Benchmarking and Sovereign Risk

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Abstract

Benchmark indices are the returns of a basket of eligible assets and are often used to evaluate investors' performance. Adding a stock to an equity index tends to increase the volatility of the stock's price. Yet, little is known about the impact of index inclusion on the volatility of bond prices, even as emerging markets (EM) seek the inclusion of local-currency sovereign bonds into global bond indices maintained by financial service providers. This paper shows that the inclusion insulates bond prices from global shocks, reducing volatility. Using micro-level data from global mutual funds, I exploit bond-level heterogeneity in investors' benchmarks. When there is a larger share of investors that benchmark to indices with EM assets holding the bond, the bond price exhibits less volatility from the global risk factor. Although benchmark inclusion increases the level of demand for a country's bonds, its damping effect on the elasticity of demand has a countervailing impact, making the overall impact of benchmark inclusion on borrowing ambiguous. I build a structural model to evaluate the quantitative implications. The level effect on the demand dominates, indicating that inclusion encourages more borrowing.

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1 Introduction

The Egyptian finance minister, in response to Egyptian bonds entering the JP Morgan GBI-EM index in 2022, stated that the efforts to "fulfill the requirements of index inclusion" had paid off[1]. Like Egypt, many emerging markets (EM) have regarded global bond index inclusions as milestones in the globalization of local-currency (LC) bond markets. Global mutual funds, a major type of creditor driving this recent wave of globalization, use bond indices as their performance benchmarks. Empirical studies on mutual funds that closely replicate benchmark weights suggest that adding assets to indices increases the volatility of asset prices (Ben-david et al., 2018) and credit supplies (Chari et al., 2022). Unlike previous work that has mainly focused on equities or a subset of investment funds, this paper offers empirical and quantitative analyses of the index inclusions of EM LC bonds, accounting for a wide range of approaches to replicating benchmarks across the entire investor base of global mutual funds. This paper shows that creating a desirable investor base—characterized by investors benchmarked to a country's assets[2]—through index inclusion reduces the volatility of LC bond prices in response to global shocks.

The overall finding is consistent with a mechanism known in micro-finance but not yet analyzed in any macro-finance context, nor in studies that endogenize decisions on the degree of index presence at both the extensive and intensive margins. Sovereign bonds are no longer priced solely by default probabilities when risk-averse managers use index weights to guide portfolio decisions. Rule-based index weights may not be directly related to the default risk of the underlying bonds, making yields less responsive to global shocks. I empirically validate this channel. Guided by conditional moments from Colombia's index inclusion, I estimate a sovereign default model incorporating this channel to examine the stabilizing role of benchmarking under different shocks.

To begin with, I introduce a binary classification of mutual funds based on their benchmark indices: funds whose benchmark indices include EM LC bonds, referred to as EM funds, and all other funds, referred to as dollar funds. Using cross-sectional holding data from the universe of US-domiciled funds, I document that dollar funds with benchmarks entirely unrelated to EM LC bonds are exposed to EM LC risk. Many existing studies on the short-run impact of index weights on asset prices focus exclusively on EM funds. This approach overlooks the potential mechanism by which global benchmarks influence the overall composition of creditors between EM and dollar funds over time. More importantly, the heterogeneity of benchmarks among bondholders provides a novel source of identification. The challenge of estimating the effects of index inclusion on the propagation of shocks is that countries endogenously select into indices. Hence, comparing outcomes of countries with and without index status introduces bias. The fact I uncovered makes the

[1] Ministry of Finance Arab Republic of Egypt (2022).

[2] A fund is considered benchmarked to a country's LC bonds if its performance benchmark indices include LC bonds issued by that country.

identification through the following comparison possible. Consider two index-eligible sovereign bonds issued by a country included in global bond indices. If one bond is only held by dollar funds, then it should not enjoy any benefit of index inclusion because the return of this bond does not directly affect the performance benchmark of dollar funds. By contrast, if another bond is only held by EM funds, then its price should reflect the effects of index inclusion over the global financial cycle. When the assignment of bondholders is as good as random, comparing the outcomes of these two bonds reveals the effects of index inclusion.

I develop a partial equilibrium model with rich creditor-side features. In this model, creditors derive utility from both the absolute return and the return relative to the LC bond index, whose weights are determined by an exogenous function estimated empirically. The model features two types of creditors who differ in their risk aversion and their choice of benchmark indices. The model predicts that a higher share of bonds held by EM funds leads to higher index weights and fewer bonds being re-priced under an increased default probability by risk-averse creditors during downturns. Therefore, ownership by EM funds mitigates the impact of changes of global risk factors to the yields of LC bonds.

Empirically, I construct a sample of global funds' portfolios, classify funds based on their benchmark indices, and test the theoretical predictions. I address the sorting of funds into various benchmarks and countries into policies related to index presence by exploiting bond-level time-varying creditor compositions while controlling for bond and country-time fixed effects. The estimates suggest that ownership by EM funds reduces the impact of VIX on bond yields, as the model predicts. I also conduct a placebo test using Colombia's index inclusion event in 2014: the set of funds benchmarking only to Colombian Peso bonds after 2014 should not reduce the shock transmission prior to index inclusion. The estimates show that creditor ownership effects shift from negligible levels prior to inclusion to negative values thereafter. Explanations based on heterogeneous financial frictions and sticky portfolios are inconsistent with the results of the placebo test.

To examine EM debt policies, I integrate creditors benchmarked to indices into a quantitative model of sovereign debt. The model features one-period LC bonds and exogenous processes for output and exchange rates, which can be interpreted as being affected by the global risk factor discussed in the empirical section. The price of bonds changes due to exogenous shocks and the announcement of an index inclusion. These features generate model moments comparable with a real index inclusion event. Moreover, a country makes optimal decisions on index inclusion or exclusion. A country's index status switches to inclusion when it voluntarily opts for inclusion and an approval from index providers arrives with an exogenous probability.

In the model, benchmarking affects borrowing through two mechanisms. First, the behavior of tracking index weights increases the demand for sovereign bonds included in the index and increases the price. Since the index weight depends on outstanding bonds, a country considers the consequence of marginal gains on index weights. While demand becomes less elastic upon inclusion, the disciplinary effects of inelastic demand could be offset by higher bond prices, leading to an ambiguous impact of index inclusion on borrowing. In principle, the existence of indices could improve a country's welfare. Second, an equilibrium effect of reduced default risk arises when the welfare-improving effects exist for certain states: the set of states in which a country chooses to default becomes smaller in a world with indices than in that without. This equilibrium effect implies higher prices when creditors are risk neutral. When creditors are risk averse, higher expected return due to lower default risk is offset by increased volatility of returns. This second mechanism offers a nuanced explanation of repaying debt: index inclusion is a reward for repaying debt because it overcomes the frictions in developed markets.

I calibrate the model to match the macro moments of the Colombian economy as well as the reduction in yields upon Colombia's index inclusion. The estimated quantitative model confirms the rise in bond prices under a country index status. In the quantitative exercise, I analyze the transmission of exchange rate and output shocks. I decompose the transmission of these two shocks by considering two counterfactual cases: one where the country, compared to the baseline under index inclusion, has a zero index weight, and another without benchmarking. These two cases differ in the amount of debt a country borrows, as the quantity of debt outstanding affects the propagation of shocks. Quantitatively, if countries had been excluded from indices, the observed external LC debt-to-output ratio would be 4% lower. Benchmarking is more potent at insulating against output shocks, while it has limited, though positive, effect on insulating exchange rate shocks. The reason behind this divergence is that global shocks to the exchange rate directly reduce the country's index weight as the index weight often follows market value. Output shocks have no direct effect on the index weight.

Related literature This paper contributes to the debate on how passive, index-driven investment and, more broadly, capital flows from mutual funds affect financial market volatility. The existing literature establishes that benchmarking increases market volatility. For equities, Appendix D in Kashyap et al. (2020) provides a theoretical argument, and Ben-david et al. (2018) presents empirical evidence on ETFs. Studies examining capital flows from ETFs and index funds report similar findings (Chari et al., 2022; Converse et al., 2020). The insulation role of benchmarks in this study contrasts with existing results mainly in equity markets and contributes to the debate on the financial stability implications of index-driven investors (Anadu et al., 2020).

As Anadu et al. (2020) note, compared to equities, there is relatively little research on how benchmarking affects valuations in bond markets. Existing studies of bond index inclusion do not address bond valuations beyond a short window of a few days, leaving longer-term pricing effects understudied. This paper examines the valuation of risky bonds both in secondary markets with fixed supply and in primary markets where issuers internalize the equilibrium effects of benchmarking. While the mechanism of perfectly inelastic demand driven by benchmarks is not new, this study explores how such demand propagates shocks, a question that remains underexamined in the literature.

Moreover, this paper builds on an extensive finance literature and recent macroeconomic studies on benchmarking and mutual fund capital flows, particularly those that use changes in benchmark weights for identifications (Broner et al., 2021; He and Beltran, 2022; Pandolfi and Williams, 2019; Moretti et al., 2024). Rather than focusing on index rebalancing or inclusion events, I investigate how indices affect sovereign risk.

My empirical approach extends research analyzing mutual fund credit supplies and asset prices, including studies on funding shocks across different types of funds (Calomiris et al., 2022), capital allocations (Chari et al., 2022; Chari, 2023; Raddatz et al., 2017), and the transmission of funding shocks to asset prices (Jotikasthira et al., 2012; Zhou, 2023). Unlike most of these works with the exception of Zhou (2023), I conduct a bond-level analysis with a particular focus on EM LC markets.

The model employs a reduced-form approach where asset managers derive utility from both absolute returns and returns relative to a benchmark, following studies that provide micro-foundations for benchmarking behavior (Kashyap et al., 2018, 2020; Duffie et al., 2014). I extend results from microfinance research on benchmarks to local-currency debt markets, which have been extensively studied regarding sovereign risk (Du et al., 2017; Ottonello and Perez, 2019; Engel and Park, 2022; Lee, 2022). Previous studies have not analyzed the impact of benchmark indices on EM LC risk. My creditor-centered analysis is related to studies on banks and sovereign risk (Bocola, 2016; Morelli et al., 2022; Perez, 2015). In contrast to these studies that focus on amplification mechanisms, I find that benchmarking acts as a stabilizing mechanism in the context of EM LC bonds.

Structure of the paper In Section 2, I introduce a new classification of mutual funds and present a new stylized fact on capital flows to EM bond markets using this classification. Section 3 presents a model with benchmarking and derives a testable prediction. In Section 4, I construct a dataset and empirically test the model's prediction. Section 5 quantifies the degree of insulation provided by benchmarking. Conclusions are drawn in Section 6. The appendix contains further evidence of EMs' index-related policies and analytical derivations.

2 A new stylized fact on benchmarking and capital flows

2.1 A brief background of indices

Mutual funds supply more credit to emerging market local currency bond markets than other types of private creditors, such as deposit-taking banks, pension funds, insurance firms, and hedge funds[3]. Moreover, mutual funds ubiquitously use performance benchmarks—indices that measure the performance of market portfolios whose underlying constituents and portfolio weights follow rules set by index providers.

Index providers set country- and bond-specific eligibility criteria. The most commonly used indices that include EM LC bonds are the FTSE Global Bond Index, the Bloomberg Global Aggregate Index, the JP Morgan Government Bond Index-Emerging Market Global Diversified (GBI-EM), along with their variants. The first two are global indices because they include developed-market sovereign bonds and corporate bonds. The JP Morgan GBI-EM index only includes EM LC sovereign bonds. Variants of these indices arise from tightening or loosening some eligibility criteria so that the index becomes broader for gauging overall market performance or narrower for replication by investors.

