by Country

	AUS	AUT	BEL	BRA	CAN	CHE	CHL	COL	CZE	DEU	DNK	EGY	ESP
ϕ_i^d	5.06	5.99	5.24	10.59	4.72	6.06	6.45	8.22	3.49	7.99	6.44	8.85	5.80
π_i^d	6.02	7.02	6.20	8.51	5.58	7.19	5.76	5.78	4.46	9.40	7.51	5.47	6.72
$\widehat{r_i^e}$	9.30	10.47	9.55	10.61	8.75	10.52	11.27	12.34	6.18	11.31	10.87	12.39	10.30
r_i^e	7.90	8.40	9.40	14.20	6.60	9.40	11.70	12.40	6.90	8.90	12.60	16.90	7.40

	FIN	FRA	GBR	HUN	IDN	IND	IRL	ISR	ITA	JPN	KOR	MEX	MYS
ϕ_i^d	8.71	6.10	4.61	7.61	8.29	9.70	5.69	2.43	3.07	3.89	10.95	4.94	3.50
π_i^d	7.54	7.14	5.57	4.52	7.39	6.86	5.03	3.58	3.59	4.64	9.56	4.33	3.07
$\widehat{r_i^e}$	11.95	10.57	8.54	12.35	11.98	11.79	10.46	3.32	5.18	7.14	9.85	9.35	6.34
r_i^e	11.20	8.60	8.00	11.20	12.20	10.10	4.30	1.70	5.60	8.90	9.40	14.70	7.40

	NLD	NOR	NZL	PER	PHL	PRT	SGP	SWE	TWN	USA	ZAF	corr($\cdot$, y_i)
ϕ_i^d	4.61	5.65	6.23	9.22	5.09	6.44	4.27	7.81	6.76	3.49	5.89	-0.45
π_i^d	5.47	6.74	5.58	6.04	4.27	5.74	5.13	9.10	5.95	4.21	4.11	0.01
$\widehat{r_i^e}$	8.56	10.08	11.07	12.19	9.61	11.27	7.91	11.35	11.54	6.23	10.86	-0.34
r_i^e	10.20	11.60	7.20	13.70	9.40	2.90	8.40	12.70	7.20	7.60	5.80	-0.42

Notes: Table reports the country-level parameters ϕ_i^d, π_i^d, the model-predicted risk premia $\widehat{r_i^e}$ and the risk premia from the data r_i^e. Data Sources: MSCI.

2.4.3 Results

Equity Risk Premia. We begin by evaluating the ability of the model to reconcile observed risk premia in the data, under the assumption that real exchange rates equal unity–i.e., in the absence of real exchange rate risk. We compute risk premia for each country by plugging the parameter values reported in Tables **??** and 2.5 into expression (2.8).

In Table 2.6, we report the summary statistics of risk premia that we compute from our model. The mean risk premium in the model is 9.8%, which is nearly identical to the mean reported in MSCI data of 9.3% in Table 2.1 above. Much like in the data, risk premia are higher in emerging markets—doubling a country's income per worker results in a 1.3 percentage point decline in risk premia. Turning to the cross-section of countries, in Table 2.5, we report excess returns implied by the model, $\widehat{r_i^e}$, and their data counterparts, r_i^e. The model-predicted excess return ranges from 3.32% in Israel to 12.39% in Egypt, and the U.S. value amounts to 6.23%. More interest-

Table 2.6: Summary Statistics for Model-Predicted Variables

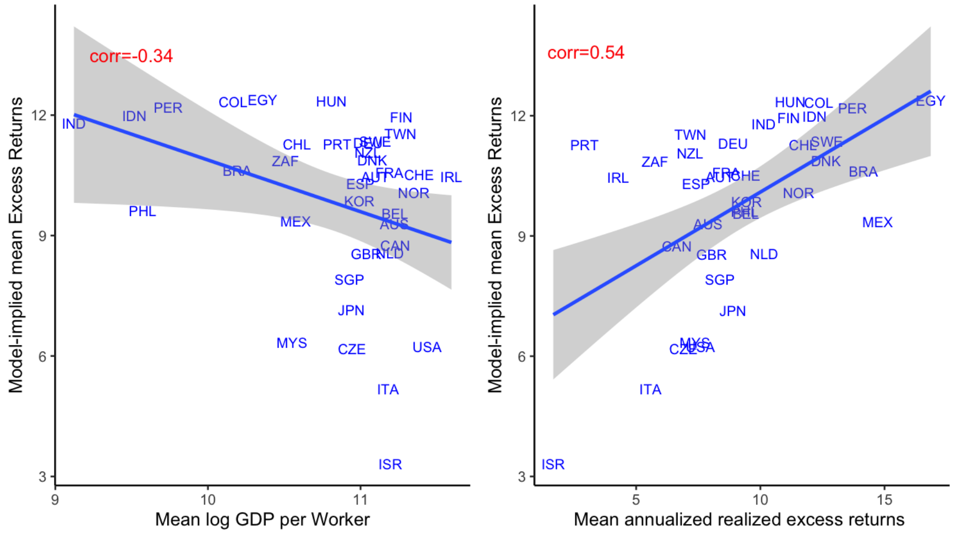
Returns								
	N	Mean	Median	Std. Dev	Q_{10}	Q_{90}	Constant	y_i
$\widehat{r_i^e}$	37	0.098	0.105	0.022	0.063	0.121	0.238***	-0.013**
							(0.065)	(0.006)
$\sigma_{\widehat{r_i}}$	37	0.365	0.357	0.150	0.174	0.552		
Dividends								
	$\frac{\mathrm{Var}(\Delta d_{\text{Long-Run}})}{\mathrm{Var}(\Delta d_{\text{Data}})}$		$\frac{\mathrm{Var}(\Delta d_{\text{Short-Run}})}{\mathrm{Var}(\Delta d_{\text{Data}})}$					
	0.217		0.126					

Notes: The upper panel reports summary statistics of the predicted risk premia from the parameterized model ($\widehat{r_i^e}$) for 37 countries, and the results of a linear regression of $\widehat{r_i^e}$ on the time-series mean log income per worker, y_i. $\sigma_{\widehat{r_i}}$ is the model-implied standard deviation of returns. Regression standard errors are in parentheses. The lower panel reports the cross-country mean fraction of dividend growth variance accounted for by the long-run and short-run components of the calibrated model. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

ingly, the model is able to reconcile the cross section of risk premia in the data. The right panel of Figure 2.6 plots model-implied against realized excess returns at the country level and the accompanying correlation between the two series, which amounts to 0.54. The left panel of Figure 2.6 plots mean model-implied excess returns against mean logged income per worker as well as the correlation between the two variables, which amounts to -0.34.

In the model, risk premia differentials across countries are driven by two parameters, ϕ_i^d and π_i^d, which capture the sensitivity of each country's dividend growth rate to persistent and transitory shocks. To evaluate the role of each parameter in delivering the results, we perform two robustness exercises. First, we set all country-specific π_i^d's to that of the U.S., and we keep the values of ϕ_i^d as in Table 2.5. The first row of Table 2.19 in Appendix 2.11 shows the summary statistics from this exercise. Mean risk premia, denoted by $\hat{r}_r^e$ (for robustness), increase to 10%, and the correlation between the model-implied and realized excess returns is effectively unchanged. Similarly, the correlation between the benchmark model-implied excess returns and the counterfactual one is effectively 1, which suggests that risk premia in the model did not change in the cross-section, only slightly in levels. This finding demonstrates that the values of ϕ_i^d generate the majority of risk

Figure 2.6: Model-Implied Excess Returns



Notes: The above figure plots the predicted risk premia from the parameterized model $\widehat{r^e}$ against income per worker (left) and the predicted risk premia from the parameterized model $\widehat{r^e}$ against the realized risk premia from the data r^e (right) for 37 countries for 1970-2020. Data Sources: MSCI, PWT, WDI and St. Louis Fred.

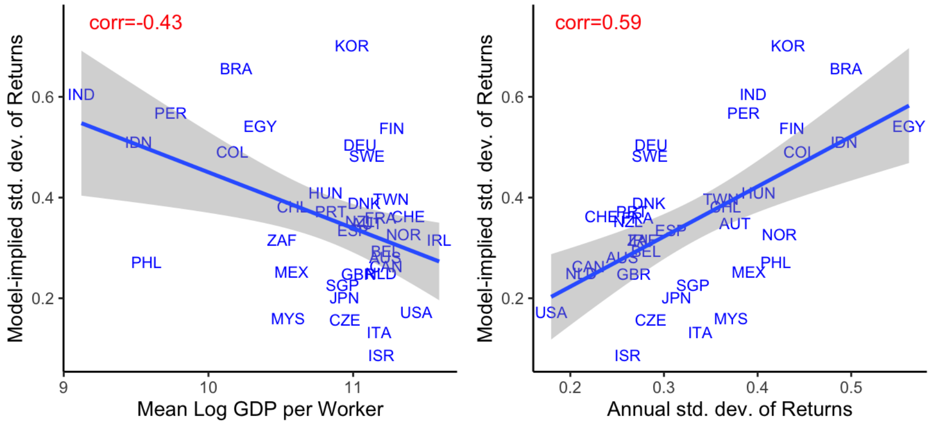
premia differentials across countries.

In the second exercise, we set all country-specific π_i^d's to that of the U.S., and we re-estimate the resulting ϕ_i^d's so as to match the covariance of each country's dividend growth rate with the world. Thus, we are effectively assigning all the cross-sectional variation of the key moment of interest—the covariance of a country's dividend growth rate with the global dividend—on the parameter ϕ_i^d. The second row of Table 2.19 in Appendix 2.11 shows the summary statistics from this exercise. While risk premia levels, denoted by $\hat{r}_r^e$, decrease slightly to 9.3% on average, there is a notable change in the cross-sectional variation. The correlation between the model-implied and realized excess returns drops to a mere 0.2, while the correlation between the baseline and the counterfactual risk premia is only 0.58. This finding implies that it is important to separately identify persistent from transitory shocks when estimating the global dividend process, even though the latter do not play an important role in governing the levels of risk premia.

Volatility of Returns to Equity. We proceed to evaluate whether the model can account for the cross-sectional volatility of equity returns reported in Section 2.3.1. By construction, the model matches the variance of each country's dividend growth rate, which is driven by three objects: long-run global component, along with the country's leverage parameter, $(\phi_i^d x_t)$, short-run global component $(\pi_i^d \eta_{t+1})$, and residual component $(\tilde{\phi}_i^d x_{it} + \tilde{\pi}_i^d \eta_{it+1} + \eta_{it+1}^d)$. The bottom panel of Table 2.6 reports the average standard deviation of dividend growth among the 37 markets in our dataset, as well as the percent of variance that is explained by the long-run and short-run global components. The long-run global component accounts for nearly double the variation than does the short-run global component—22% versus 13%. Not surprisingly, the largest part of the variation is explained by the residual idiosyncratic component.