Income levels and market access hurdles, such as capital controls and liquidity, are two important criteria for country eligibility. A country may experience index inclusion or exclusion when its income changes. For example, the JP Morgan GBI-EM Broad Diversified included Czech Republic in February 2017 after JP Morgan reclassified it as an emerging market country based on GNI per capital (Broner et al., 2021). Index providers require certain degree of market "liquidity". This condition differentiates all-inclusive broad market indices from narrowly defined indices for replication.

Moreover, indices set bond-level criteria related to maturity, issuance size, and coupon types[4]. For example, in the JP Morgan GBI-EM Broad Diversified index, the size is greater than $1 billion for domestic bonds and $500 million for global bonds, the residual maturity is greater than two and half years, and the coupon type is restricted to fixed-rate non-inflation-index coupons. Once index providers select a set of bonds, they often use market values to determine portfolio weights. Providers may fulfill a diversification objective by capping a country's index weight[5].

2.2 A new stylized fact

I propose a binary classification of mutual funds based on their performance benchmarks. I classify funds based on whether or not their performance benchmark indices include emerging market

[3]See Appendix B for a comparison of mutual fund EM LC bond holdings with the holdings from other types of creditors using ECB and U.S. TIC data.

[4]Interested readers may refer to Arslanalp et al. (2020) for more details of these rules.

[5]For more on exploiting this feature to construct capital flow shocks that may be exogenous to a country's fundamentals, I refer interested readers to Pandolfi and Williams (2019), He and Beltran (2022), and Moretti et al. (2024).

local-currency denominated sovereign bonds. I refer to these funds as EM and dollar funds.

I validate this classification by showing that funds whose benchmarks exclude EM LC bonds are actively exposed to EM LC risk. I apply this classification to portfolios reported between December 2019 and February 2020 from the universe of mutual funds domiciled in the US. I use portfolio data from mutual funds' mandatory filings in Form N-PORT-P[6]. Form N-PORT-P is used by investment companies to disclose their portfolios to the U.S. Securities and Exchange Commission. The data are accessible and free to the public. I describe this dataset in Appendix A.2.

Table 1: Characters of US bond funds holding (USD bn) EM LC sovereign debts reported between 12/2019 and 02/2020

Fund benchmarks	Total	Common EM countries		Other developing countries	
		Non-index fund	Index fund	Non-index fund	Index fund
Include EM LC bonds (*EM fund*)	46.95	29.40	12.67	4.72	0.16
Exclude EM LC bonds (*Dollar fund*)	33.67	27.77	0.01	5.89	0

Notes: This table summarizes the distribution of local-currency sovereign bond holdings across the universe of U.S. fixed-income funds. Common EM countries are those that have ever been included in EM LC bond indices. They are Argentina, Brazil, Chile, China, Colombia, Czechia, Dominican Republic, Egypt, Hungary, Indonesia, Malaysia, Mexico, Peru, Philippines, Poland, Romania, Russian, South Africa, Thailand, and Turkey. Other countries are Costa Rica, India, Kazakhstan, Sri Lanka, Pakistan, Uruguay, Ukraine, and Venezuela.

Table 1 lists the EM local-currency credit supplied by US fixed-income mutual funds in Q4 2019, based on mandatory filings in N-PORT-P. This binary classification detects investment strategies that deviate from the benchmark index at the extensive margin. The key insight from the table is that over 40% of credit supplied to EM local-currency bond markets is unrelated to the asset type underlying mutual fund performance benchmarks. Moreover, index funds, whether they are ETFs or not, all have active strategies in the U.S. and have discretion over country weights and the degree of benchmark index tracking errors. That is why index funds are exposed to $160 million in local-currency bonds issued by countries which are absent from indices widely used as benchmarks.

Appendix Table A.2 summarizes non-benchmark driven credit across different categories of mutual funds, such as allocation, alternative, and equity funds. EM and dollar allocation funds

[6]See Sikorskaya (2023) for another application of the data from Form N-PORT-P.

hold EM LC bonds with market values \$1.14 billion and \$0.77 billion respectively. The empirical evidence in existing studies has not yet documented the size of EM LC holdings from allocation funds. The total size of credit to EM LC markets is small from US allocation funds, and this fact motivates me to focus on fixed-income funds throughout my analysis.

To provide further context on non-benchmark related capital flows, existing studies have almost exclusively focused on EM funds. Studies such as He and Beltran (2022) and Pandolfi and Williams (2019) that examine the effect of changing country weights on benchmarks have focused only on funds whose benchmarks are related to the asset class they study. An exception is Calomiris et al. (2022), which analyzes EM corporate bond holdings by funds which mainly invest in bonds issued by developed markets. Another exception is Arslanalp et al. (2020). It mentions benchmark-driven funds versus those tracking global bond indices, termed unconstrained funds due to small portfolio weight in EM countries.

The new stylized fact is that, among mutual funds investing in emerging market local-currency sovereign bonds, there is considerable heterogeneity in benchmarks. For instance, funds benchmarked to dollar-denominated assets often invest in emerging market local-currency bonds. Recognizing this heterogeneity is important to the profession, as researchers of emerging markets who lack access to the complete universe of global mutual fund portfolios often take a shortcut by selecting only mutual funds benchmarked to emerging market debt. However, my stylized fact shows that constructing a dataset based solely on mutual funds benchmarked to emerging market local-currency indices omits 42% of capital flowing into emerging markets. Acknowledging this heterogeneity offers a nuanced view of capital volatility by mutual fund types: certain funds may exhibit more "flightiness" than others and warrant closer examination.

3 The transmission of global shocks with benchmarking creditors

I present a model that includes benchmark indices and creditor portfolio decisions. Despite its simplicity, the model provides insights into the determinants of how secondary market prices respond to shocks.

3.1 Global asset managers

Two types of asset managers of a unit mass domicile in the developed economy and are managing mutual funds. Mutual funds publish the indices used as performance benchmarks. The binary classification of benchmark indices divides funds into EM and dollar funds. Throughout this paper, I use the $j \in \{\text{EM fund}, \text{Dollar fund}\}$ to denote the type of manager or fund. Each type of manager receives funding D_0. The mass of managers managing EM funds is $\mathcal{M}_{\text{EM fund}} = 1 - \mu$. The risk-bearing ability of managers $\gamma_j \in \{\gamma_L, \gamma_H\}$ with $\gamma_L < \gamma_H$ characterizes their investment strategies. I assume that managers of EM and dollar funds have risk aversions of γ_L and γ_H, respectively.

A market index prescribes a portfolio whose weights follow a specific rule. EM funds use an index that includes EM LC bonds, while dollar funds use an index that excludes them. Index providers construct indices for EM funds according to a rule summarized by the function $\Omega(.)$. Time-varying weights depend on the market value of EM bonds, denoted as $s_t^{-1} q_t B_{t+1}$. q_t is the price of the bond denominated in local currency. B_{t+1} is the principal value of a one-period EM LC bond issued at time t and maturing at time $t+1$. s_t is the real exchange rate . A rising s_t means depreciation. I summarize the weight of EM LC bonds w_j^b in the benchmark index for a fund of type j below.

$$w_{j,t}^b = \begin{cases} \Omega(s_t^{-1} q_t B_{t+1}) & \text{if } j = \text{EM Fund} \\ 0 & \text{if } j = \text{Dollar Fund} \end{cases} \tag{1}$$

The manager of type j chooses a portfolio consisting of claims $b_{j,t+1}^{DM}$ on developed market risk-free bonds and local-currency bonds $b_{j,t+1}$. Managers have a mean-variance utility based on a weighted sum of the absolute return, $W_{j,t+1}$, and the return relative to the portfolio prescribed by their benchmark index, $W_{j,t+1} - W_{j,t+1}^b$. The weight assigned to the absolute return is $\alpha \in (0,1)$. The portfolio problem for the manager of type j is:

$$\begin{aligned} \max_{b_{j,t+1}^{DM}, b_{j,t+1}} \; & \mathbb{E}_t(\alpha W_{j,t+1} + (1-\alpha)(W_{j,t+1} - W_{j,t+1}^b)) - \frac{\gamma_k}{2} \mathrm{Var}_t(\alpha W_{j,t+1} + (1-\alpha)(W_{j,t+1} - W_{j,t+1}^b)) \\ \text{s.t. } & D_0 = q_t^{DM} b_{j,t+1}^{DM} + s_t^{-1} q_t b_{j,t+1} \\ & W_{j,t+1} = b_{j,t+1}^{DM} + s_{t+1}^{-1}(1 - d_{t+1}) b_{j,t+1} \end{aligned} \tag{2}$$

where q_t^{DM} is the price of risk-free developed market bonds b_{t+1}^{DM}, and q_t is the local-currency price of EM bonds. $1 - d_{t+1}$ is the fraction of the principal value in local-currency terms recovered when the bond matures. $d_{t+1} = 1$ means default, and $0 < d_{t+1} < 1$ means partial default.

3.2 Discussion of managers' utility function

The main assumption regarding asset managers is that the relative returns influence their decisions. This functional form is directly from Kashyap et al. (2020), which derives the optimal benchmark compensation contract between investment firms and asset managers of mutual funds. A higher $1 - \alpha$ implies more frictions in the asset management industry. Investment firms, on behalf of savers, use compensation contracts to overcome two frictions. One is managers' private portfolio-management costs for risky assets, which are borne by asset managers and not shared with other agents. The other friction is unobserved portfolios. As a result, investors cannot directly instruct managers to hold an optimal portfolio, as would be the case in a frictionless world.

The main difference between this paper and Kashyap et al. (2020) is the introduction of heterogeneous asset managers and the assumption that they have different benchmarks. Kashyap et al. (2020) considers the contract for only one representative asset manager and leaves aside the question of why manager's benchmarks differ from one another in practice.

In Appendix C.5, I address this question by extending Kashyap et al. (2020) with two additional assumptions. I assume that the private cost of management depends on the type of asset under management. Moreover, investment firms cannot punish managers based on positive returns relative to benchmarks: an asset manager's salaries cannot be deducted if his performance is better than any benchmarks.

A final note on the utility function is that benchmarking under a CRRA preference implies benchmarking alters the risk aversion of managers. α affects risk aversion so that the risk tolerance becomes γ_k/α. I formalize this point in Appendix C.

3.3 Market equilibrium

To align this stylized model of one-period bonds with the empirical analysis using secondary market prices, I introduce two sub-periods between the issuance and maturity of bonds, occurring between time t and $t+1$. I consider how changes in the global risk factor from z_0 to z_1 during these two sub-periods affect the current and expected exchange rates as well as the expected probability of default. Throughout this section, I use the yield to maturity, y. The timeline in the secondary market is:

1. Given the outstanding local-currency debt B, the exchange rate s_0, the expected value of repayment per unit of bond at time $t+1$, $\mathbb{E}(\psi_0) \equiv \mathbb{E}(s_{t+1}^{-1}(1-d_{t+1})|z_0)$, the yield corresponding to the secondary market clearing price is y_0. The index weight is $w_0^b = w_0^b(s_0^{-1}y_0^{-1}B)$. It is useful to define a new variable θ_0 to measure the bond ownership by the EM fund:

$$\theta_0 \equiv \frac{\text{Market value of bonds held by EM Fund}}{\text{Total market value}}$$

2. In the second sub-period, a negative shock to the risk factor z occurs. The exchange rate unexpectedly changes to s_1 and the expected value of repayment is $\mathbb{E}(\psi_1) \equiv \mathbb{E}(s_{t+1}^{-1}(1-d_{t+1})|z_1)$. Creditors adjust their portfolios, and the bond yield y_1 clears the market.

The market demand B_ν^d for local-currency bonds in two sub periods $\nu = 1, 2$ from mutual funds satisfy the asset demand equation:

$$D_0 \sum_{j \in J} \mathcal{M}_j w_j(s_\nu, B, z_\nu) = s_\nu^{-1} q(s_\nu, B_\nu^d, z_\nu) B_\nu^d, \quad \nu = 0, 1 \tag{3}$$

Given an exogenous bond supply B, an exogenous process of risk factor z_ν, the asset market clearing condition is

$$B_\nu^d = B, \quad \nu = 0, 1 \tag{4}$$

The equilibrium in the secondary market consists of exogenous processes, prices, and portfolio weights derived from the first-order conditions of portfolio optimization problems, ensuring that the asset market clears.

To understand how benchmarks affect prices, I derive the portfolio of EM funds prior to the shock.

$$w_{\text{EM Fund},0} = \frac{\frac{\mathbb{E}(\mathcal{E}')}{q_0(1+r^f)} - 1}{\gamma_L (1+r^f) D_0 \operatorname{Var}(\frac{\mathcal{E}'}{q_0(1+r^f)})} + (1-\alpha) w^b_{\text{EM Fund}} \left(s_0^{-1} q_0 B \right) \tag{5}$$

where the expected value of repayment prior to the risk factor shock is

$$\mathcal{E}' = \frac{s_0}{s_{t+1}} \times (1 - d_{t+1}) \equiv s_0 \psi_0 \tag{6}$$

Throughout the analytical derivations, I use the approximation $\operatorname{Var}(\frac{\mathcal{E}'}{q_0(1+r^f)}) \approx \operatorname{Var}(s_0\psi_0)$. This approximation omits the second-order effect of prices when the difference between the EM bond yield and risk-free rate is small. The portfolio weight for dollar funds has a similar expression. Imposing the market clearing condition, $B_\nu^d = B, \nu = 0, 1$, the price of bonds in local currency is

$$q_0 = \frac{\mathbb{E}(s_0\psi_0)/(1+r^f)}{\underbrace{E_\gamma^{-1}}_{(1)} \operatorname{Var}(s_0\psi_0)(1+r^f) D_0 \underbrace{\left(\frac{s_0^{-1} q_0 B}{D_0} - (1-\mu)(1-\alpha) w_0^b\right)}_{(2)} + 1} \tag{7}$$

where the term for average risk aversion is $E_\gamma = \frac{1-\mu}{\gamma_L} + \frac{\mu}{\gamma_H}$. The price depends on two components: (1) the average risk aversion of creditors; and (2) the bond supply in excess of benchmark weights.

An observation from the above pricing equation—Eq. 7—is that the price under a positive index weight is always higher than that with a zero weight. Appendix C.1 shows the proof. This observation is consistent with the large literature on the short-run valuation effect of index inclusion (Broner et al., 2021; Shleifer, 1986). It is useful to analyze how benchmarking affects the slope of the demand curve. I examine how changes in the index weight affect the responsiveness of demand to price changes. When there is an exogenous increase in the benchmark weight, $\frac{\partial^2 B^d}{\partial q \partial w^b} > 0$. This result holds regardless of how the weighting function is specified. In other words, the inactive components of demand imply that demand becomes more inelastic.

3.3.1 Creditor compositions and the transmission of shocks

I analyze the impact of a risk factor shock that leads to an unexpected currency devaluation, represented by a rise in s_0, and a lower expected repayment value, represented by a decline in $\mathbb{E}(\psi_0)$. I perform a first-order approximation of the market clearing condition. I use $\wedge$ to denote the log deviation from the initial value in the secondary market. To simplify notation, I define the following variables: $V_0 \equiv (1+r^f)D_0 \operatorname{Var}(\frac{s_0(1-d_1)}{s_1})$, $R_0 \equiv \mathbb{E}(\frac{s_0(1-d_1)}{s_1})$, $w_0 \equiv \frac{s_0^{-1}B}{y_0 D_0}$, and $\epsilon^b \equiv \frac{\partial w^b}{\partial(s_0^{-1}B/y_0)} \cdot \frac{s_0^{-1}B}{y_0}$.

The impact consists of supply-side effects through the devaluation of the dollar value of the debt, both directly due to the risk factor shocks on the exchange rate and indirectly through the response of bond prices, as shown in (1) in equation 8. This effect includes changes in the value of outstanding bonds (a) and the benchmark weight (b).

$$
\begin{aligned}
\Big[1+\frac{(1+r^f)V_0}{R_0 y_0 E_\gamma}\overbrace{\Big(\underbrace{w_0}_{(a)}-\underbrace{\epsilon^b(1-\mu)(1-\alpha)}_{(b)}\Big)}^{(1)}\Big]\hat{y}_0 = &\underbrace{-\Phi_R}_{(2)}\hat{z}_0 + \\
&\Phi_v \hat{z}_0 \frac{(1+r^f)V_0}{R_0 y_0 E_\gamma}\underbrace{\left(w_0-(1-\mu)(1-\alpha)w_0^b\right)}_{(3)} + \\
&\Phi_s \hat{z}_0 \frac{(1+r^f)V_0}{R_0 y_0 E_\gamma}\left(\left(-w_0+\epsilon^b(1-\mu)(1-\alpha)\right)\right)
\end{aligned}
\tag{8}
$$

On the demand side, the risk factor shock affects bonds through direct impacts on the expected value in part (2) and indirect effects on risk premium of active portfolios (3). Risk factor shocks have direct impacts on the risk premium, as in $\Phi_v \equiv \frac{z_0}{V_0}\frac{\partial V_0}{\partial z_0}$, the expected return, as in $\Phi_R \equiv \frac{z_0}{R_0}\frac{\partial R_0}{\partial z_0}$, and the exchange rate, as in $\Phi_s \equiv \frac{z_0}{s_0}\frac{\partial s_0}{\partial z_0}$. The key effect of creditor composition works in part (3): a higher θ_0 implies a higher w_0^b and hence the risk factor shock affects less outstanding debt. Rewriting the above expression by substituting w_0^b with θ_0 yields the following result.

Proposition 1 *The following relationship holds in the secondary market*

If $w^b > 0$, then

$$\hat{y}_0 = \left(-\Phi_s + \Phi_1\left(-\Phi_R + \Phi_s + \Phi_v\Phi_2\right)\right)\hat{z}_0$$

where $\Phi_1 = \left(1+\frac{w_0-\epsilon^b(1-\mu)(1-\alpha)}{\phi+\frac{E_\gamma}{(1+r^f)V_0}}\right)^{-1}$, $\Phi_2 = \frac{\phi}{\phi+\frac{E_\gamma}{(1+r^f)V_0}}$, *and* $\phi \equiv (1-\theta_0)w_0(1+\frac{(1-\mu)\gamma_H}{\mu\gamma_L})$. *The term* Φ_1 *and* Φ_2 *are decreasing in* θ_0 *when* $w_0 > \epsilon^b(1-\mu)(1-\alpha)$.

If $w^b = 0$, then $\phi = w_0$ and θ_0 does not enter into the term $\Phi_1\left(-\Phi_R+\Phi_s+\Phi_v\Phi_2\right)$.

Appendix C.2 shows the proof.

If a rising z_0 represents a negative shock that depreciates the currency, lowers the expected repayment, and raises the yield, then

$$\frac{\partial \hat{y}_0}{\partial \hat{z}_0} = -\Phi_s + \Phi_1 \left(-\Phi_R + \Phi_s + \Phi_v \Phi_2\right) > 0$$

Currency devaluation implies that $\frac{\partial s}{\partial z} > 0$ and $-\Phi_s < 0$. Given that $\Phi_1 > 0$ under the assumption $w_0 > \epsilon^b (1-\mu)(1-\alpha)$. The term $\left(-\Phi_R + \Phi_s + \Phi_v \Phi_2\right)$ must be greater than zero for the yield to rise.

The above proposition implies the following

$$\frac{\partial^2 \hat{y}_0}{\partial \hat{z}_0 \partial \theta_0} \begin{cases} \frac{\partial \Phi_1}{\partial \theta_0} \left(-\Phi_R + \Phi_s + \Phi_v \Phi_2\right) + \Phi_1 \Phi_v \frac{\partial \Phi_2}{\partial \theta_0} < 0, & \text{if } w^b > 0, \Phi_v > 0, \text{ and } w_0 > \epsilon^b (1-\mu)(1-\alpha) \\ = 0, & \text{if } w^b = 0 \end{cases} \tag{9}$$

A prediction of the current model with benchmarks is that the higher θ_0 is, the smaller the changes in bond yields. Other than the requirement of the weighting function, the prediction only requires that the shock to the risk factor increases the volatility of the expected return, which depends on changes in the exchange rate s_0 and the expected value of repayment ψ_0. This condition is likely to be met when the risk factor shock generates a large currency deviation.

When the valuation effect on index weight is small ($w_0 - \epsilon^b (1-\mu)(1-\alpha) > 0$), as the EM fund creditor share θ_0 increases, the coefficients Φ_1 and Φ_2 decrease[7]. Therefore, exchange rate shocks have less impact on bond yields when θ_0 is high. Another feature is the term w_0 enters the expression $\Phi_1 \left(-\Phi_R + \Phi_s + \Phi_v \Phi_2\right)$. This suggests an empirical strategy to construct the empirical counterpart of w_0 when testing the model's prediction. The second part of the proposition says that θ_0 would not affect the transmission of the exchange rate shocks to bond yields if benchmarks were irrelevant to portfolio decisions. This is intuitive because the creditor composition reflects the current risk factor driving expected return and should not predict future shocks, which determines changes in yields.

Index inclusion and asset price volatility Does index inclusion make asset prices less or more volatile to the global risk factor? The answer to this question depends on

$$\frac{\partial \hat{y}_0}{\partial \hat{z}_0}\Big|_{\text{exclusion}} - \frac{\partial \hat{y}_0}{\partial \hat{z}_0}\Big|_{\text{inclusion}}$$

and ultimately on how index inclusion changes the value of $\Phi_1 \left(-\Phi_R + \Phi_s + \Phi_v \Phi_2\right)$. As Φ_2 decreases upon index inclusion, the volatility depends on Φ_1, whose value decreases if $w^b - \epsilon^b - \frac{E_\gamma}{(1+r^f)V_0} \frac{\epsilon^b}{w^b} > 0$. A sufficient condition for reduced volatility is that the benchmark weight is insen-

[7] ϵ^b is generally small. See footnote 19 for additional information.

sitive to changes in current bond prices ($\epsilon^b = 0$). In practice, benchmark weights are infrequently updated, and accordingly $\epsilon^b = 0$, implying that the Φ_1 term decreases and index inclusion reduces volatility from the global risk factor. The following proposition summarizes the result.

Proposition 2 (Equivalence) *When index weights are infrequently updated, there is no response of the weight to changes in current prices. This weighting rule is a sufficient condition for reduced bond price volatility to the global risk factor. The impact of index inclusion, $\frac{\partial \hat{y}_0}{\partial \hat{z}_0}|_{exclusion} - \frac{\partial \hat{y}_0}{\partial \hat{z}_0}|_{inclusion}$, to the first order is $-\gamma$, where*

$$\hat{y}_0 = \alpha \hat{z}_0 + \beta \hat{z}_0 \times w_0 + \gamma \hat{z}_0 \times w_0 \times \theta_0$$

Appendix C.3 shows the proof. This proposition also states that, to the first order, the effect of index inclusion on the transmission of global shocks to bond prices is identical to the coefficient on the EM fund share in an OLS regression.