The second row of the top panel of Table 2.6 reports summary statistics of the volatility of equity returns implied by the model. The model generates a standard deviation of returns to equity of 0.365 for the average country, which is nearly identical to (and somewhat exceeds) the mean of 0.336 reported in MSCI data in Table 2.1 in Section 2.3.1.

Figure 2.7: Model-Implied Standard Deviation of Equity Returns



Notes: The above figure plots the predicted standard deviation of risk premia against the time-series standard deviation of risk premia from the data (left) and time-series mean log income (right) for 37 countries from 1970-2020. Data Sources: MSCI, PWT and WDI.

To evaluate the cross-sectional predictions of the model, in the left panel of Figure 2.7, we plot the model-implied standard deviation of equity returns against the logged income per worker for the 37 countries in our sample, as well as the correlation between the two variables. Consistent with the data, the model generates higher volatilities in emerging markets. Since return volatility contributes negatively to model-predicted excess returns in expression (2.7), it follows that the negative relationship between model-predicted returns and per-capita income is driven by the covariances of returns with the U.S. agent's SDF.

To evaluate the fit of the model to the volatility data, in the right panel of Figure 2.7, we plot the model-implied standard deviation of returns against the standard deviation observed for the 37 countries in MSCI data. The correlation between the two variables is remarkably high—0.59. With these statistics at hand, we conclude that our parsimonious model, which excludes currency risk and relies on a single moment on dividend growth comovement with the world, can reconcile at least 50% of the variation in levels and volatilities of stock market returns across rich and poor countries.

A few observations are in order. Recall that, when we derive risk premia in the model, we assume that the U.S. agent does not price idiosyncratic shocks for country i ($e_{it}, \eta_{it}, \eta_{it}^d$). If we were to relax this assumption, our model-implied measure of risk premia, which excludes real exchange rate risk, would only change to the extent that the term $\text{var}(r_i)$ would change in expression (2.7). This follows directly from the fact that the U.S. SDF does not reflect any i-specific idiosyncratic shocks.[14] This does not mean that the U.S. agent does not price events that affect the dividend that she receives from country i; indeed, shocks to those dividends are at the very heart of the risk premium that the agent demands to hold that asset. It is simply that the agent prices the portion of the variations in dividends that covary with her SDF, and that portion is heterogeneous across assets and captured by parameters ϕ_i^d and π_i^d.

The assumption of pricing global shocks simplifies the quantitative analysis

[14] It is worth to note that the risk premium for U.S. equities would change by an extra term because the U.S. agent would be pricing U.S. idiosyncratic shocks, e_{it} and η_{it}.

significantly, as we do not need to estimate parameters that relate to idiosyncratic persistent and transitory shocks for each country ($\rho_i, \sigma_{ei}, \sigma_{\eta i}, \sigma_{\eta_i^d}$). Notice that, even if we were to estimate these parameters, the ultimate result would be an increase in the model-implied variance of returns and a decrease in the mean returns in each country. Under the current calibration, the model predicts a variability in returns that is at par with the data–and in fact slightly higher. Hence, a model that features idiosyncratic shocks would further raise this variance and worsen the model's fit to the data on average. Thus, given the moments that we choose in our estimation, the model suggests that idiosyncratic shocks are not critical to account for observed risk premia and the variability of returns in the average country. However, if we consider the cross-section in the right panel of Figure 2.7, a number of countries lie below the 45-degree line, which implies that the volatility of returns in those economies is below what the model predicts. It is possible that idiosyncratic shocks may account for the residual volatility in those markets, but the outcome would be a fall in mean model-implied risk premia for the same markets. Thus, it would be more fruitful to consider other sources of risk (or frictions) to improve the fit of the model.

2.4.4 Real Exchange Rate Risk

In the model, we assume that the U.S. agent is pricing all assets—domestic and foreign. Therefore, she faces real exchange rate risk from dividend income incurred from abroad. If the volatility of real exchange rate change, Δq_i, is non-zero, the U.S. agent would be pricing it, which would result in the following risk-premium expression for equity from country i:

$$\mathbb{E}\left[\widehat{r_i^{e,rer}}\right] = -\operatorname{cov}\left(m_{US}, r_i\right) - \frac{1}{2}\operatorname{var}\left(r_i\right) - \operatorname{cov}\left(m_{US}, \Delta q_i\right) - \frac{1}{2}\operatorname{var}\left(\Delta q_i\right) - \operatorname{cov}\left(r_i, \Delta q_i\right). \tag{2.11}$$

Relative to the risk premium equation (2.7) that we quantify, currency risk adds the last three terms in equation (2.11). The last term, which is referred to as the "cross term" is roughly 0 in the data (see Chernov et al. (2024)). Furthermore, $\frac{1}{2}\text{var}(\Delta q_i)$ is a small number (less than 0.1% for a typical country—see Table 2 in Colacito and Croce (2011) for U.S.-U.K. for example). Hence, if currency risk premia were to be incorporated in the model, model-implied risk premia would change by the amount corresponding to the third term, cov$(m_{US}, \Delta q_i)$.

There are a number of theories of the real exchange rate in the existing literature (see Itskhoki, 2021, for summary). In frictionless environments, the change in the exchange rate is

$$\Delta q_{it+1} = m_{it+1} - m_{USt+1}, \tag{2.12}$$

which obtains directly from the Euler equations for domestic and foreign agents who price a given asset using their respective SDF. If agents' SDFs reflect consumption growth, as in our model, then the third covariance term in expression (2.11) would be non-zero (see ex. Colacito and Croce (2011), Colacito and Croce (2013b), and Colacito et al. (2018b) within the context of long-run risk models of the real exchange rate in developed markets, Verdelhan (2010) for a model that builds on consumption habits or the seminal work by Lustig and Verdelhan (2007) that emphasizes the role of U.S. consumption growth).[15] Any market friction would decouple variability in real and financial variables (see ex. Itskhoki and Mukhin (2021), Itskhoki (2021) and Lustig and Verdelhan (2019) among others for discussion on the "exchange rate disconnect puzzle"). Furthermore, exchange rate regimes, which are notably different between emerging and developed markets (as documented by Ilzetzki et al. (2019)), could introduce further sources of exchange rate risk premia for a U.S. agent. Finally, a segment of the finance literature aims to explain the behavior of currency risk premia using reduced-form SDFs that are entirely orthogonal to macroeconomic variables (ex. Verdelhan (2018) and related work).

Given the vast literature, it is outside of the scope of this paper to incorporate the various mechanisms of exchange rate determination. Instead, we conduct a robustness exercise to quantify how large the contribution of ex-

[15]Recently, Hassan et al. (2024) re-examine long-run risk and habit models and their implications for the cross-section of currency risk premia.

change rate risk could be within the context of our model. Observe that, if the real exchange rate is not unity, we would need to denominate dividend growth rates in our quantitative exercise in each country's local currency. This follows directly from the third line in the Euler equation for foreign equity in expression (2.4), where R_{it+1} is denominated in local currency by definition and it reflects local-currency denominated dividends, which becomes more apparent from the approximation methods detailed in Appendix 2.8. Therefore, to evaluate the role of currency risk, we re-calibrate the model's parameters related to dividend growth using dividend series denominated in local currency, as described in Section 2.2.2. Recall that, MSCI data coverage is identical in local currency and USD, which implies that our robustness exercise focuses on the exact same time period for each country as our benchmark calibration above.

Table 2.7: Parameters and Resulting Risk Premia, in local currency

	(i)	(ii)	(iii)
	$\phi_i^{d,lc}$	$\pi_i^{d,lc}$	$\widehat{r^{e,lc}}$
mean	4.88	5.23	8.20
std. dev	1.89	1.57	2.62
corr(.,ϕ_i^d)	0.80	-	-
corr(.,π_i^d)	-	0.72	-
corr(.,r^e)	-	-	0.56
corr(.,$\widehat{r^e}$)	-	-	0.70
corr(.,y_i)	-0.35	0.01	-0.32

Notes: Table reports estimated dividend growth parameters using local currency data and corresponding model-predicted risk premia $\widehat{r^{e,lc}}$, as well as correlations with income, parameters and risk premia from the benchmark calibration, and risk premia in the data. Data Sources: MSCI.

Specifically, we keep the values of all global and U.S. parameters as outlined in Table **??**, and we re-estimate all the country-specific parameters related to dividend growth, ϕ_i^d and π_i^d, using dividend growth series denominated in local currency. We report summary statistics for the estimated country-level parameters and the model-implied risk premia in Table 2.7. Notice that the mean value for the key parameter ϕ_i^d across countries drops to 4.88, compared to the mean of 6.21 for the values reported in Table 2.5.

The implication for model-implied risk premia is a fall in the mean to 8.2% from the benchmark prediction of 9.8%. Hence, risk premia fall on average by roughly 1 percentage point. More importantly, risk premia do not change considerably in the cross section of countries. In particular, the correlation between model-implied and actual risk premia is nearly unchanged–0.56– which is due to the high correlation between model-predicted risk premia under the two specifications. Indeed, notice that the correlation between the resulting parameter values for ϕ_i^d between the two specifications is remarkably high—0.8.

The findings from the robustness exercise suggest that exchange rate risk premia do not have a first-order effect on equity risk premia in this model, when the key model parameters are recovered from countries' dividend growth comovement with the world.[16] In other words, this moment is robust to the currency of denomination. This finding is not surprising in light of the fact that this moment is predominantly driven by the volatility of dividend growth rates, which is very similar whether denominated in local or foreign currency for a typical country. Similarly, Chernov et al. (2024) document that the volaitlity of equity returns is near identical when denominated in local currency and in USD for most countries.

2.5 Conclusion

In this paper, we compile a comprehensive panel of international stock market returns and we document: (1) higher and more volatile stock market returns in poorer over richer countries, and (2) higher stock market returns in countries with higher co-movement of dividends with the world. We find that long-run risk, i.e., risk due to persistent fluctuations in economic growth rates, is a promising channel to reconcile these facts. Key to our results is that emerging markets not only feature large fluctuations in growth

[16] Colacito et al. (2018b) show that currency risk premia account for a significant portion of equity risk premia for 10 developed economies during the 1987-2013 period. We reach a different conclusion using a sample of 37 developed and emerging markets during the 1970-2020 period.

rates, but also that the shocks are systemically related across countries, i.e., these markets are highly exposed to global growth-rate shocks.

In our quantitative analysis, one parameter is critical in generating risk premia differentials across countries—the exposure of a country's dividend growth rate to the world persistent process. This parameter is directly governed by one moment in the data—the co-movement of countries' dividend growth rates with the world. Our parsimonious model accounts for over a half of the observed cross-sectional variation in equity returns and volatilities, but a large amount of variation remains unexplained. Similarly, the behavior of real exchange rates across countries remains unaccounted for, even though real exchange rate risk premia do not appear to be critical in reconciling the cross-section of equity returns.