4 Testing the stabilizing role of benchmarking

4.1 Constructing samples of global fund holdings and identifying non-benchmark driven credit

From the variables "primary prospectus benchmark" and "secondary prospectus benchmark" in Morningstar, I can retrieve information related benchmark indices. However, I am unable to accurately identify mutual funds exposed to EM LC bonds, particularly those that have already been liquidated, using existing variables on Morningstar. My objective is to identify a representative sample of global mutual funds that can be classified as EM and dollar funds without downloading the holdings of the entire universe of funds in Morningstar.

I describe a procedure to identify EM LC positions from EM and dollar funds based on the accessible commercial data from Morningstar. My focus is on fixed-income funds due to the presence of small positions in U.S. allocation funds, which invest in both stocks and bonds, as observed in public filings. I compile a list of funds domiciled in developed markets, including Canada, developed European countries[8], Hong Kong, Japan, Singapore, the United Kingdom, and the United States. Morningstar categorizes funds into over 200 detailed categories based on their investment objectives and locations. I identify Morningstar categories mostly likely containing funds exposed to EM LC bonds and extract the positions from all the funds within these categories.

[8]Belgium, Denmark, Finland, France, Germany, Ireland, Italy, Luxembourg, Netherlands, Norway, Portugal, Spain, Sweden, Switzerland, and United Kingdom.

I proceed in two steps. First, using Form N-PORT-P filing data, which covers the universe of U.S. funds between 2019 and 2022, I identify Morningstar categories containing funds exposed to EM LC bonds. I then identify the corresponding categories for EU funds. I exclude seven U.S. categories with the least exposure— representing 3.7% of the total U.S. fund exposure in Q4 2019—and their corresponding non-U.S. fund categories.

Moreover, Morningstar calculates the percentage of fixed-income holdings in different currencies over time for each fund. Their calculation has a good coverage for surviving funds. I compile a list of all the fixed-income funds domiciled in Luxembourg and Ireland and calculate the share of bonds denominated in EM currencies in their portfolios. After identifying funds with exposure to EM currencies, I obtain a list of the Morningstar categories to which they belong. I then find the corresponding categories for funds domiciled in Canada, Singapore, Hong Kong, Japan, and other EU countries.

These two steps result in 61 Morningstar categories, 6200 funds, and 160,000 quarterly portfolios. I download portfolios from Morningstar Direct, add bond characteristics by merging a dataset from Bloomberg, and identify EM LC positions.

The primary and secondary benchmarks of funds that have ever been exposed to EM LC bonds yield approximately 1,000 unique indices. I proceed with my country-level binary index classification using publicly available documentation from index providers. Appendix Table D.6 lists historical country inclusion and exclusion events in major indices. Appendix A.4 outlines the steps taken to clean fund holdings data. Appendix A.1 details the procedure for identifying unique bonds and constructing a dataset of bond characteristics and yields using Bloomberg Terminal.

The creditor composition θ_{it} for bond i from country $c(i)$ measures the proportion of a bond's face or market values held by mutual funds benchmarked to indices which include the issuer's country, relative to the total face or market values held by global funds.

$$\theta_{it} = \frac{\sum_{f|c(i)\in\mathcal{B}_{ft}} H^m_{ift}}{\sum_f H^m_{ift}}$$

where H^m_{ift} is the local-currency face value (m = face) or dollar market value (m = market) of a bond held by mutual fund f. A fund is benchmarked to bonds issued by country c at time t if its performance benchmark $\mathcal{B}_{ft}$ includes LC bonds from country c. A related variable $I_{\text{fund},it}$ measures the proportion of the market value of a bond held by global funds relative to the total net assets of the funds holding it. This measure is the direct analog of $w_t = s_t^{-1} q_t B_t / D_0$, the market value relative to the funding D_0, in the model.

$$I_{\text{fund},it} = \frac{\sum_f H^{market}_{ift}}{\sum_f \text{Net Asset}_{f(i)t}} \tag{10}$$

Table 2: Bond-level summary statistics

	Funding	EM fund share	Funding residualized	EM fund share residualized
Mean	0.78	59.85	0.00	0.00
Median	0.70	62.35	-0.04	1.17
Std Dev	0.43	20.35	0.40	18.27

Note: This table summarizes the measures of funding, $I_{\text{fund},it}$, and EM fund share, θ_{it}, and the residual values of them after controlling covariates in eq (11), (residual maturity and country × time and bond fixed effects). Statistics are computed at the bond level, then averaged across bonds. The sample includes 650 bonds with an issuance amount exceeding one billion dollars and a residual maturity of more than two years, denominated in the following currencies: BRL, CLP, COP, CZK, HUF, IDR, ILS, MXN, MYR, PEN, PLN, RON, THB, TRY, and ZAR.

Table 2 presents the summary statistics of key variables: bond-level funding, $I_{\text{fund},it}$, and EM fund share, θ_{it}, and their residual values after controlling the covariates of residual maturity, bond and country × time fixed effects. On average, the share of a bond held by EM funds is 59.85%, with a standard deviation of 20.35%. These values align with the notion that mutual funds often invest in assets outside of their performance benchmarks' underlying asset class. The last column in Table 2 shows a considerable variation in the EM fund share even controlling country × time fixed effects.

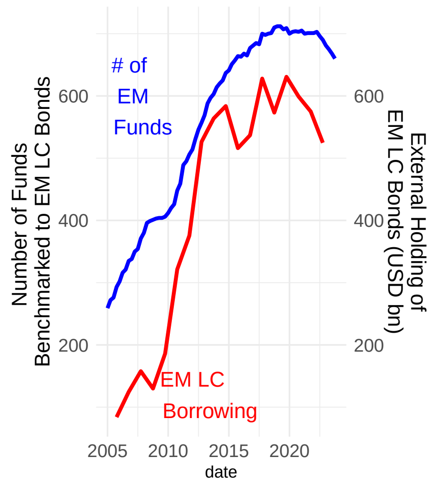
I describe two key facts unique to my setting with benchmarks.

Entries of EM funds and globalization of local-currency sovereign bonds. Figure 1 compares the number of EM funds with the total face value of externally held EM LC bonds over the past two decades. The two statistics show a correlation: the stagnation in fund entries around the 2008 crisis corresponded with a decline in borrowing, and both the number of EM funds and the borrowing plateaued in 2018 before subsequently declining. Given that the net assets of EM funds have remained stable over time, this correlation suggests that the sorting of funds as EM funds contributes to the fluctuations in EM LC borrowing.

Extensive margin responses. To further examine the extensive margin responses, I use current bond ownership status to predict future status within two years. Specifically, I estimate at the bond level with and without conditional on a fund's survival. I estimate

$$D_{if,t+h} = \alpha_i + \alpha_{c(i)t} + \beta D_{if,t} + \gamma D_{if,t} \times \mathbf{1}\{i(c) \in \mathcal{B}_{ft}\} + \epsilon_{ift}$$

Figure 1: Benchmark usages and EM LC borrowing



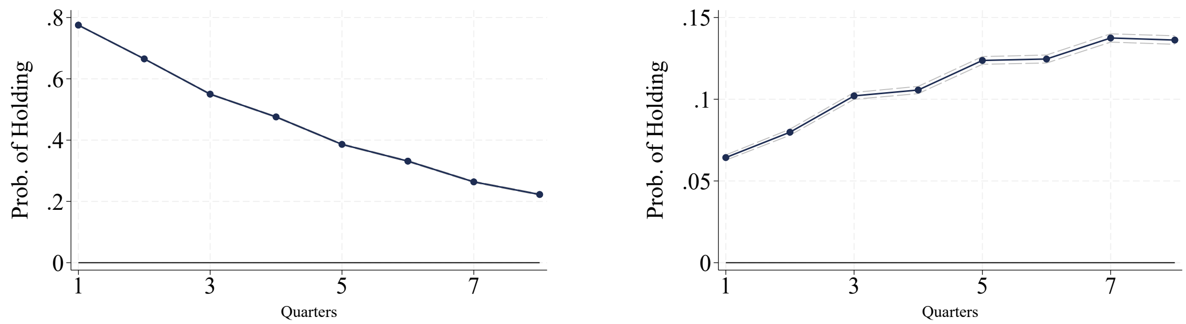
Notes: To measure the number of EM funds over time, I obtain a list of EM funds based on the binary classification of their benchmark indices. Given that each fund has multiple share classes, I use the earliest inception and the latest obsolete dates among share classes as a fund's surviving period. The total EM LC borrowing is summing the total externally held debts from EM countries (BR, CO, CL, CZ, HU, ID, IL, MY, MX, PE, PH, PL, RO, RU, TH, TR, ZA) using the data from Onen et al. (2022). In the borrowing data, the valuation is the face value of debts in dollars whenever the data exists.

where the asset ownership indicator $D_{if,t}$ switches on if fund f holds bond i at time t. The indicator $\mathbf{1}\{i(c) \in \mathcal{B}_{ft}\}$ is 1 if an EM fund holds bond i. β indicates the persistence of the asset ownership, while γ indicates the difference in ownership between EM and dollar funds. $D_{if,t+h}$ for $1 \leq h \leq 8$ may switch from 1 to 0 when a fund sells a bond or when a fund is liquidated. To examine the contribution of mutual fund exits to observed positions, I estimate the model conditional on the fund not being liquidated during the entire period from t to $t + h$. Appendix Figure E.5 shows estimates unconditional on fund liquidation. After two years, the probability of holding a bond is only 20% for dollar funds, whereas EM funds have a slightly higher probability of holding the same bond. Fund exits play a minimal role in this observed pattern.

4.2 Empirical results of creditor benchmarking and the pass-through of global shocks to LC bond yields

The main challenge in testing the model implication of asset ownership is the sorting of creditors into benchmark indices and the selection of countries into policies that ultimately affect their index presence based on unobserved variables. I address this issue by using time-varying creditor compositions at the bond level by controlling for country-time fixed effects. Given that shocks at the bond level affect trading and ultimately bond prices, sorting at the bond level would not be

Figure 2: Extensive margin of fund holdings



(a) Extensive margin, fund holding effect

(b) Extensive margin, benchmarking effect

Notes: For a bond currently in a fund's portfolio, these two graphs show the hazard rate that a fund continues holding this bond for two years conditional on the date being one year before a fund's liquidation. The estimation controls bond and country-time fixed effects. The first graph shows the hazard rate of dollar funds, and the second shows the difference in hazard rates between dollar and EM funds.

a concern if the following assumptions hold: a global fund does not select an EM LC index as its performance benchmark upon inception, or a country does not implement a policy related to its index presence because a single bond experiences relatively lower trading costs and relatively higher expected returns compared to other bonds from the same country.

I use the pass-through of the global risk factor VIX to secondary-market yields y_{it} to test the model implications from Section 3. I use VIX because it is a widely used approximation of the global risk factor in the literature (Miranda-Agrippino and Rey, 2021). The estimation equation is:

$$\begin{aligned}\log(1+y_{it+h})-\log(1+y_{it-1}) &= \beta^h \Delta\log(\text{VIX}_t)\times I_{\text{fund},it-1}+\gamma^h\Delta\log(\text{VIX}_t)\times I_{\text{fund},it-1}\times\theta_{it-1}\\ &+\alpha_i^h+\alpha_{c(i)t}^h+\Gamma^h\mathbf{X}_{it}+\epsilon_{it},\quad h=0,1,\ldots,8\end{aligned} \tag{11}$$

The focus is on the coefficient γ— the effect of creditor composition towards EM funds on the pass through of global shocks to yields. The vector $\mathbf{X}_{it}$ includes $I_{\text{fund},it-1}$, θ_{it-1}, and bonds' residual maturities. Figure 3 panels 3a) and 3b) show that a 1 percentage point increase in VIX raises bond yields, with the peak effect of 0.09 percentage points occurring three quarters after the impact. The credit ownership by EM funds is associated with a smaller impact of VIX on bond yields, with the reduction peaking at 0.11 percentage points four quarters after the impact. Both estimates are significant at the 90% confidence level under robust standard errors.

I also check robustness using face values to construct $I_{fund,it}$ and θ_{it}. $I_{fund,t}$ represents the proportion of the face amount in local-currency units of a bond held by funds relative to the total issuance in local-currency units:

$$I_{\text{fund},it} = \frac{\sum_f H_{ift}^{face}}{\text{Amount Issued}_i}$$

Figure 3 panels 3c) and 3d) confirm that a higher creditor ownership share from EM funds is associated with a lower pass-through of the VIX to bond yields under specifications using face values. Moreover, within this specification, I divide bonds based on whether they are issued in the global or domestic market. Panels E.6a) and E.6b) in Figure E.6 show that the response of global bonds is much more pronounced.

The potential limitation comes from portfolio decisions of many sovereign bonds at the same time: the yield of bond i depends on the ownership and the funding related to this bond and *all other bonds*. Therefore, estimates may be biased due to time-varying omitted variables at the bond level.

To access the degree of this limitation, I check the robustness of the empirical test by interacting $\Delta \log(\text{VIX}_t)$ with two additional time-varying controls: $\bar{I}_{fund,c(-i)t-1}$ and $\bar{I}_{fund,c(-i)t-1} \times \bar{\theta}_{c(-i)t-1}$. $\bar{I}_{fund,c(-i)t}$ averages the fund credit supplies across all bonds issued by country c, excluding bond i at time t. The same method is applied to construct a leave-one-out average of EM fund ownership, $\bar{\theta}_{c(-i)t}$. These two controls capture the influence of all other bonds issued by the same issuer. Appendix Figure E.6 panels E.6c) and E.6d) show that the creditor ownership effect becomes slightly more negative, with a peak reduction of 0.14 percentage points.

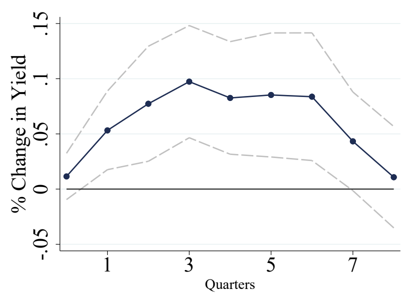
Another minor limitation is that, at the bond level, my classification does not distinguish whether an index includes specific bonds. I address this issue by selecting sovereign bonds eligible for the JP Morgan GBI-EM index, which has the most stringent eligibility criteria. These bonds are likely to be included in other bond indices with less stringent criteria.

4.3 Robustness to alternative explanations

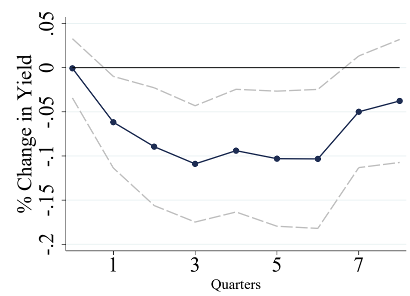
This section addresses the concern that the empirical finding may stem from other types of heterogeneity between EM and dollar funds, rather than their differences in the use of benchmark indices.

Models with alternative interpretations of VIX's impact on EM borrowing. The VIX correlates with a global risk factor and common components of EM exchange rates. However, the VIX is not a well-identified shock. Funding shocks (or creditor wealth shocks) among dollar funds could be alternative structural shocks generating previous results: the VIX correlates more strongly with larger funding shocks among dollar funds than among EM funds. I indirectly address this concern

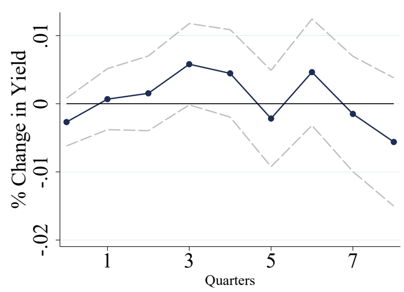
Figure 3: Effect of benchmarking



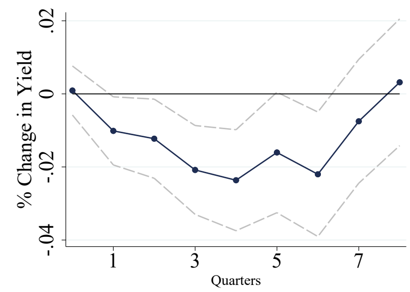
(a) VIX, fund holding effect (market value measures)



(b) VIX, benchmarking effect (market value measures)



(c) VIX, fund holding effect (face value measures)



(d) VIX, benchmarking effect (face value measures)

Notes: These figures show the direct effect of bond holdings by global funds and the interaction effect of the creditor composition of these funds on the impact of a 1% rise in the VIX on bond prices. The sample of bonds is restricted to bonds with issuance greater than one billion dollars and residual maturities greater than two years. Currencies are BRL, CLP, COP, CZK, HUF, IDR, ILS, MXN, MYR, PEN, PLN, RON, THB, TRY, ZAR. Dotted lines plot the 90% robust standard error bands.

by presenting results from short-term LC interest rate shocks originating in individual countries. Since the main mechanism of benchmarking is equivalent to reducing the quantity of assets subject to shocks, ownership of long-term bonds by EM funds would imply a smaller impact of rising short-term bond yields on long-term yields.

I use identified EM monetary policy shocks as proxies for short-term interest rate shocks. The results in this section are the weakest in the paper, and future tasks should extend the sample of countries and provide a more thorough analysis.

These shocks are identified using changes in forward premia around monetary policy announcements. I collect the dates of EM monetary policy press releases from central banks and measure changes in the one-year forward premium within a four-day window around each release, specifically the difference between the two-day average before and after the release. Short-term LC interest rates y_t^{LC} depend on the CIP wedge λ_t, the U.S. short-term rate $y_{US,t}^{\$}$, and the forward premium $fp_{c,t,t+1} = \frac{F_{lc/\$,t,t+1}}{E_{lc/\$,t}}$, as shown below. The identifying assumption is that the unexpected component of the monetary policy announcement does not affect the CIP wedge or U.S. interest rates. Witheridge (2024) empirically assesses this assumption and demonstrates this high-frequency approach identifies EM monetary policy shocks.

$$y_t^{LC} \times \lambda_t = y_{US,t}^{\$} \frac{F_{lc/\$,t,t+1}}{E_{lc/\$,t}} \tag{12}$$

Let $\Delta\tilde{y}_t^{LC}$ denote the monetary policy shock. I calculate changes in bond yields $\Delta \log(1 + y_{ict})$ as the difference between the 7-day average yields before and after the monetary policy press release. I then estimate how short-term borrowing rates are passed through to the actual borrowing costs of long-term bonds y_{it} across different bonds using the baseline specification, which includes country-time and bond fixed effects.

The estimates in Table 3 show that bond ownership by EM funds reduces the pass-through of short-term interest rate shocks to long-term bond yields. The overall point estimate of the interaction term with creditor composition measure is negative.

Models with financial frictions and sticky portfolios. Another explanation for the existing results is that dollar funds face tighter financial constraints than EM funds. Possible financial frictions include the costs of raising funds from households and value-at-risk type constraints from risk management. These frictions suggest that EM funds are likely to maintain a sticky portfolio of EM LC bonds, thereby dampening the transmission of shocks through their bond ownership. If this explanation holds true, then EM funds' bond ownership would reduce shock transmission even before a country's index inclusion. I use Colombia's inclusion in the JP Morgan GBI-EM index to test another prediction of benchmarking effects. Specifically, the bond ownership by global funds

Table 3: Pass through of EM short-term interest rate shocks

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
$\Delta \log(1 + y_{ict})$	All	Brazil	Colombia	Mexico	Peru	Philippines	Turkey
$\Delta\tilde{y}^{lc}_{c,t} \times I_{\text{fund},it-1}$	2.185*	0.598	-1.904	3.631**	1.407	1.334	3.321
	(0.949)	(1.699)	(1.050)	(1.110)	(1.681)	(0.778)	(5.777)
$\Delta\tilde{y}^{lc}_{c,t} \times I_{\text{fund},it-1} \times \theta_{i,t-1}$	-1.852	2.953	10.23	-3.293*	5.519	-1.274	-5.036
	(1.213)	(4.465)	(17.01)	(1.561)	(3.563)	(0.939)	(8.713)
Obs	8261	371	961	1439	1052	2917	1521

Notes: All specifications include country-time and bond fixed effects. Time-varying controls are residual maturities, the amount held by global funds, and the share held by EM funds. * $p < 0.1$,** $p < 0.05$ *** $p < 0.01$

benchmarked to the GBI-EM index should not affect shock transmission before Colombia's index inclusion in 2014, whereas the same ownership should reduce the transmission after the index inclusion.

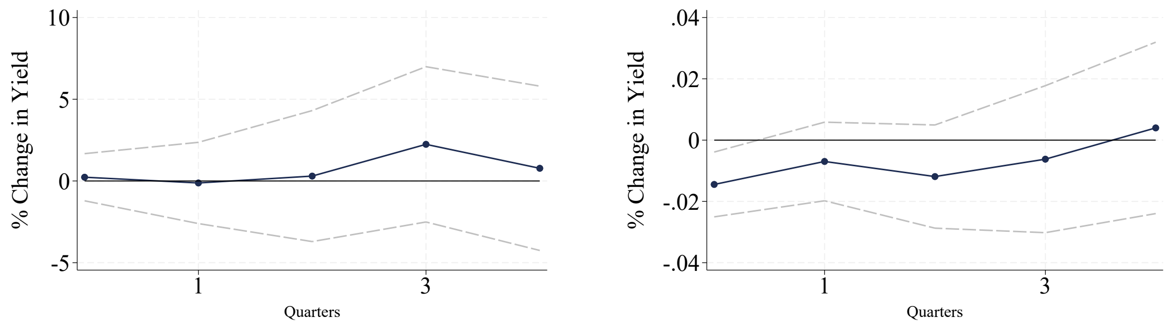
Colombia is an ideal case study due to its significant post-inclusion index weights and consistent domestic current account policies over the past two decades. Despite having foreign capital gains withholding taxes, Colombia did not impose explicit capital flow controls. The proportion of LC bonds held externally by global funds exceeded 20% even before 2014. The next big index inclusion event for Colombia was its inclusion in the Bloomberg Global Aggregate Index in September 2020. Therefore, I focus on global funds benchmarked to JP Morgan GBI-EM indices and analyze the period between 2009 and 2019 to avoid the impact of two global crises and the second major index inclusion event.