We leave for future work a more detailed investigation into the sources of the differences in long-run risk that we measure. The implications of such an analysis would be important on many dimensions; from the point of view of our analysis, in reducing required risk premia associated with investments in poor countries and so potentially attracting additional investment flows. Potential avenues of research include understanding how institutional differences across countries shape firms' borrowing constraints (**?**) and dividend payments, which may provide insights into the mechanisms that result in high exposure of emerging markets to global shocks.

We have focused on consumption-based risk due to uncertainty regarding dividend payoffs, both in the short and long run. By doing so, we have abstracted from a number of other sources of risk that may play a role in leading to return differences such as default risk or expropriation risk. Additionally, our model does not shed light on the fundamental source of long-run risk, i.e., changing prospects for technological progress, etc. Further work investigating these issues and their interaction with rates of return on capital around the world could be fruitful.

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Appendix

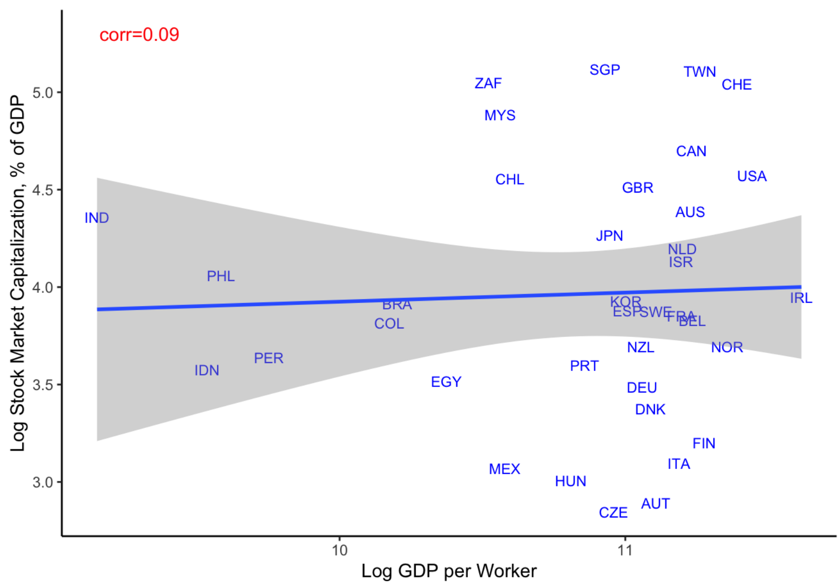
2.6 Supplemental Data Sources

2.6.1 Stock Market Capitalization

We obtain stock market capitalization to GDP ratios from World Development Indicators (WDI) during the 1975-2020 period. Stock market capitalization for the U.K. during 2012-2020, France during 2019-2020, and

Taiwan during 1983-2020 are from CEIC. We compute the ratio relative to GDP using nominal GDP series in USD from IMF for Taiwan. For France and the U.K., we use nominal GDP per capita in USD and population from St. Louis Fred to arrive at total nominal GDP. Figure 2.8 plots the log of the stock market capitalization to GDP ratio against a country's level of development.

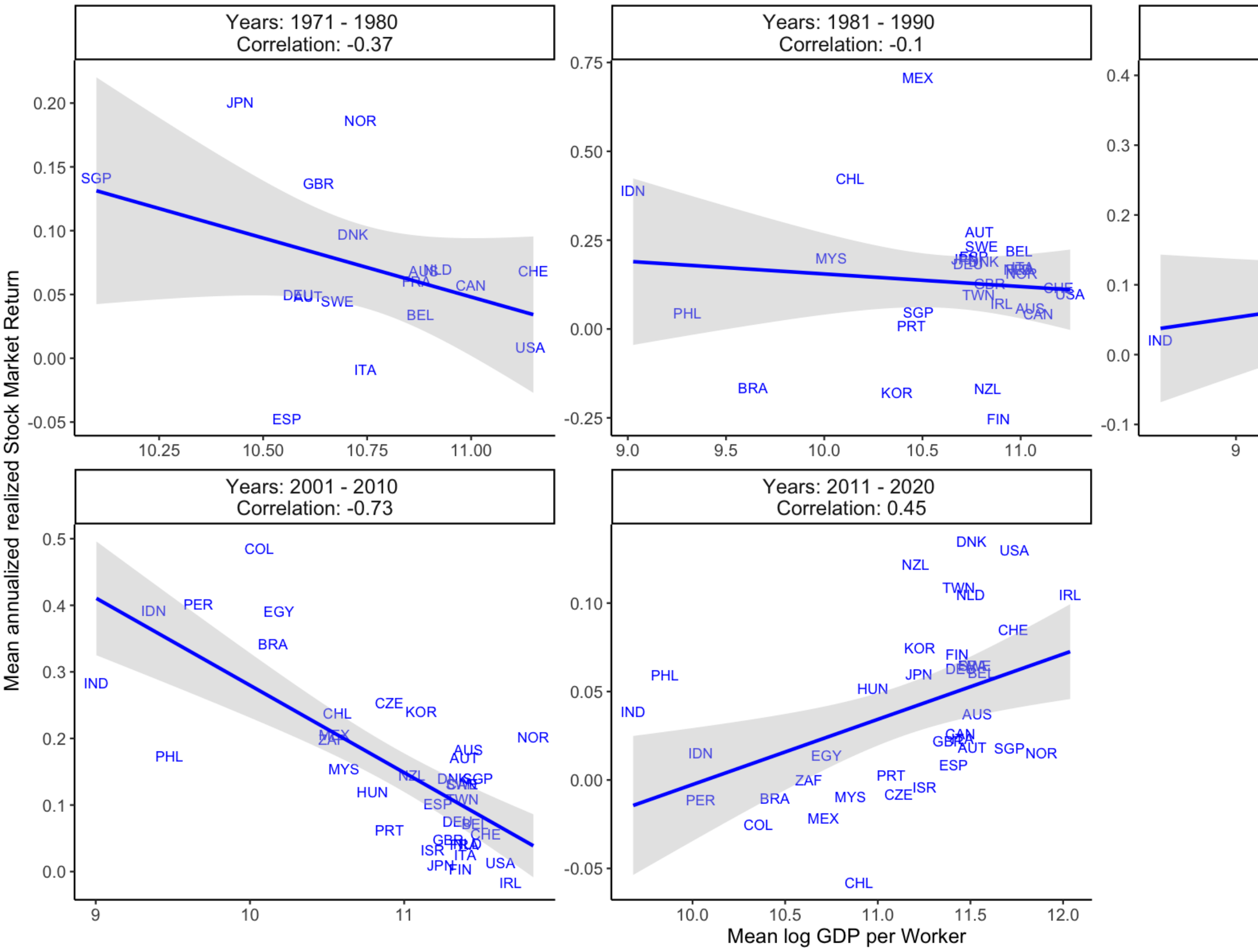
Figure 2.8: Log Stock Market Capitalization (% of GDP)



Notes: Figure plots (log) stock market capitalization (% of GDP) against time-series mean log income per worker for 37 countries from 1975-2020. Data Sources: WDI, CEIC, PWT and WDI.

2.7 Data Description: Supporting Tables and Figures

Figure 2.9: Stock Market Returns and Income Per Worker Over Time



Notes: Figures plot mean annualized USD-denominated stock market returns against time-series mean log income per worker for 37 countries for three sub-periods of equal duration. Earlier decades exclude emerging markets due to data unavailability. Data Sources: MSCI, PWT and WDI.

Table 2.8: Summary Statistics for Equity Returns, by Country

	(i)	(ii)	(iii)	(iv)	(v)	
Country	N	Mean	Median	Std. dev	Q_{10}	Q_{90}
AUS	50	0.085	0.086	0.256	-0.186	0.411
AUT	49	0.090	0.019	0.376	-0.268	0.472
BEL	50	0.100	0.076	0.281	-0.202	0.431
BRA	30	0.163	0.138	0.495	-0.412	0.750
CAN	50	0.073	0.090	0.219	-0.192	0.300
CHE	50	0.100	0.114	0.233	-0.159	0.325
CHL	32	0.138	0.109	0.366	-0.227	0.576
COL	27	0.147	0.121	0.445	-0.341	0.727
CZE	24	0.093	-0.014	0.286	-0.145	0.427
DEU	50	0.095	0.111	0.286	-0.241	0.334
DNK	50	0.133	0.116	0.284	-0.175	0.436
EGY	23	0.193	0.108	0.562	-0.465	0.889
ESP	50	0.080	0.026	0.308	-0.229	0.456
FIN	32	0.132	0.095	0.437	-0.339	0.496
FRA	50	0.092	0.086	0.272	-0.227	0.363
GBR	50	0.087	0.087	0.268	-0.170	0.327
HUN	24	0.136	0.145	0.401	-0.298	0.659
IDN	32	0.142	0.071	0.492	-0.472	0.770
IND	27	0.125	0.073	0.395	-0.286	0.726
IRL	32	0.063	0.115	0.275	-0.241	0.402
ISR	26	0.040	0.057	0.261	-0.327	0.308
ITA	50	0.062	0.020	0.339	-0.252	0.369
JPN	50	0.095	0.081	0.314	-0.264	0.404
KOR	32	0.115	0.090	0.432	-0.402	0.537
MEX	32	0.168	0.149	0.391	-0.238	0.572
MYS	32	0.094	0.018	0.372	-0.225	0.513
NLD	50	0.108	0.109	0.212	-0.159	0.338
NOR	50	0.122	0.051	0.423	-0.299	0.533
NZL	32	0.092	0.094	0.262	-0.223	0.368
PER	26	0.160	0.169	0.385	-0.315	0.659
PHL	31	0.114	0.092	0.419	-0.484	0.602
PRT	32	0.050	0.031	0.266	-0.250	0.416
SGP	49	0.090	0.041	0.331	-0.311	0.503
SWE	50	0.133	0.133	0.285	-0.254	0.478
TWN	32	0.092	0.084	0.361	-0.316	0.496
USA	50	0.082	0.118	0.179	-0.150	0.292
ZAF	27	0.081	0.153	0.277	-0.259	0.428

Notes: Table reports summary statistics for annual equity returns (in decimal form). Annual returns are computed by the authors using daily observations of each country's Total Return Gross Index from MSCI. The sample period begins in the year the country enters the dataset and ends in 2020.