Unlike the baseline, I use the face value of holdings to construct $I_{fund,it}$ and θ_{it}, as in Figure 3 panels e) and f). Figure 4 graphs the cumulative impact of creditor composition on yields over a year for the pre-inclusion period 2009q1 $\leq \{t + h|h, t\} <$ 2014q1 and the post-inclusion period 2014q1 $\leq \{t+h|h, t\} <$ 2020q1. The point estimates change from zero to negative between the two periods. Existing theories based on financial frictions cannot explain why the same types of funds would not affect the pass through of the shock before Colombia's index inclusion but dampen the shock thereafter.

4.4 Further empirical supports from flow-induced demand shocks

I further address the concerns that the global risk factor approximated using VIX is not a well identified shock and hence the lack of causal claims. I construct bond-level demand shocks by

Figure 4: Falsification test using pre-inclusion periods



(a) Pre-inclusion effect (2009 to 2013)

(b) Post-inclusion effect (2014 to 2019)

Notes: This graph come from estimating the baseline specification for Colombia. To construct creditor composition measure, I use a sub sample of EM funds whose benchmarks include JPM GBI-EM family indices. The graph presents a 90% confidence band from robust standard errors.

exploiting idiosyncratic funding outflow shocks. Capital outflows from mutual funds reduce the size of funds and demand for EM sovereign bond. This part follows Zhou (2023), which uses funding flows to construct an instrument for bond prices. Similarly, Sander (2019) uses changes in mutual funds' funding to construct an instrument for capital flows to emerging economies. The demand shock z_{it} to bond i due to the funding outflows by global mutual funds depends on a share term share$_{ift}$ and a shifter term $\tilde{g}_{ft}$. share$_{ift}$ is the lag value of a bond in fund f's portfolio relative to the total market value of this bond held by global mutual funds. This measure represents the exposure of the mutual fund to this bond. The shift term attempts to capture the funding (i.e. a mutual fund's assets under management) outflow due to fund specific conditions unrelated to broad macroeconomic conditions. Denote g_{ft} as the reduction in the funding normalized by the lag size of the mutual fund $g_{ft} \equiv -\frac{D_{f,t}-D_{f,t-1}}{D_{f,t-1}}$, where $D_{f,t}$ is the asset under management.

$$z_{i,t} = \sum_{f \in \mathcal{J}(i)_t} \text{share}_{ift} \times \widehat{\tilde{g}}_{ft} = \sum_{f \in \mathcal{J}(i)_t} \left(\frac{w_{if,t-1} D_{f,t-1}}{\sum_f w_{if,t-1} D_{f,t-1}} \right) \times \widehat{\tilde{g}}_{ft}$$

Since mutual fund returns directly affect outflows, I extract outflows unrelated to a fund's return. $\widehat{\tilde{g}}_{ft}$ is the residual of funding flows obtained by estimating

$$g_{ft} = \alpha_t + \sum_{l=-1}^{L=2} \beta_l r_{f,t-l} + \tilde{g}_{ft} \tag{13}$$

The construction of the demand shock is similar to a shift-share instrument. The assumption is that the funding shock $\widehat{g}_{ft}$ is as good as randomly assigned to each fund, such that each mutual fund is expected to face identical funding shocks regardless of the unobservable factors affecting a fund's choice of the portfolio of bond i (Borusyak et al., 2022). To examine how EM funds affect the transmission of demand shocks, I estimate

$$\begin{aligned}\log(1+y_{it+h})-\log(1+y_{it-1}) &= \beta^h z_{it} + \delta^h \theta_{i,t-1} + \gamma^h z_{it} \times \theta_{i,t-1} \\ &\alpha_i^h + \alpha_{c(i)t}^h + \Gamma_1^h \mathbf{X}_{it} + \Gamma_2^h z_{it} \mathbf{X}_{it} + \epsilon_{i,t+h}, \quad h = 0, 1, ..., 8\end{aligned} \tag{14}$$

where z_{it} is the outflow-induced demand shock defined above, and $\alpha_{c(i)t}^h$ is the country-time fixed effect. Controls in $\mathbf{X}_{it}$ include the change in yield prior to the shock $(\log(1+y_{it-1})-\log(1+y_{it-2}))$, bond i's residual maturities, and the log GDP of the issuing country. Figure 5 shows that outflow-induced demand shock raise yields, but the negative interaction effect indicates that the more a bond is held by EM funds, the less impact of demand shock has on bond yield.

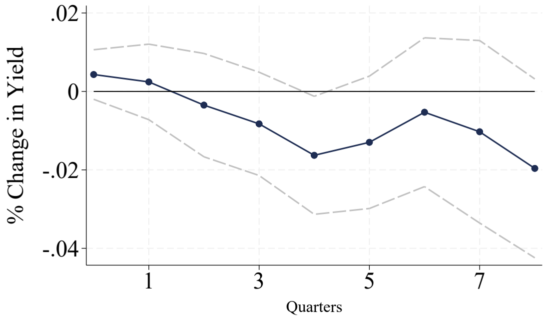
5 A sovereign debt model with bond benchmarks

I use a model-based approach to address two questions difficult to answer empirically. Previous results show that benchmarking reduces the impact of the global risk factor VIX on local bond markets. Changes in VIX imply a combination of shocks. I examine the stabilizing role of benchmarking under different shocks. Moreover, the supply-demand framework of sovereign bonds implies that the quantity of bonds outstanding in the secondary market affects the transmission of shocks. I endogenize the quantity of bonds in the secondary market by introducing a primary bond market in an Eaton-Gersovitz style model of sovereign default. The counterfactual experiment allows me to disentangle the direct effect due to benchmarking from the equilibrium effect of changing the quantity of debt carried into the secondary market.

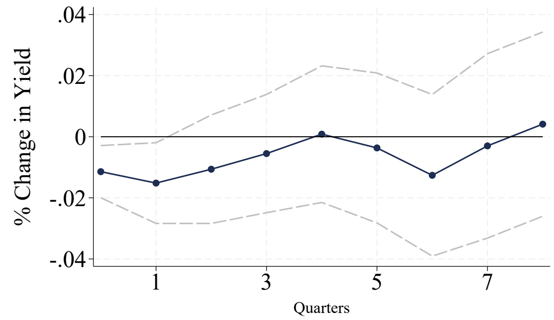
In a preliminary calibration, I demonstrate that the model can generate moments consistent with the observed reduction in yields following a country's index inclusion. However, the current model includes only a single bond issued by a small open economy. Consequently, replicating the empirical evidence using simulated data and employing micro-level empirical moments to calibrate the model's parameters are tasks reserved for future extensions. Specifically, the extended model will incorporate multiple local-currency bonds with idiosyncratic risk.

5.1 Model environment

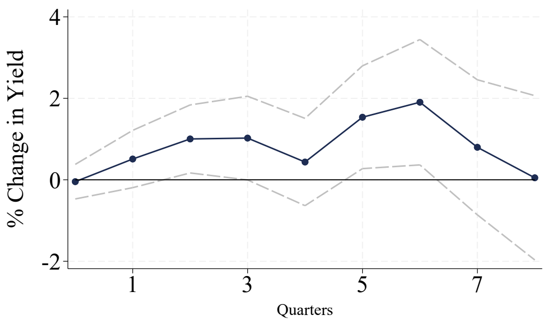
Figure 5: The transmission of outflow-induced demand shocks



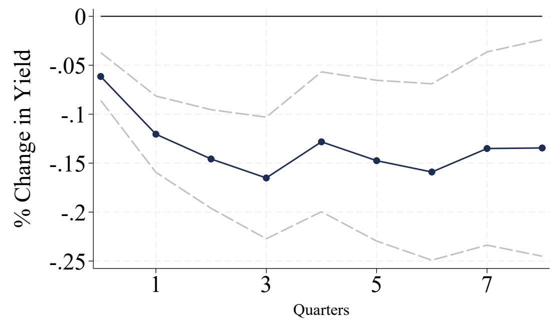
(a) Outflow-induced demand shock (Fixed effect:$\alpha^h_{c(i)t}$)



(b) Benchmarking effect (Fixed effect: $\alpha^h_{c(i)t}$)



(c) Outflow-induced demand shock (Fixed effect: α^h_t, $\alpha^h_{c(i)}$)



(d) Benchmarking effect (Fixed effect: α^h_t, $\alpha^h_{c(i)}$)

Notes: This graph shows the effects of outflows on bond yields. Panel (a) shows that outflows reduce the demand for sovereign bonds and increase yields. However, the negative interaction effect indicates that outflows to EM funds have a smaller impact compared to outflows to dollar funds. The graph shows a 90% confidence band from robust standard errors.

5.1.1 Primary bond market

The emerging market is modeled as a single representative country. Identical households with the following preference populate the country.

$$\mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t_{EM} u(c_t) \tag{15}$$

where $u(.)$ is increasing and concave, c_t denotes consumption of the representative EM household in period t, and $\beta_{EM} \in (0,1)$ is the subjective discount factor of EM households.

Each period, EM households receive an endowment of tradable goods Y_t. After the realization of exchange rate s_t and the current index status h, households choose whether to repay their debt or to default, where default is indicated by $d_t = 1$. Bonds are interpreted as local-currency bonds auctioned in the domestic market.

Upon default, households are unable to raise additional funds from private foreign investors, and consumption becomes $c_t = Y_t - \mathcal{L}(Y_t)$, where $\mathcal{L}(.)$ represents losses associated with default. I assume an exogenous capital account policy such that, following a default, foreign capital re-enters with a probability of ϕ_{market}.

If the country chooses not to default, it borrows local-currency index eligible bonds, denoted B_{t+1} which are due in the next period $t + 1$.

Once the country has an open capital account, it chooses whether to join global bond indices ($k = 1$) or to refrain from satisfying index inclusion criteria by choosing the policy $k = 0$. At the end of the period, the index provider has a probability ϕ_{index} of including a country committed to inclusion ($k = 1$) when its initial index status is $h = 0$. Similarly, the index provider excludes this country from the index if $k = 0$, even if the initial index status is $h = 1$.

5.1.2 Secondary market

Denote $\mathbf{S} = (Y, s)$ as the tuple representing the current state of output and exchange rates. A secondary market consisting of three stages opens for outstanding bonds before the next period. Stage 0 inherits the states $(\mathbf{S}, B')$ and has outcomes identical to the primary market. In Stage 1, new endowment and exchange rate are observed. Asset managers adjust their portfolios based on the default probability revealed by the new state. The price $q_{sm,1}$ clears the market. At this point, if a country has exclusion status and opts for index inclusion, then Stage 2 begins. The index provider announces the country's inclusion with probability ϕ_{index} and updates h' to 1. Otherwise, h' remains at 0. After Stage 2, the secondary market ends, and agents observe a new state in the next period. Table 4 summarizes the sequence of events.

Global asset managers enter EM local bond markets and can purchase any bonds issued or traded there. Their problem is identical to that described in Section 3.1.

5.2 Discussion of assumptions

"Defaulting" domestically issued local-currency bonds. Unlike outright default on hard-currency denominated bonds, countries can suspend, delay, or defer coupon or principal payments on domestically issued bonds held by foreign creditors. Sovereigns may implement these policies through capital outflow or currency exchange controls. In this context, "defaulting" refers to any of these actions, which can be potentially costly for mutual funds[9]. For example, in 2014, JP Morgan justified Nigeria's exclusion from the index due to its currency exchange controls and reduced

[9] Beers et al. (2023) reports that domestic LC debt in arrears accounted for 2.5% of the total arrears in 2022.