Table 2.9: Summary Statistics for Dividend Growth, by Country

	(i)	(ii)	(iii)	(iv)	(v)	
Country	N	Mean	Median	Std. dev	Q_{10}	Q_{90}
AUS	50	0.041	0.025	0.240	-0.248	0.266
AUT	49	0.043	0.047	0.411	-0.365	0.438
BEL	50	0.022	-0.014	0.239	-0.177	0.225
BRA	30	0.146	0.084	0.462	-0.443	0.704
CAN	50	0.035	-0.018	0.209	-0.177	0.254
CHE	50	0.084	0.046	0.266	-0.155	0.252
CHL	32	0.065	-0.005	0.352	-0.267	0.481
COL	27	0.155	0.077	0.465	-0.328	0.860
CZE	24	0.090	-0.020	0.534	-0.377	0.815
DEU	50	0.075	-0.006	0.360	-0.245	0.416
DNK	50	0.075	0.044	0.295	-0.302	0.423
EGY	23	0.036	-0.049	0.499	-0.464	0.769
ESP	50	0.018	-0.009	0.311	-0.271	0.270
FIN	32	0.123	0.005	0.453	-0.358	0.907
FRA	50	0.041	0.021	0.252	-0.246	0.337
GBR	50	0.027	0.027	0.215	-0.211	0.312
HUN	24	0.057	0.070	0.486	-0.472	0.489
IDN	32	0.100	0.094	0.533	-0.547	0.791
IND	27	0.104	0.101	0.344	-0.305	0.420
IRL	32	0.020	-0.003	0.259	-0.228	0.274
ISR	26	-0.002	0.091	0.394	-0.516	0.504
ITA	50	0.021	-0.007	0.301	-0.382	0.401
JPN	50	0.044	0.033	0.203	-0.165	0.288
KOR	32	0.145	0.041	0.536	-0.371	0.771
MEX	32	0.192	0.272	0.500	-0.364	0.851
MYS	32	0.051	-0.004	0.284	-0.258	0.464
NLD	50	0.037	0.010	0.202	-0.191	0.300
NOR	50	0.072	0.077	0.328	-0.325	0.426
NZL	32	0.007	-0.022	0.285	-0.278	0.382
PER	26	0.241	0.152	0.607	-0.421	1.011
PHL	31	0.082	0.014	0.530	-0.423	0.721
PRT	32	0.067	0.035	0.368	-0.385	0.306
SGP	49	0.056	0.047	0.243	-0.224	0.366
SWE	50	0.085	0.014	0.341	-0.253	0.435
TWN	32	0.118	0.079	0.407	-0.406	0.721
USA	50	0.027	0.026	0.134	-0.141	0.188
ZAF	27	0.035	-0.007	0.227	-0.261	0.300

Notes: Table reports summary statistics of dividend growth. Data Sources: MSCI.

Table 2.10: Summary Statistics for Consumption Growth, by Country

	(i)	(ii)	(iii)	(iv)	(v)	
Country	N	Mean	Median	Std. dev	Q_{10}	Q_{90}
DEU	81	0.022	0.020	0.046	-0.009	0.068
FRA	81	0.021	0.021	0.072	-0.000	0.053
GBR	81	0.017	0.020	0.028	-0.009	0.045
JPN	81	0.031	0.027	0.081	-0.007	0.091
USA	81	0.021	0.019	0.022	-0.002	0.046

Notes: Table reports summary statistics of annual consumption growth. Data are for 1940-2020 period from MacroHistory Database provided by Jordà et al. (2019) and Jordà et al. (2017).

Table 2.11: Summary Statistics for Price-Dividend Ratios, by Country

	(i)	(ii)	(iii)	(iv)	(v)	
Country	N	Mean	Median	Std. dev	Q_{10}	Q_{90}
DEU	79	3.71	3.61	0.65	3.22	4.16
FRA	81	3.56	3.46	0.53	3.04	4.20
GBR	81	3.16	3.15	0.28	2.85	3.48
JPN	81	3.90	3.95	0.78	2.80	4.86
USA	81	3.45	3.42	0.46	2.88	4.04

Notes: Table reports summary statistics of pd ratios. Data are for 1940-2020 period from MacroHistory Database provided by Jordà et al. (2019) and Jordà et al. (2017).

2.8 Model Solution

Processes in our environment:

$$\begin{aligned}
\Delta c_{it+1} &= \mu_i + \phi_i x_t + x_{it} + \pi_i \eta_{t+1} + \eta_{it+1} \\
x_{t+1} &= \rho x_t + e_{t+1} \\
x_{it+1} &= \rho_i x_{it} + e_{it+1} \\
\Delta d_{it+1} &= \mu_i^d + \phi_i^d x_t + \tilde{\phi}_i^d x_{it} + \pi_i^d \eta_{t+1} + \tilde{\pi}_i^d \eta_{it+1} + \eta_{it+1}^d
\end{aligned} \tag{2.13}$$

For a consumption paying domestic asset the Euler equation is:

$$1 = E_t \left[M_{it+1} R_{ict+1} \right]$$

with the stochastic discount factor in country i:

$$m_{it+1} := \log M_{it+1} = \theta \log \delta - \frac{\theta}{\psi} \log \left(\frac{C_{it+1}}{C_{it}} \right) + (\theta - 1) \log R_{ict+1}$$

equivalently:

$$m_{it+1} = \theta \log \delta - \frac{\theta}{\psi} \Delta c_{it+1} + (\theta - 1)\, r_{ict+1}$$

where ψ is IES, γ risk aversion and $\theta = \frac{1-\gamma}{1-\frac{1}{\psi}}$.

The return to the domestic dividend paying asset, R_{idt+1} assumes a similar Euler equation.
Following Bansal and Yaron (2004), we approximate returns as:

$$r_{ict+1} = k_{i0} + k_{i1} z_{it+1} - z_{1t} + \Delta c_{it+1}$$

$$r_{idt+1} = k_{i0} + k_{i1} z_{it+1}^d - z_{1t}^d + \Delta d_{it+1}$$

where

$$z_{it+1} = A_{i0} + A_{i1} x_{t+1}$$

$$z_{it} = A_{i0} + A_{i1} x_t$$

$$z^d_{it+1} = A_{i0} + A^d_{i1} x_{t+1} \tag{2.14}$$

$$z^d_{it} = A_{i0} + A^d_{i1} x_t$$

and $z^d = pd = \log(\frac{P}{D})$. The standard asset pricing condition is:

$$E_t \left[M_{USt+1} R_{i,t+1} \right] = 1$$

since

$$m_{USt+1} + r_{it+1} \quad = \quad \theta \log \delta - \frac{\theta}{\psi} \Delta c_{USt+1} + (\theta - 1) r_{USct+1} + r_{it+1}$$

the Euler Equation is equivalent to:

$$E_t \left[\exp \left(\theta \log \delta - \frac{\theta}{\psi} \Delta c_{USt+1} + (\theta - 1) r_{USct+1} + r_{it+1} \right) \right] = 1$$

for any asset from country i. We focus on the asset from the US. For any realization of the state variable, the following equation must be constant:

$$\theta \log \delta - \frac{\theta}{\psi} \Delta c_{USt+1} + (\theta - 1) r_{USct+1} + r_{USct+1}$$

Thus if you plug in the expressions for Δc_{USct+1}, r_{USct+1}:

$$\begin{aligned}
&= \quad \theta \log \delta - \frac{\theta}{\psi} \left(\mu_{US} + \phi_{US} x_t + x_{USt} + \pi_{US} \eta_{t+1} + \eta_{USt+1} \right) \\
&+ \quad (\theta) \left(\kappa_{i0} + \kappa_{i1} \left(A_{i0} + A_{i1} x_{t+1} \right) - \left(A_{i0} + A_{i1} x_t \right) \right) \\
&+ \quad (\theta) \left(\mu_{US} + \phi_{US} x_t + x_{USt} + \pi_{US} \eta_{t+1} + \eta_{USt+1} \right)
\end{aligned}$$

solving for A_1:

$$A_{i1} \quad = \quad \frac{\phi_{US} - \frac{\phi_{US}}{\psi}}{1 - \kappa_{i1} \rho}$$

similarly:

$$m_{it+1} + r_{idt+1} \;=\; \theta \log \delta - \frac{\theta}{\psi} \Delta c_{it+1} + (\theta - 1) r_{ict+1} + r_{idt+1}$$

equivalently

$$m_{USt+1} + r_{idt+1} \;=\; \theta \log \delta - \frac{\theta}{\psi} \Delta c_{USt+1} + (\theta - 1) r_{USct+1} + r_{idt+1}$$

this is equivalent to:

$$\begin{aligned}
m_{USt+1} + r_{idt+1} \;=\;& \theta \log \delta - \frac{\theta}{\psi} \Delta c_{USt+1} \\
& + (\theta - 1)(k_{US0} + k_{US1} z_{USt+1} - z_{USt} + \Delta c_{USt+1}) \\
& + k_{i0} + k_{i1} z^d_{it+1} - z^d_{it} + \Delta d_{it+1} \\
=\;& \theta \log \delta - \frac{\theta}{\psi} (\mu_{US} + \phi_{US} x_t + x_{USt} + \pi_{US} \eta_{t+1} + \eta_{USt+1})) \\
& + (\theta - 1)(k_{US0} + k_{US1}(A_{US0} + A_{US1} x_{t+1}) - (A_{US0} + A_{US1} x_t) \\
& + (\mu_{US} + \phi_{US} x_t + x_{USt} + \pi_{US} \eta_{t+1} + \eta_{USt+1})) \\
& + k_{i0} + k_{i1}(A_{i0} + A^d_{i1} x_{t+1}) - (A_{i0} + A^d_{i1} x_t) \\
& + (\mu^d_i + \phi^d_i x_t + \tilde{\phi}^d_i x_{it} + \pi^d_i \eta_{t+1} + \tilde{\pi}^d_i \eta_{it+1} + \eta^d_{it+1})
\end{aligned}$$

This yields:

$$A^d_1 \;=\; \frac{\phi^d_i - \frac{\phi_{US}}{\psi}}{1 - \kappa_{i1} \rho}$$

The demeaned stochastic discount factor in country i:

$$m_{it+1} - \mathbb{E}_t[m_{it+1}] \;=\; (-\frac{\theta}{\psi} + (\theta - 1))(\pi_i \eta_{t+1} + \eta_{it+1}) + (\theta - 1)(\kappa A_1 e_{t+1} + \kappa e_{it+1})$$

The demeaned return on consumption in country i:

$$r_{ict+1} - \mathbb{E}_t[r_{ict+1}] \;=\; \pi_i \eta_{t+1} + \kappa A_1 e_{t+1}$$

The demeaned return to dividends in country i:

$$r^d_{it+1} - \mathbb{E}_t[r^d_{it+1}] \quad = \quad \pi^d_i \eta_{t+1} + \kappa A^d_1 e_{t+1}$$

We assume m, e and r are jointly log-normal.

$$\begin{aligned} 0 &= \log(\mathbb{E}_t[\exp(r^d_{it+1} + \Delta e + m_{USt+1})]) \\ &= \mathbb{E}_t[r^d_{it+1}] + \mathbb{E}_t[\Delta e] + \mathbb{E}_t[m_{USt+1}] \\ &\quad + \frac{1}{2}\operatorname{var}(r^d_{it+1}) + \frac{1}{2}\operatorname{var}(\Delta e) + \frac{1}{2}\operatorname{var}(m_{USt+1}) \\ &\quad + \operatorname{cov}(r^d_{it+1}, \Delta e) + \operatorname{cov}(r^d_{it+1}, m_{USt+1}) + \operatorname{cov}(\Delta e, m_{USt+1}) \end{aligned}$$

Combining these terms gives the risk premium.
Total US return:

$$\mathbb{E}\left(r_{US} - r_{f_{us}}\right) = -\operatorname{cov}\left(m_{US}, r_{US}\right) - \frac{1}{2}\operatorname{var}\left(r_{US}\right)$$

Total foreign return:

$$\mathbb{E}\left[\hat{r}^e_i\right] \equiv \log \mathbb{E}\left[R_i\right] + \mathbb{E}\left[\Delta q_i\right] - \log \mathbb{E}\left[R_f\right] = -\operatorname{cov}\left(m_{US}, r_i\right) - \frac{1}{2}\operatorname{var}\left(r_i\right) - \frac{1}{2}\operatorname{var}\left(\Delta q_i\right),$$

where

$$\operatorname{cov}\left(m_{US}, r_i\right) = \left(-\frac{\theta}{\psi} + \theta - 1\right)\pi_{us}\pi^d_i\sigma^2_\eta + (\theta - 1)\kappa^2\left(\frac{\phi_{US} - \frac{\phi_{US}}{\psi}}{1 - \kappa\rho}\right)\left(\frac{\phi^d_i - \frac{\phi_{US}}{\psi}}{1 - \kappa\rho}\right)\sigma^2_e$$

$$\operatorname{var}(m_{US}) = \left(-\frac{\theta}{\psi} + \theta - 1\right)^2 \pi^2_{US}\sigma^2_\eta + (\theta - 1)^2\kappa^2\left(\frac{\phi_{US} - \frac{\phi_{US}}{\psi}}{1 - \kappa\rho}\right)^2\sigma^2_e$$

$$\operatorname{var}(m_i) = \left(-\frac{\theta}{\psi} + \theta - 1\right)^2 \pi^2_i\sigma^2_\eta + (\theta - 1)^2\kappa^2\left(\frac{\phi_i - \frac{\phi_i}{\psi}}{1 - \kappa\rho}\right)^2\sigma^2_e$$

$$\operatorname{var}\left(r_i\right) = \kappa^2\left(\frac{\phi^d_i - \frac{\phi_{US}}{\psi}}{1 - \kappa\rho}\right)^2\sigma^2_e + \left(\pi^d_i\right)^2\sigma^2_\eta$$

2.8.1 Kappas

We estimate κ using a symmetric balanced growth path, so that the terms are constant across countries. On a balanced growth path,

$$\begin{aligned}\bar{z} = A_0 &= \frac{\log\beta + \left(1 - \frac{1}{\psi}\right)\mu_{US} + \kappa_0}{1 - \kappa_1} \\ &= \frac{\log\beta + \left(1 - \frac{1}{\psi}\right)\mu_{US} + \log\left(1 + e^{\bar{z}}\right) - \frac{e^{\bar{z}}}{1+e^{\bar{z}}}\bar{z}}{1 - \frac{e^{\bar{z}}}{1+e^{\bar{z}}}}\end{aligned}$$

Similarly,

$$\begin{aligned}\bar{z}_m = A_{0m} &= \frac{\log\beta + \left(1 - \frac{1}{\psi}\right)\mu_{US} + \kappa_{0m}}{1 - \kappa_{1m}} \\ &= \frac{\log\beta + \left(1 - \frac{1}{\psi}\right)\mu_{US} + \log\left(1 + e^{\bar{z}_m}\right) - \frac{e^{\bar{z}m}}{1+e^{\bar{z}}m}\bar{z}_m}{1 - \frac{e^{\bar{z}}m}{1+e^{\bar{z}}m}}\end{aligned}$$

$$\kappa = \frac{e^{A_0}}{1 + e^{A_0}} = \frac{e^{A_{0m}}}{1 + e^{A_{0m}}}$$

Given $\mu_{US} = 0.021$, $\kappa = 0.9975$, which is consistent with the estimate of Bansal and Yaron (2004), who report $\kappa = 0.997$.

2.8.2 Risk-Free Rate

To derive the U.S. risk-free rate

$$\mathbb{E}_t\left[\theta\log\delta - \frac{\theta}{\psi}\Delta c_{USt+1} + (\theta - 1)r_{USct+1} + r_{ft+1}\right] = 0$$

which yields:

$$r_{f,t} = -\theta\log(\delta) + \frac{\theta}{\psi}\mathbb{E}_t\left[\Delta c_{USt+1}\right] + (1-\theta)\mathbb{E}_t\left[r_{USct+1}\right] - \frac{1}{2}\operatorname{var}_t\left[\frac{\theta}{\psi}\Delta c_{USt+1} + (1-\theta)r_{USct+1}\right],$$

following the approach Bansal and Yaron (2004), subtract $(1-\theta)r_{f,t}$ from both sides and divide by θ:

$$r_{f,t} = -\log(\delta) + \frac{1}{\psi}\mathbb{E}_t\left[\Delta c_{USt+1}\right] + \frac{(1-\theta)}{\theta}\mathbb{E}_t\left[r_{USct+1} - r_{f,t}\right] - \frac{1}{2\theta}\,\mathrm{var}_t\left[\frac{\theta}{\psi}\Delta c_{USt+1} + (1-\theta)r_{USct+1}\right.$$

2.9 Estimation of Consumption Growth Processes

Global consumption growth parameters. In our model, there are both global and idiosyncratic sources of risk, but only the former are priced by the U.S. agent. We model global consumption growth, measured by expression (2.1) in the data, as:

$$\Delta c_{Wt+1} = \mu_W + x_t + \eta_{t+1}, \tag{2.15}$$

where x_t is the global persistent component defined in expression (2.6) and idiosyncratic components have been averaged out.[17] The global process closely mimics the consumption growth process for the U.S. in Bansal and Yaron (2004). We will rely on second moments from the global consumption process in our identification strategy, so we need not specify a value for the mean growth rate, μ_W.

To identify the global persistent component, x_t, we follow the methodology in Colacito et al. (2018b) and we proceed in two steps. First, based on insights in Bansal et al. (2012), we exploit the model's prediction that a country's logged price-to-dividend ratio is a function of the global persistent process only (see expression (2.14) in Appendix 2.8). This implies that a projection of future consumption (or dividend) growth on lagged values of the (logged) price-to-dividend ratio is able to recover the time series of the persistent process. The challenge with this strategy is to estimate parameters pertaining to the "world" over a long period of time as we want to capture global long-run risks. As we describe in Section 2.2.3, we de-

[17]One can view this process as a special case of the following process: $\Delta c_{Wt+1} = \mu_W + \phi_W x_t + \pi_W \eta_{t+1}$, where the exposure parameters ϕ_W and π_W are normalized to unity.

fine the world to consist of five major economies, each denoted by k below: U.S., U.K., France, Germany and Japan, due to reliable data coverage during the 1940-2020 period. Tables 2.10 and 2.11 in Appendix 2.7 report summary statistics for consumption growth and price-dividend ratios for each country.

Specifically, we estimate the parameter α from the following pooled regression using data on all five countries during the 1940-2020 period:

$$\Delta c_{it+1} = \alpha \cdot pd_{it} + \epsilon_{it+1} \quad \forall t, i, \tag{2.16}$$

where pd_{it} is the logged price-dividend ratio in country i in year t. In the second step, we define the global persistent component as:

$$x_{t+1} \equiv \frac{1}{5} \sum_k \Delta \hat{c}_{kt+1} = \frac{1}{5} \sum_k \hat{\alpha} \cdot pd_{kt}, \tag{2.17}$$

where $\Delta \hat{c}_{kt+1}$ is a fitted value for country k of the pooled linear regression in expression (2.16).

Table 2.12: Global Persistent Component

	Dependent variable:		
	Δc_{t+1}		x_t
pd_t	0.006***	x_{t-1}	0.758***
	(0.001)		(0.073)
		Constant	0.005***
			(0.002)
Observations	395		79
R^2	0.149		0.584

Notes: Table reports (left) a pooled linear regression of country k's per-worker consumption growth on the log price-to-dividend ratio, pd_{kt}, and the estimated coefficient $\hat{\alpha}$, where k = U.S., U.K., France, Japan, Germany, and $K = 5$. On the right, the table reports the autoregression results of the global persistent process, x_t. $x_{t+1} \equiv \frac{1}{5} \sum_k \Delta \hat{c}_{kt+1} = \frac{1}{5} \sum_k \hat{\alpha} \cdot pd_{kt}$. Data are for 1940-2020 period from MacroHistory Database. Standard errors statistics in parentheses. $^{*}p < 0.1$; $^{**}p < 0.05$; $^{***}p < 0.01$.

We report the estimate of α in the left panel of Table 2.12. Our estimate of 0.006 compares favorably to the estimate of 0.005 in Colacito et al. (2018b),

Table 2.13: Summary Statistics for World (5 countries) Consumption Growth

	N	Mean	Median	Std. Dev	Q_{10}	Q_{90}	Constant	Δc_{Wt}	R^2
Δc_{Wt+1}	81	0.022	0.022	0.034	0.000	0.049	0.013**	0.459***	0.217
							(0.004)	(0.100)	

Notes: Table reports summary statistics of world consumption growth Δc_{Wt+1} and results of linear regression of Δc_{Wt+1} on its lag, Δc_{Wt}, where the world is computed as the average consumption of U.S., U.K., France, Germany and Japan. Data are for the 1940-2020 period from MacroHistory Database. Standard errors statistics in parentheses. $^{*}p < 0.1$; $^{**}p < 0.05$; $^{***}p < 0.01$

and similarly to the authors, we find that country-specific estimates of the parameter are not statistically different from each other, which supports the choice in favor of a pooled regression.[18] Having obtained a series for the global persistent component, x_t, we estimate ρ from an AR(1) regression corresponding to the second line in expression (2.6), and we report the results in the right panel of Table 2.12. The estimate amounts to a sizeable 0.76, which compares favorably to estimates reported by the existing literature (see for ex. Table 4 in Colacito and Croce (2011) for U.S. and U.K.).

We recover the variance of the persistent global shock, σ_e^2, from an autoregression of the global consumption growth process. Specifically, taking the time-series variance of Δc_W in expression (2.15) yields $\sigma_e^2 = \frac{(1-\rho^2)\beta_W^c \text{var}(\Delta c_W)}{\rho}$, where $\beta_W^C \equiv \frac{\text{cov}(\Delta c_{Wt+1}, \Delta c_{Wt})}{\text{var}(\Delta c_W)}$ is the coefficient estimate of the following regression:

$$\Delta c_{Wt+1} = \beta_W^C \Delta c_{Wt} + \epsilon_{t+1}, \tag{2.18}$$

and it is reported in the right panel of Table 2.13, along with summary statistics of the series in the left panel. The variance of the persistent component, σ_x^2, is a direct function of the variance of the innovations to the persistent component, $\sigma_x^2 = \sigma_e^2/(1-\rho^2)$. Finally, the residual variance of the transitory shock follows from the global consumption growth series, af-

[18] Colacito and Croce (2011) identify the persistent process from a projection of consumption growth on the price-dividend ratio and the risk-free rate of a country. For our sample of countries and period of study, the coefficient estimates of the risk-free rate are not statistically different from zero, so we exclude risk-free rates from the analysis as they do not seem to contain additional information.

ter accounting for the persistent component, $\sigma_\eta^2 = \text{var}(\Delta c_W) - \sigma_x^2$. This approach to recovering the persistent and transitory innovations to global consumption growth parallels Bansal and Yaron (2004), who aim to account for observed variations in consumption growth (in the U.S.) over a long horizon.