Table 4: Timing of events

Current period				Next period
Primary market	Secondary market			Primary market
	Stage 0	Stage 1	Stage 2 (if $h=0, k=1$)	
−Observe states $(\mathbf{S}, h)$ −Inherit state B −Decide to default or not −If not, decide k and borrow B' − Market clearing at q	− Identical to the primary market	−Observe state S_{sm} −Inherit states B' −Market clearing at q_{sm1}	−Announce h' $Pr(h'=1) = \phi_{\text{index}}$ −Market clearing at q_{sm2}	−Observe states $(\mathbf{S}', h')$ −Inherit state B'

liquidity in its currency markets (Mu'azu, 2015). Iceland forced foreign investors to reinvest capital gains in local bonds instead of allowing the funds to leave the country (Bianchi and Lorenzoni, 2021; Central Bank of Iceland, 2021).

Moreover, accumulating interest and principal arrears or deferring coupon payments on local-currency denominated bonds was common during the COVID-19 pandemic. U.S. mutual funds report whether or not a bond is in default in their Form N-PORT-P filings[10]. Appendix Table D.5 summarizes reported cases of defaults or deferred interest/principal payments, most of which have an average market value of about \$1 million. Overall, empirical evidence supports modeling LC defaults similarly to defaults on hard-currency denominated debt.

Optimal policies related to index presence. I argue that sovereigns have policy instruments to influence a country's eligibility for index inclusion and the degree of its presence. Capital controls influence a country's eligibility, and the characteristics of bonds influence a country's index weights. Appendix Section B.2 provides narrative and quantitative evidence of issuing index-eligible bonds to support this argument.

Omitting domestic index ineligible bonds, foreign-currency bonds, and long-term bonds. If the EM country chooses the optimal portfolio of index-eligible and ineligible bonds, then index-eligible and ineligible bonds issued by a country included in the index should have an identical price. However, these two bonds are still imperfect substitutes because they face different supply elasticities[11]. Therefore, depending on the index weight, the channel in the model gives rise to en-

[10]The form asks, "Currently in default?" and "Are there any interest payments in arrears or have any coupon payments been legally deferred by the issuer?"

[11]The coexistence of two bonds with identical prices in a standard Eaton-Gersovitz class of model requires adding bond-specific default costs. Dellas and Niepelt (2016) analyzes a country's portfolio of private and official loans, which have identical prices because a country cannot selectively default on one type of creditor. They argue that additional default costs of official creditors explain the coexistence of both debts.

dogenous portfolios of bonds with different index status. In fact, I show in Appendix Section B.2.3 that for countries persistently in the index, the issuance of index-eligible bonds relative to the total domestic debt is procyclical: countries tilt away from index-eligible bonds during global downturns. Depending on benchmark weights, the model could generate this result qualitatively.

Existing studies have extensively analyzed dollar-denominated bonds underwritten in developed markets and long-term bonds. Given that these studies focus on the trade-off between defaulting and repaying, they do not directly interact with the operative mechanism of sovereign benchmarks explored here. Nonetheless, long-term bonds could introduce new mechanisms with bond benchmarks: countries may swap index ineligible bonds with index eligible ones when the index status changes. Analyzing long-term bonds is left for future research.

Further discussion. The model is designed to examine the transmission of shocks in the secondary market. However, it is silent on how realized shocks and capital inflows thereafter in the secondary market affect future debt policies. Given the nature of forward-looking prices, anticipated changes in future debt policies would be reflected in prices. The answer to the question in this paper may be different when considering changes in future debt policies.

Broner et al. (2021) suggest capital flows related to index inclusion and alternating weights affect exchange rates and could have distributional effects across sectors. Although the aggregate effect remains an open question, the model in this paper shuts off any feedback channels.

How secondary market prices affect *future* debt policies and borrowing cost in the primary market[12] is still an open research agenda. The consensus view is that prices in primary and secondary markets are linked. The detailed transmission channel across these two markets remains elusive. Zhou (2023) analyzes how prices in primary and secondary markets are linked for 41 cases where new issues have identical characteristics between the two markets in the appendix of his paper. In fact, he shows the price in the primary market is lower than in the secondary market. The most widely recognized and analyzed linkage is through liquidity in the secondary market (Chaumont, 2020).

5.3 Recursive equilibrium

I characterize a Markov perfect equilibrium. The state space is summarized by the tuple $(\mathbf{S}, h, B)$. $\mathbf{S}$ is the tuple of current exogenous states (Y, s), where Y is the endowment and s is the exogenous

[12] Analyzing primary market is not a common theme in the sovereign debt literature. Some papers examining the frictions in the primary markets are Cole et al. (2022) and Cole et al. (2024). An alternative view of the secondary market is that the functioning of the secondary market determine nature of the default. Broner and Ventura (2016) argue that creditors do not expect the country to default when foreign creditors are able to trade with domestic creditors in the secondary market.

exchange rate. h is the country's index inclusion status, taking values 1 or 0. B is the bond issued in the previous period and due in the current period.

If the government has an open capital account, it chooses the policy d, deciding whether to default ($d = 1$) or not ($d = 0$).

$$V(\mathbf{S}, h, B) = \max_{d \in \{0,1\}} \left\{ V^r(\mathbf{S}, h, B), V^d(\mathbf{S}) \right\} \tag{16}$$

When the government repays, it chooses the optimal level of debt B' and consumption. If the country is currently in the index ($h = 1$), then it can choose to be excluded $k = 0$. The state of the index status in the primary market and in stages 0 and 1 depend on $h \times k$. It means that the index status is inclusion when index providers approve (i.e. inheriting an index status $h = 1$) and the actual optimal policy is inclusion ($k = 1$).

$$\begin{aligned} V^r(\mathbf{S}, h, B) = &\max_{c, B', k} u(c) + \mathbb{E}_{\mathbf{S}_{sm}|\mathbf{S}} V^{sm}(\mathbf{S}_{sm}, h \times k, B', k) \\ &\text{s.t. } c = Y + q(\mathbf{S}, h \times k, B', k)B' - B \end{aligned} \tag{17}$$

During the secondary market, the value function and the index status are updated according to exogenous processes and the updating rule:

$$\begin{aligned} V^{sm}(\mathbf{S}_{sm}, h \times k, B', k) = &\beta_{EM} \mathbb{E}_{\mathbf{S}'|\mathbf{S}_{sm}} (V(\mathbf{S}', h', B')) \\ &\text{s.t. } h' = \begin{cases} 0, & \text{if } h = 0, k = 0 \\ 1, & \text{if } h = 1, k = 1 \\ 0, & \text{if } h = 1, k = 0 \end{cases} \\ &\Pr(h' = 1) = \phi_{\text{index}}, \text{ if } h = 0, k = 1 \end{aligned} \tag{18}$$

If the country is currently excluded from the index ($h = 0$), then it remains excluded by choosing the optimal policy $k = 0$. If the government defaults, then

$$\begin{aligned} V^d(\mathbf{S}, 0, 0) = &u\big(Y - \mathcal{L}(Y)\big) \\ &+ \beta_{EM} \mathbb{E}_{\mathbf{S}'|\mathbf{S}} \big(\phi_{market} V(\mathbf{S}', 0, 0) + (1 - \phi_{market}) V^d(\mathbf{S}', 0)\big) \end{aligned} \tag{19}$$

The first order condition of the portfolio choice problem of asset manager type j during the primary market is

$$w_j(\mathbf{S}, h\times k, B', k) = (1-\alpha) w_j^b(\mathbf{S}, h\times k, B', q) + \frac{\mathbb{E}_{\mathbf{S}',h'|\mathbf{S},h,k}(\Delta R')}{(1+r^f) D_0 \gamma_j \operatorname{Var}_{\mathbf{S}',h'|\mathbf{S},h,k}(\Delta R')}, \quad j = \{\text{EM fund}, \text{Dollar fund}\} \tag{20}$$

where the expressions of depreciation adjusted for default and the excess return are:

$$\mathcal{E}' = \frac{s}{s'} \times (1 - d'(\mathbf{S}', h', B')), \quad \Delta R' = \frac{\mathcal{E}'}{q(\mathbf{S}, h\times k, B', k)(1+r^f)} - 1$$

The weighting function accounting for a country's index status is:

$$w_j^b = \begin{cases} \Omega(s^{-1} q B') & \text{if } j = \text{EM Fund}, h\times k = 1 \\ 0 & \text{if } j = \text{Dollar Fund, or } j = \text{EM Fund}, h\times k = 0 \end{cases}$$

The secondary market starts with Stage 0, where the clearing price is identical to that in the previous primary market. In Stage 1, once asset managers observe the new state $\mathbf{S}_{sm} = (Y_{sm}, s_{sm})$, the portfolio choice is

$$w_{j,sm1}(\mathbf{S}_{sm}, h\times k, B', k) = (1-\alpha) w_{j,sm1}^b(S_{sm}, h\times k, B', q_{sm1}) + \frac{\mathbb{E}_{S',h'|S_{sm},h,k}(\Delta R'_{sm})}{(1+r^f) D_0 \gamma_j \operatorname{Var}_{S',h'|S_{sm},h,k}(\Delta R'_{sm})}, \quad j = \{\text{EM fund}, \text{Dollar fund}\} \tag{21}$$

where the expressions of depreciation adjusted for default in the secondary market and the excess return in the secondary market are

$$\mathcal{E}'_{sm} = \frac{s_{sm}}{s'} \times (1 - d'(\mathbf{S}', h', B')), \quad \Delta R'_{sm} = \frac{\mathcal{E}'_{sm}}{q(\mathbf{S}_{sm}, h\times k, B', k)(1+r^f)} - 1$$

When an index inclusion announcement occurs in the Stage 2 in the secondary market, the portfolio choice is

$$w_{j,sm2}(\mathbf{S}_{sm}, 1, B') = (1-\alpha) w_{j,sm2}^b(\mathbf{S}_{sm}, 1, B', q_{sm2}) + \frac{\mathbb{E}_{\mathbf{S}'|\mathbf{S}_{sm}}\left(\frac{\frac{s_{sm}}{s'}\times(1-d'(\mathbf{S}',1,B'))}{q_{sm2}(\mathbf{S}_{sm},1,B')(1+r^f)} - 1\right)}{(1+r^f) D_0 \gamma_j \operatorname{Var}_{\mathbf{S}'|\mathbf{S}_{sm}}\left(\frac{\frac{s_{sm}}{s'}\times(1-d'(\mathbf{S}',1,B'))}{q_{sm2}(\mathbf{S}_{sm},1,B')(1+r^f)} - 1\right)}, \quad j = \{\text{EM fund}, \text{Dollar fund}\} \tag{22}$$

The asset market clearing in the primary market is

$$D_0 \sum_{j \in J} \mathcal{M}_j w_j(\mathbf{S}, h \times k, B', k) = s^{-1} q(\mathbf{S}, h \times k, B', k) B' \tag{23}$$

When the new state $\mathbf{S}_{sm}$ is realized, the market clearing condition for Stage 1 is

$$D_0 \sum_{j \in J} \mathcal{M}_j w_{j,sm1}(\mathbf{S}_{sm}, h \times k, B', k) = s_{sm}^{-1} q_{sm1}(\mathbf{S}_{sm}, h \times k, B', k) B', \tag{24}$$

If the announcement of index inclusion occurs in the secondary market, then the market clearing condition for Stage 2 is

$$D_0 \sum_{j \in J} \mathcal{M}_j w_{j,sm2}(\mathbf{S}_{sm}, 1, B') = s_{sm}^{-1} q_{sm2}(\mathbf{S}_{sm}, 1, B') B', \tag{25}$$

A Recursive Markov Equilibrium consists of exogenous states $\{\mathbf{S}(Y, s), \mathbf{S}_{sm}(Y_{sm}, s_{sm}), h\}$, value functions $\{V(.), V^r(.), V^{sm}(.), V^d(.)\}$, policy functions $\{d(.), B'(.), k(.), \{w_j(.)\}_{j=\{\text{EM fund,Dollar fund}\}}\}$ in the primary market and bond pricing function $q(.)$ in the primary market such that:

1. Taking as given the bond price function $q(.)$, the government's policy functions $B'(.)$ and $k(.)$, and choices of default $d(.)$ solve the optimization problem in equations (16), (17), and (19). $V(.), V^r(.)$, and $V^d(.)$ are the associated value functions. Value functions $V^{sm}(.)$ and $V(.)$ are consistent with equation (18).