In Appendix 2.10, we re-estimate the key autoregressive parameter of the global persistent process, ρ, using historical data from the MacroHistory Database for six additional economies: Australia, Switzerland, Spain, Italy, Netherlands and Sweden. The resulting parameter estimate of 0.81 compares favorably to our benchmark and to estimates in the existing literature.

Idiosyncratic consumption growth parameters for the U.S. To identify the parameters that govern the U.S. consumption growth process in the first line of expression (2.6), we use the series described in Section 2.2.3 for the 1940-2020 period. Under the assumption that innovations are independent, we recover the consumption growth exposure parameter to the global persistent process for the U.S. from a linear regression of U.S. consumption growth on the global persistent process, x_t,

$$\Delta c_{USt+1} = \phi_{US} x_t + \epsilon_{t+1}. \tag{2.19}$$

Given the specifications for U.S. and global consumption growth in expressions (2.6) and (2.15), the exposure parameter to the global temporary shock follows from a regression of U.S. consumption growth on the residual component of global consumption growth,

$$\Delta c_{USt+1} = \pi_{US}(\Delta c_{Wt+1} - x_t) + \epsilon_{t+1}. \tag{2.20}$$

The results from these regressions are displayed in Table 2.14. The mean U.S. consumption growth rate is 2.1% as reported in Table 2.10 in Appendix 2.7 and corresponds to parameter μ_{US} in the consumption process. The leverage parameters to the temporary and persistent global shocks are precisely estimated and correspond to 2.13 and 0.42, respectively.

Table 2.14: U.S. consumption growth

	Dependent variable: Δc_{USt+1}	
	(1)	(2)
x_t	2.133*	
	(1.195)	
$\Delta c_{Wt+1} - x_t$		0.425***
		(0.059)
Constant	-0.027	0.020***
	(0.027)	(0.002)
Observations	79	79
R^2	0.04	0.401

Notes: Table reports results of linear regression of U.S. consumption Δc_{USt+1} on persistent process x_t, and on $\Delta c_{Wt+1} - x_t$ to estimate ϕ_{US} and π_{US}, respectively. Data are for the 1940-2020 period from MacroHistory Database. Standard errors in parentheses. $^{*}p < 0.1$; $^{**}p < 0.05$; $^{***}p < 0.01$

2.10 Global Variables: Robustness

As described in the main text, we define a 'global' variable to be the mean of five economies: U.S., U.K., France, Germany and Japan. In Table 2.15, we show that adding observations for Australia, Switzerland, Spain, Italy, Netherlands and Sweden yields estimates for the parameters that govern the global persistent shocks that are very similar to our benchmark estimates.[19]

Furthermore, we show that estimates of country-level exposures to global persistent shocks are robust to different definitions of global dividend growth. First, recall that, as we describe in Appendix 2.6, stock market capitalization data coverage begins in 1975, while dividend growth data begins in 1970 for developed economies. Hence, the key moment of interest that identifies dividend growth exposure to global persistent shocks–namely, the covariance of a country's dividend growth with the global dividend growth

[19] The MacroHistory Database also includes consumption data for Portugal, Canada, and Ireland, but price-dividend data are missing for these countries. Price-dividend observations for Belgium are not reliable during the post-war period as the variance is several orders of magnitude higher than what we observe for other countries. Data are available for Finland, Denmark and Norway, but these countries accounted for a negligible share of world economic activity historically. Furthermore, the financial variables from Norway are pooled from a variety of data sources and the coverage of stocks changes multiple times over the period of study, which makes the data difficult to interpret.

Table 2.15: Global Persistent Component, Robustness

	Dependent variable:		
	Δc_{t+1}		x_t
pd$_t$	0.006***	x_{t-1}	0.810***
	(0.000)		(0.068)
		Constant	0.004***
			(0.001)
Observations	845		79
R^2	0.166		0.648

Notes: Table reports (left) a pooled linear regression of country k's per-worker consumption growth on the log price-to-dividend ratio, pd_{kt}, and the estimated coefficient $\hat{\alpha}$, where k = Australia, Switzerland, Germany, Spain, France, U.K., Italy, Japan, Netherlands, Sweden and U.S. and $K = 11$. On the right, the table reports the autoregression results of the global persistent process, x_t. $x_{t+1} \equiv \frac{1}{11}\sum_k \Delta\hat{c}_{kt+1} = \frac{1}{11}\sum_k \hat{\alpha} \cdot pd_{kt}$. Data are for 1940-2020 period from MacroHistory Database provided by Jordà et al. (2019) and Jordà et al. (2017). Standard errors statistics in parentheses. *p<0.1; **p<0.05; ***p<0.01.

rate—is computed beginning in 1975, even though returns and dividend growth rates date back to 1970 for developed economies. We believe that it is important to weigh stock-market moments by stock market capitalization as it is standard practice when constructing an index; for ex. MSCI indices, that combine groups of countries, weigh countries' equity indices by stock market capitalization.[20] This weighting scheme aims to capture the importance of each country in driving global economic activity. Second, recall that our definition of global dividend growth reflects five countries: U.S., U.K., France, Germany and Japan.

In Table 2.16, we explore several different definitions of global dividend growth. Keeping all other parameters of the model fixed as in Table ?? the main text, we re-estimate π_i^d using expression (2.10) and ϕ_i^d from the covariance implied by expression (2.3) for each country, where Δd_W is derived under an alternative definition. In column (i), we let global dividend growth be the stock-market-capitalization weighted mean of all 37 countries in our sample. In column (ii), we focus on the countries that make up MSCI's World Index, which are the following developed economies:

[20] An individual country's index weighs firm observations by their respective stock market capitalization.

Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Ireland, Israel, Italy, Japan, Netherlands, New Zealand, Norway, Portugal, Singapore, Spain, Sweden, Switzerland, U.K. and U.S. In column (iii), we include the 11 countries from the robustness exercise that re-estimated the global persistent process in Table 2.15, and in column (iv), we include the G-10 countries (which add up to 11 including Switzerland): Belgium, Canada, France, Germany, Italy, Japan, Netherlands, Sweden, Switzerland, U.K., and U.S.

The mean level of ϕ_i^d increases as we increase the sample of countries that constitute the world. Compared to the mean in our benchmark specification in Table 2.5 of 6.21, the most significant increase occurs when we include all 37 countries in the definition of the 'world', which mechanically reflects the increased correlation of each country with the 'world'. But more importantly, in the cross-section of countries, the correlation of estimated ϕ_i^d's under the alternative specifications with our benchmark estimates is almost unity. This finding implies that different definitions of the 'world' do not affect the cross-sectional predictions of our model, and only affect the levels of risk premia.

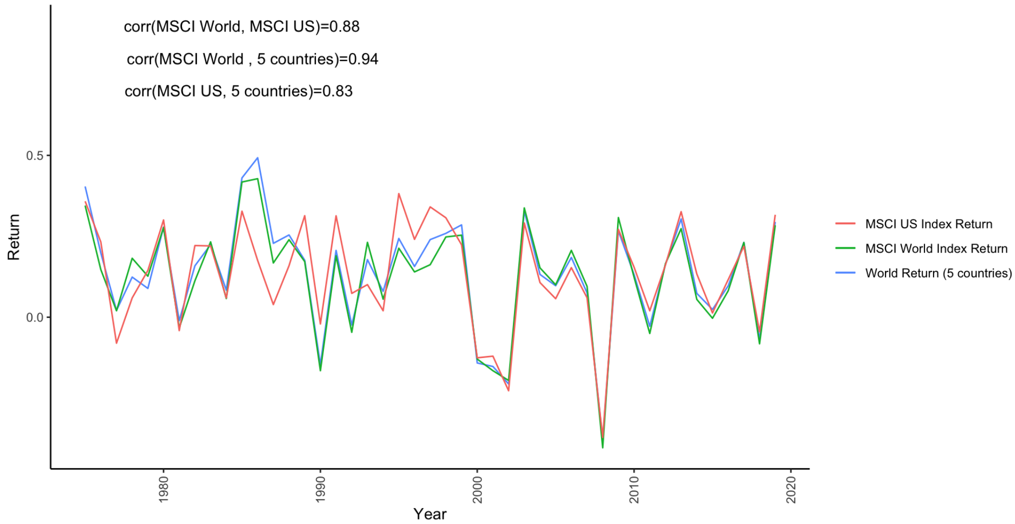
Finally, in Figure 2.10, we plot equity returns during the 1975-2020 period for: (i) the U.S., (ii) the definition of World Index provided by the MSCI, and (iii) our definition of a global variable, which constitutes the stock-market-weighted mean of returns for U.S., U.K., France, Germany, and Japan. It is apparent that the three series comove very closely, which reassures our definition of a global variable.

Table 2.16: Estimated ϕ_i^d Under Different Global Dividend Processes

	(i)	(ii)	(iii)	(iv)
Country	$\phi^d_{AllCountries}$	$\phi^d_{MSCIWorld}$	ϕ^d_{C-11}	ϕ^d_{G-10}
AUS	7.14	6.39	5.96	5.60
AUT	9.00	7.97	7.32	6.77
BEL	6.64	6.18	5.97	5.64
BRA	15.22	13.06	12.19	11.40
CAN	6.77	5.99	5.47	5.36
CHE	8.13	7.41	7.13	6.73
CHL	10.12	8.45	7.73	7.37
COL	11.25	10.00	9.24	8.78
CZE	5.75	4.93	4.15	3.68
DEU	10.69	9.74	9.29	8.81
DNK	8.24	7.63	7.33	7.01
EGY	13.43	11.19	9.96	10.03
ESP	8.27	7.49	7.04	6.46
FIN	12.16	10.87	9.98	9.79
FRA	8.12	7.47	7.12	6.69
GBR	5.93	5.55	5.32	4.97
HUN	10.17	9.32	8.32	8.06
IDN	13.40	11.25	10.20	9.64
IND	14.20	12.01	10.99	10.81
IRL	6.81	6.66	6.46	6.12
ISR	3.74	3.16	2.62	2.63
ITA	3.84	3.83	3.76	3.38
JPN	4.75	4.39	4.34	4.10
KOR	15.50	12.90	12.08	11.78
MEX	7.95	6.42	5.63	5.47
MYS	5.87	4.68	4.07	4.13
NLD	5.92	5.57	5.37	5.02
NOR	8.36	7.29	6.69	6.44
NZL	8.31	7.61	7.12	6.81
PER	12.71	11.38	10.68	10.33
PHL	7.66	6.68	5.98	5.84
PRT	9.04	8.25	7.60	7.26
SGP	6.06	5.28	4.93	4.65
SWE	10.08	9.37	8.95	8.49
TWN	9.41	7.91	7.47	7.18
USA	4.37	4.09	3.91	3.79
ZAF	8.20	7.12	6.58	6.35
mean	8.74	7.72	7.16	6.85
median	8.24	7.47	7.12	6.69
std. dev	3.11	2.59	2.40	2.34
corr(.,y_i)	-0.55	-0.51	-0.47	-0.48
corr(., ϕ_i^d)	0.98	0.99	1.00	1.00

Notes: Table reports country-level estimated parameters ϕ_i^d using different definitions of the Δd_W process. All countries refers to a stock-market weighted Δd_W using all countries in our sample. C-11 refers to a stock-market weighted Δd_W process using the 11 countries used for the robustness exercise in Table 2.15. G-10 refers to a stock-market weighted Δd_W process using G-10 countries. MSCI World refers to a stock-market weighted Δd_W process using countries that are in the MSCI World database. Data Sources: MSCI.

Figure 2.10: Time Series of Stock Market Returns



Notes: Figure plots various return series over time. Data Sources: MSCI.

2.11 Supporting Tables

Table 2.17: Realized Stock Market Returns Regression

	Dependent variable:		
	r_{it}		
	(1)	(2)	(3)
y_{it}	-0.041^{***}	-0.036^{***}	-0.082^{***}
	(0.014)	(0.011)	(0.028)
Constant	0.545^{***}		
	(0.155)		
Observations	1,466	1,466	1,466
R^2	0.006	0.007	0.006
Country fixed effects	N	N	Y
Year fixed effects	N	Y	N

Notes: Table reports the results of a linear regression of equity returns r_{it} on income per worker, y_{it}. Data Sources: MSCI, PWT and WDI. Standard errors statistics in parentheses. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$

Table 2.18: Realized Stock Market Returns Regression

	Dependent variable:		
	r_{it}		
	(1)	(2)	(3)
y_{it-1}	-0.050^{***}	-0.039^{***}	-0.117^{***}
	(0.014)	(0.011)	(0.027)
Constant	0.642^{***}		
	(0.151)		
Observations	1,502	1,502	1,502
R^2	0.009	0.009	0.013
Country fixed effects	N	N	Y
Year fixed effects	N	Y	N

Notes: Table reports the results of a linear regression of equity returns r_{it} on lagged income per worker, y_{it-1}. Data Sources: MSCI, PWT and WDI. Standard errors statistics in parentheses.

Table 2.19: Robustness Results

Condition	Statistic	N	Mean	Median	St. Dev.	Q_{10}	Q_{90}	corr$(\cdot, r_i^e)$	corr$(\cdot, \widehat{r_i^e})$
$\pi_i^d = \pi_{i,US}^d$ & $\phi_i^d = \phi_i^d$	$\widehat{r_{ri}^e}$	37	10.050	10.840	2.486	6.231	12.533	0.558	0.993
$\pi_i^d = \pi_{i,US}^d$ & $\phi_i^d = \phi_i^d(\pi_{US}^d)$	$\widehat{r_{ri}^e}$	37	9.299	10.761	4.084	2.865	12.568	0.201	0.581

Notes: Table reports summary statistics of predicted risk premia $(\widehat{r_{ri}^e})$ from the parameterized model under two alternative parameter estimates. The first specification sets $\pi_i^d = \pi_{US}^d$ and keeps ϕ_i^d as in the baseline. The second specification sets $\pi_i^d = \pi_{US}^d$ and re-estimates ϕ_i^d from the same moment conditions as in the baseline.

Chapter 3

De-Dollarization? Not So Fast

with Jon Hartley

Abstract

De-dollarization refers to the reduction of the reliance of foreign countries on the US dollar. This phenomenon generates concern about the U.S. dollar as a global currency. We construct new data on the currency denomination of central bank currency reserves, foreign exchange transaction volume, denomination of global debt securities, and the invoicing of trade. This paper presents empirical evidence suggesting that these concerns are misplaced, finding US dollar dominance remains unchanged up through late 2023, nearly two years after the 2022 Russian invasion of Ukraine and several years after the 2020 COVID-19 pandemic. Meanwhile euro and renminbi influence have since declined. These findings have implications for reserve currency resilience, U.S. dollar dominance, U.S. sanctions policy, international spillovers of U.S. monetary policy, and U.S. government borrowing costs.

3.1 Introduction

A unipolar currency world (as opposed to a multipolar currency world) is highly important in international finance as it can improve the hegemon's borrowing costs, stability in the international monetary system, the ease of international trade, and the ability of the hegemon to implement sanctions (Farhi and Maggiori (2018)).

In recent years, the future of dollar dominance, that is the status of the United States dollar as the dominant global currency (a position held since World War 2), has been called into question, owing to various geopolitical events and changing economic sentiments. The 2022 Russian invasion of Ukraine, accompanied by sanctions imposed by the United States on Russian foreign exchange reserve assets, has sparked discussions regarding the resilience of the dollar's dominance. Moreover, the emergence of anti-dollar sentiment, exemplified by the 2023 bilateral trade agreement between China and Brazil to reduce their reliance on the US dollar as an intermediary currency, has raised concerns about the future of the dollar's global standing. Financial institutions, such as Goldman Sachs, JP Morgan and State Street, have undertaken research efforts to dissect this issue (Chandan and Popescu (2023), Tse (2023), Thiagarajan et al. (2023)). However, many of these reports tend to disregard the extensive body of literature in international economics that supports the strength of dollar dominance across various dimensions, including trade invoicing, central bank currency reserves, foreign exchange transaction volume, and the denomination of global debt securities (see Maggiori (2017), Maggiori and Farhi (2018), He, Krishnamurthy, and Milbradt (2019), Eichengreen (2014) among others.). More recently, Maggiori, Neiman, and Schreger (2019) have offered insights into the increasing predominance of the dollar and the declining status of the euro, a trend that persisted until the mid-2010s, across various dimensions such as trade invoicing, central bank currency reserve denominations, foreign exchange transaction volume, and syndicated loans.

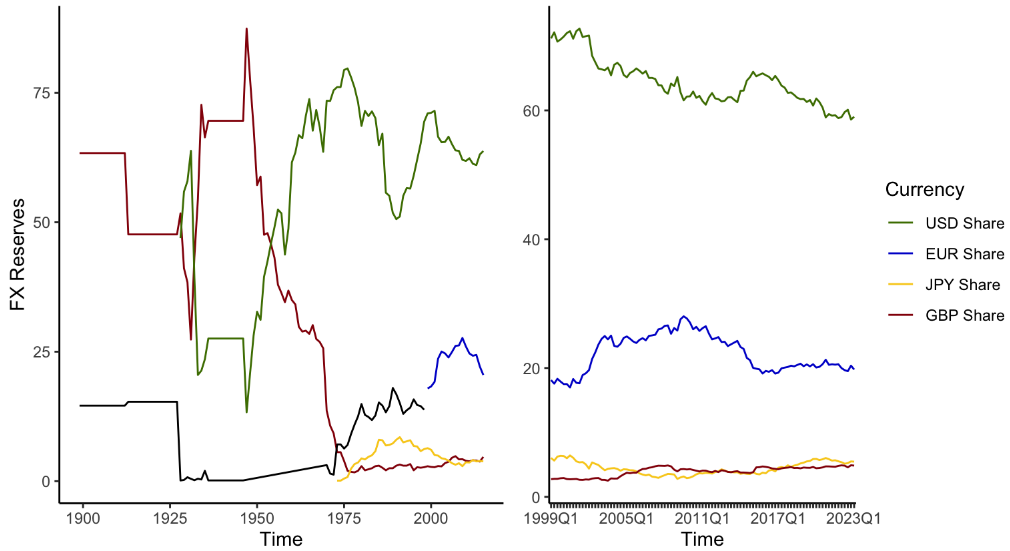
This paper aims to provide an update on the dollar dominance facts provided by Maggiori, Neiman, and Schreger (2019), utilizing more recent data that accounts for developments post COVID-19 and the 2022 Russian invasion of Ukraine, both of which were viewed as potential destabilizing factors in the global trade and financial landscape. Since dollar dominance is a multifaceted phenomenon, it is important to consider various indicators beyond just the recent decline in dollar-denominated foreign reserves. In fact, other key indicators, such as trading volume and the prevalence of global debt securities in dollars, present a contrasting perspective. To contribute to the existing literature and policy discussions, we analyze important indicators such as the proportion of central bank reserves denominated in dollar, the volume of foreign exchange transactions conducted in dollars and currency invoicing of US Dollars in global trade. Additionally, we provide data that encompasses the currency denomination of global debt securities, offering a more comprehensive measure of currency denomination in debt markets.

The paper is organized as follows. First we report new data on the proportion of central bank reserves denominated in dollars. The second section is displaying updated data on the volume of foreign exchange transactions conducted in dollars. We then discuss the percentage of global debt securities issued in dollars. Finally, we show evidence of non-declining currency invoicing in U.S. dollars in global trade.

3.2 Central Bank Reserves by Currency

We start our analysis by extending the time series of central bank currency reserves compiled by Eichengreen (2014). A widely referenced graph in the discourse on "de-dollarization" is Figure 1 (right), displaying central bank currency reserves from 1999 to 2023, sourced from the IMF COFER database. This graph might erroneously suggest that the U.S. dollar's share of currency reserves is at an all-time low. However, when we extend the dataset to 1900 (left), a different narrative unfolds. While the U.S. share of central bank currency reserves has declined since the early 2000s, it has since remained higher than its lows in the early 1990s when the U.S. dol-

Figure 3.1: Central Bank Foreign Exchange Reserves By Currency



Notes: Central bank foreign exchange reserves by currency denomination data from 1900 through 1999 (left) is annual from Eichengreen (2014) and obtained from Barry Eichengreen's website. Data from 1999 through 2023 (right) is quarterly data obtained from the International Monetary Fund (IMF) COFER database.

lar's share of currency reserves reached as low as 50.6 percent. Potentially driving the recent decline in the dollar share of central bank reserves since the early 2000s could be globalization which substantially increased the volume of international trade and established new global trading routes (Brancaccio, Kalouptsidi, and Papageorgiou (2022)).

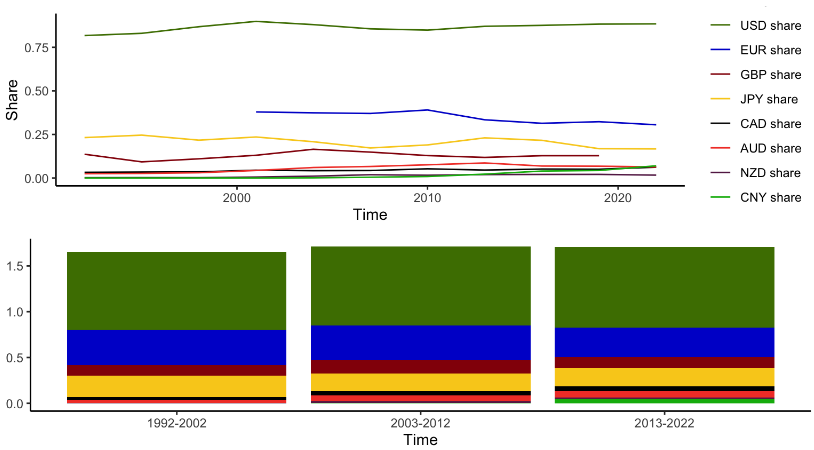
While not pictured, the Chinese reminbi share of global central bank reserves (as a fraction of total world central bank FX reserves) has fallen from 2.83 percent at the time of the 2022 Russian invasion of Ukraine in the first quarter of 2022 to 2.37 percent as of the third quarter of 2023.

This level is significantly below the current figure of 59.2 percent as of the third quarter of 2023. Notably, the U.S. dollar's share of currency reserves has experienced lower levels in the past, particularly during periods of inflation in the 1970s and 1980s, rebounding during the era of low and stable inflation associated with central bank inflation targeting, which began in the 1990s.

3.3 Foreign Exchange Transaction Volume by Currency

An examination of foreign exchange transaction volume by currency pairs, as depicted in Figure 3, underscores that the U.S. dollar's share in foreign exchange transactions has remained remarkably stable over the past 30 years since the inception of the BIS Triennial Survey. As of 2022, the U.S. dollar is involved in approximately 90 percent of all foreign exchange transactions. It is worth noting that this data was collected after the 2022 Russian invasion of Ukraine and the subsequent sanctions on Russian central bank currency reserves, reinforcing the U.S. dollar's enduring dominance in foreign exchange transaction volumes. Interestingly, the euro's share of foreign exchange transaction volume has declined from nearly 40 percent in 2001 to around 30 percent in 2022.

Figure 3.2: Foreign Exchange Transaction Volume By Currency



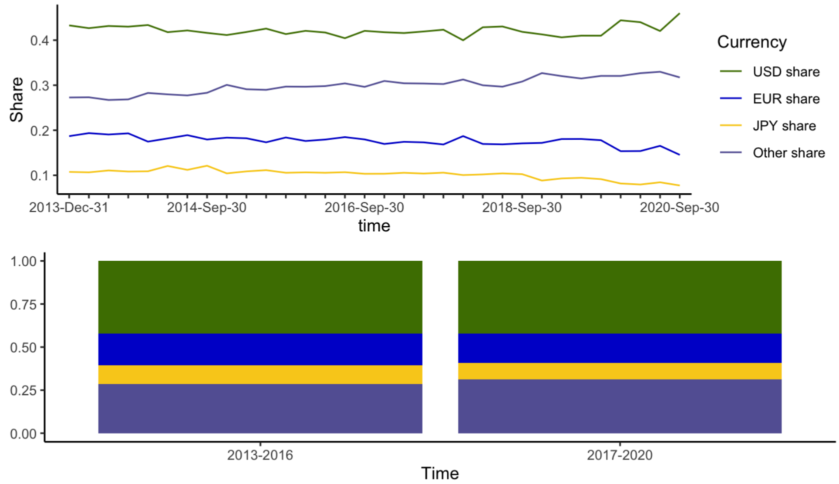
Notes: The upper graph displays the time series of the data, while the bottom one shows averages across time periods. Data on foreign exchange transaction volume from 1992 to 2022 is from the Bank of International Settlements (BIS) Triennial Survey. All shares sum to 200% (with a maximum of 100% for any single currency) since there are two currencies in every foreign exchange transaction. The U.S. dollar share of foreign exchange volume remains close to 90% through 2022 and has been close to such a level since the survey began in 1992.

3.4 Global Debt Securities Denomination by Currency

Regarding the denomination of global debt securities, the majority of the world's debt is still denominated in US dollars, even in the aftermath of the COVID-19 pandemic and the 2022 Russian invasion of Ukraine. Our analysis covers the currency denomination of global debt securities, encompassing both domestic and international debt, from the first quarter of 2015 to the first quarter of 2023, as presented in Figure 4. The dollar's share in debt contracts has remained relatively stable, hovering between 40 and 45 percent, with a slight increase following the events of the COVID-19 pandemic and the 2022 Russian invasion of Ukraine. Notably, corporate debt and firms' currency preferences significantly contribute to this pattern. Eren and Malamud (2022) have documented the widespread use of the U.S. dollar for debt contracts worldwide, even in the post-COVID-19 era. Subsequent analysis of BIS international debt statistics reaffirms this empirical observation.

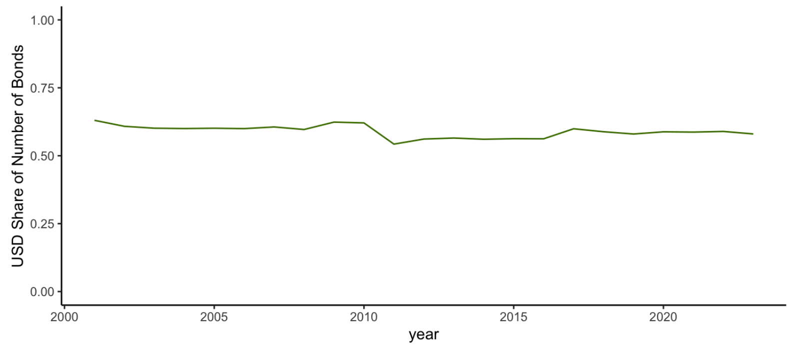
In the realm of sovereign borrowing, the phenomenon known as "original sin," referring to the denomination of external debt in foreign currencies, remains a persistent challenge, particularly in emerging markets, as originally highlighted by Eichengreen, Hausmann, and Panizza (2002, 2007). Onen, Shin, and Peter (2023) provide further insights into this matter.

Figure 3.3: Currency Denomination of Global Debt Securities Outstanding (Q1 2013 – Q4 2022)



Notes: The upper graph displays the time series of the data, while the bottom one shows averages across time periods. Data on global debt securities denomination by currency from Q1 2013 to Q4 2022 is from the Bank of International Settlements (BIS). The U.S. dollar share of global debt securities has been above 40% since 2013. At the end of 2022, the U.S. dollar share is now higher than what it has been since 2013. Meanwhile the euro and yen shares have decreased while other currencies have increased.

Figure 3.4: Share of the number of bonds denominated in USD



Notes: U.S. dollar share of the number of total bonds in the Bloomberg Fixed Income Multiverse. Data obtained from Bloomberg.

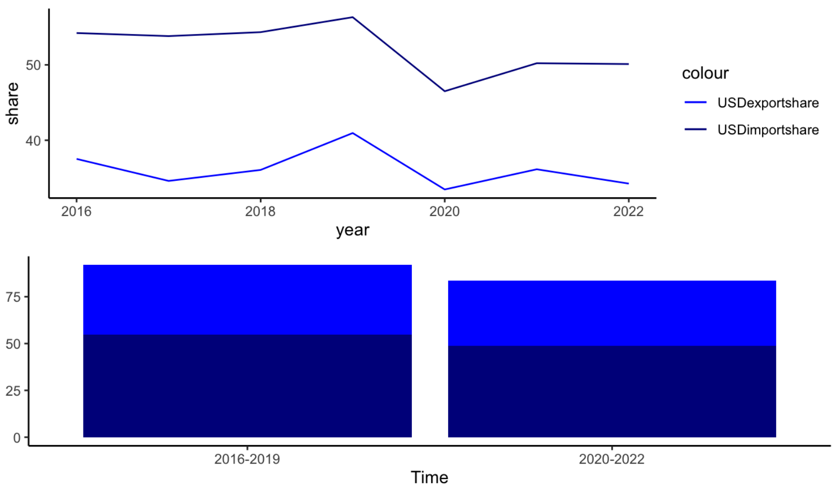
3.5 Currency Invoicing in Global Trade

Expanding upon the previous work of Gopinath et al. (2020), which computed trade invoicing dollar shares up to 2018, our paper updates this series with more recent data, including the pandemic years of 2022 and the Russian invasion of Ukraine. This was achieved by collecting data on trade invoicing currencies from the central banks, finance ministries, and government agencies of several major trading nations, including the United

Kingdom, all European Union member countries, Japan, and Australia. In the context of global currency invoicing in international trade, a significant proportion is still denominated in U.S. dollars, a pattern that has remained largely unchanged post-COVID-19 pandemic and the Russian invasion of Ukraine.

Figure 5, displaying a simple average from 2016 to 2022, illustrates that the U.S. dollar's share of exports has remained relatively constant, ranging between 47 and 54 percent. Similarly, the U.S. dollar's share of imports has seen little variation, fluctuating between 34 and 38 percent. This trend aligns with the findings of Gopinath et al. (2020) and Maggiori, Neiman, and Schreger (2019). Notably, there was a slight decline in U.S. dollar invoicing in 2020 during the peak of the COVID-19 pandemic, followed by a subsequent rebound.

Figure 3.5: Currency Denomination of Trade Invoicing (2016–2022)



Notes: The upper graph displays the time series of the data, while the bottom one shows averages across time periods. Data for the UK is from non-EU imports and exports, for the European Union imports and exports with extra-EU geopolitical entity partners, for Japan the component Ratio by Contract Currency in the Export and Import Price Index Decomposition, and Australia Export and Import invoicing currencies data. The U.S. dollar share of exports has approximately been between 30% and 40% since 2016 while the U.S. dollar share of imports has approximately been between 45% and 55%.

3.6 Conclusion

In conclusion, this paper seeks to offer an updated assessment of U.S. dollar dominance through the mid-2020s, considering the potential impact of both the COVID-19 pandemic and the 2022 Russian invasion of Ukraine, events that were perceived as potential disruptors in the global trade and financial landscape. Notably, following these events, the US dollar remains a prominent force in trade invoicing, central bank currency reserves, foreign exchange transaction volumes, and the denomination of global debt securities. The findings of Dooley, Folkerts-Landau, and Garber (2022), who argue that U.S. sanctions reinforce the dollar's dominance, as well as McDowell (2023), who suggests that such sanctions may provoke political backlash against the dollar, appear to find support in our analysis.

Furthermore, scholars such as Farhi and Maggiori (2018) and Maggiori (2017) offer theoretical models for understanding the dynamics of the international monetary system and the role of dominant global currencies. Historically, transitions between dominant currencies have often been prompted by wars and conflicts. The absence of such events in recent times, possibly due to international deterrence mechanisms, may have contributed to the enduring dominance of the U.S. dollar. Nonetheless, the rise of an internationally ascendant China, as documented by Clayton et al. (2022), could potentially pave the way for a transition to a multipolar currency regime in the coming decades. That said, the renminbi's share of central bank reserves has fallen since the 2022 Russian invasion of Ukraine as of late 2023.

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