2. Given $d(.)$ and $B'(.)$, the bond price $q(.)$, portfolio choices in the primary market $\{w_j(.)\}$ solve equation (20).

3. The primary market clears as in equation (23).

Additional policy functions of the portfolio choices are denoted $\{w_{j,sm\nu}(.)\}_{,j=\{\text{EM fund,Dollar fund}\},\nu=\{1,2\}}$ in the secondary market, and prices $\{q_{sm\nu}\}_{\nu=\{1,2\}}$ in the secondary market such that given $d(.)$ and exogenous states, the secondary market clears as in equations (24) and (25).

5.4 Calibration

I calibrate the model at a quarterly frequency in two steps. First, I fix a subset of parameters to standard values from the literature or based on the business cycle statistics of Colombia. I estimate the remaining parameters using macro moments, changes in the yields during Colombia's index inclusion.

I assume that EM households' utility function is $u(c) = \frac{c^{1-\sigma}}{1-\sigma}$. I set the inter-temporal elasticity of substitution σ at 2, as is standard in the literature. I set the quarterly risk-free rate at $r^f = 0.01$, which is consistent with the average real risk-free rate observed in the U.S. For the risk aversion parameter of dollar fund managers, I fix $\gamma_H = 21$, following the value used by Kekre and Lenel (2024) for the risk-bearing capacity of U.S. households. I pick the risk averse $\gamma_L = 14$ for managers of EM funds. I choose the initial wealth of mutual funds $D_0 = 1.5$ which is aligning with the existing sovereign debt literature using risk averse creditors (Lee, 2022; Aguiar et al., 2016).[13]

There is no evidence so far on the average duration of exclusion from the domestic market. Therefore, I conservatively set this duration at 4 years, based on empirical studies (Richmond and Dias, 2009; Gaston Gelos et al., 2004) focused on dollar debts. This corresponds to a re-entry probability of 0.0613. I also assume a conservative value for the exclusion from the index. According to Fick (2015), a country is not eligible for reentry into JP Morgan indices for one year after its exclusion. In practice, Egypt has faced an exclusion spell of 8 years, and Nigeria has been excluded since 2015. I set the index inclusion probability equal to the market re-entry probability.

I assume a multivariate AR(1) process for output and the (real) exchange rate of Colombia[14]. I log-linearly detrend quarterly real GDP and exchange rate values from Q1 2000 to Q3 2023. I assume that exogenous states follow the same processes in the primary and secondary markets. The transition from the secondary market at $t-1$ to the primary market at t and the transition from primary market to the secondary market at t are:

$$\begin{bmatrix} ln(Y_t) \\ ln(s_t) \end{bmatrix} = A \begin{bmatrix} ln(Y_{t-1,sm}) \\ ln(s_{t-1,sm}) \end{bmatrix} + \epsilon_t, \text{ and } \begin{bmatrix} ln(Y_{t,sm}) \\ ln(s_{t,sm}) \end{bmatrix} = A \begin{bmatrix} ln(Y_t) \\ ln(s_t) \end{bmatrix} + \epsilon_{t,sm}, \text{ where } \Sigma_{\epsilon_t} = \Sigma_{\epsilon_{t,sm}} \equiv \Sigma \tag{26}$$

The shock to the primary and secondary market both follow the same variance-covariance matrix Σ.

I adjust the variance-covariance matrix to ensure that it is consistent with the primary and the secondary markets each occupying half of a period. I first estimate a multivariate AR(1) using quarterly data from Colombia to obtain an estimate of the transition matrix A and a variance-covariance matrix $\hat{\Omega}$.

$$\hat{A} = \begin{bmatrix} 0.704, -0.047 \\ 0, 0.986 \end{bmatrix} \text{,and } \hat{\Omega} = 10^{-4} \times \begin{bmatrix} 5.33, -2.53 \\ -2.35, 27.13 \end{bmatrix}$$

[13] Lee (2022)'s model implicitly implies $D_0 = 1$. In Table 10, Aguiar et al. (2016) set the wealth of investors to be 2.5-2.7 times the average endowment.

[14] Colombia's quarterly real GDP is from its central bank. I obtain the nominal exchange rate from FRED and adjust it by CPI inflation from IFS.

I estimate the variance-covariance matrix Σ to ensure that:

$$\text{Var}\left(\begin{bmatrix} ln(Y_{t+1}) \\ ln(s_{t+1}) \end{bmatrix} | \begin{bmatrix} ln(Y_t) \\ ln(s_t) \end{bmatrix}\right) = \hat{A}\Sigma\hat{A}^T + \Sigma = \hat{\Omega} \Rightarrow \hat{\Sigma} \approx 10^{-4} \begin{bmatrix} 3.5, -1.0 \\ -1.0, 13.8 \end{bmatrix}$$

Output costs take the the form $L(y_{EM}) = \max\{0, a_1 y_{EM} + a_2 y_{EM}^2\}$ as in Chatterjee and Eyigungor (2012).

For the mass of EM funds, I use the full sample of funds downloaded from Morningstar. I identify the EM funds using my classification methodology. I estimate the proportion of their net assets relative to the net assets of all the funds in my sample. Specifically, I approximate each fund's net assets by dividing the market value of its holdings by the portfolio share. Next, I sum over two groups of funds each quarter. This approach is consistent with the procedure in Section 4, where I estimate the funding from global funds. I then take the average across time. The value $\mu = 0.63$ reflects the dollars funds, on average, have 63% funding in the mutual fund industry in the data. 35% of the managers' compensation comes from performance-based pay, corresponding to $\alpha = 1 - 0.35 = 0.65$. This value falls within the typical range of 20% to 40% for bonus pay in the asset management industry. Using micro-level data to discipline the value of α is left as a task in an extended model with multiple bonds.

I choose the following functional form for the benchmark weighting function $\Omega(d) = \frac{d}{d+\kappa}$. This functional form is motivated by the real-world practice of benchmark indices including multiple countries, where κ represents the outside assets in indices.

I select the remaining four parameters $\{\beta_{EM}, a_1, a_2, \kappa\}$ to match four moments, with three macro moments and one moment from the secondary market.

I use Colombia's index inclusion event (see Appendix B.2.2 for a discussion of the event and Colombian debt policies). I calculate the average changes in the yields of index-eligible bonds in a two-week window around the announcement of Colombia's bonds in the JP Morgan GBI-EM index. I define an elasticity term η as follows[15]:

$$\eta = \frac{\Delta y}{\Delta debt/GDP} \times \frac{1}{y} = \frac{-0.51\%}{\text{COL\$}\,86.75 \times 10^{12} / \text{COL\$}\,191 \times 10^{12}} \frac{1}{6.95\%} = -0.16 \tag{27}$$

I identify outstanding bonds qualified for inclusion Δdebt into the JP Morgan GBI-EM index. I choose parameters so that the simulation features index inclusion events and the model-implied elasticity aligns with Colombia's case. Colombia is not the only country with index inclusion

[15] I do not multiply the term debt/GDP in the expression as in the usual definitions of elasticity, since the model does not include index-ineligible bonds during a country's inclusion period. $\Delta debt$ presents the sum of principal values in COP and GDP is the annualized nominal GDP in COP for Q4 2014

Table 5: Parameter values of the baseline calibration

Panel A: Fixed Parameters			Panel B: Calibrated Parameters		
Param.	Description	Value	Param.	Description	Value
σ	Risk aversion—EM households	2.00	β_{EM}	Discount rate of EMs	0.73
ϕ_{market}	Market re-entry probability	0.0613	a_1	Default cost—level	-0.30
ϕ_{index}	Index re-entry probability	0.0613	a_2	Default cost—curvature	0.325
D_0	Funding	1.5	κ	Weighting rules— other assets	5.0
r^f	Risk-free interest rate	0.01			
γ_H	Risk aversion—Dollar funds	21			
γ_L	Risk aversion— EM funds	14			
α	Compensation-absolute return	0.65			
μ	Mass of dollar funds	0.63			

events, and Appendix F.1 discusses why Colombia's case is better than other inclusion event used for calibrating the model.

I use macro data from Q1 2005 to Q1 2021 from Colombia to construct three moments. Two of these moments are the mean and volatility of the LC bond spreads. For the spread, I use the five-year LC interest and US interest rate complied by Du and Schreger (2016). I use external LC debt data from Onen et al. (2022) to calculate the average of externally held LC debt to GDP for Colombia. These three moments are commonly used in the literature to calibrate the subjective discount factor and the cost of default (Uribe and Schmitt-Grohé, 2017).

I approximate the equilibrium dynamics through value-function iteration over a discretized state space. For the exogenous states of output and exchange rate, I discretize using 10 points for log output and 10 points for log exchange rate, covering ±3 times the unconditional standard deviation of their estimated process. To construct the transition probability matrix, I apply the iterative procedure proposed by Schmitt-Grohe and Uribe (2009)[16]. Finally, the stock of external debt is discretized on a grid of 25 equally spaced points, starting at 0 and ending at 0.8.

Discussion. Table 5 lists all the fixed and calibrated parameters. The discount factor of the emerging market country $\beta_{EM} = 0.73$, which is very low but not uncommon in the sovereign debt literature. Table 6 compares five empirical moments used for calibration and their counterparts from simulation. The model matches data well except for the volatility of spread. Existing literature tweaks the emerging market discount factor and two parameters governing the cost of

[16] https://www.columbia.edu/~mu2166/tpm/tpm.pdf

default to adjust the first and the second moments of spreads. The goal of endogenously generating episodes of index inclusion poses addition challenges. For example, raising β_{EM} reduces the volatility of spread and lowers the probability of default, as agents assign higher values to future consumption and avoid costly defaults. Given fixed parameters and the functional form of weight, index inclusion is always beneficial. As a result, a country does not choose to be excluded from the index. Generating index inclusion episodes relies on episodes of default followed by the index thereafter in the simulation. A lower probability of default under a higher β_{EM} results in no observed default and index inclusion. Since the elasticity η is informative on the deep parameters κ related to benchmarking, the current calibration chooses to sacrifice the match of the volatility of the spread in favor of the index inclusion elasticity.

Moreover, I show the evidence that EM funds reduce the transmission of shocks. I create a time series of secondary market prices and bond ownership between EM and dollar funds. Let $\Delta \log(s_{sm,t})$ be defined as $log(s_{sm_1,t}) - log(s_{sm_0,t})$ and $\Delta \log(y_{sm,t})$ as $log(y_{sm_1,t}) - log(y_{sm_0,t})$. Using model-simulated EM fund share θ, I estimate the reduced-form equation:

$$\Delta \log(y_{sm,t}) = \alpha + \beta \Delta \log(s_{sm,t}) + \gamma \Delta \log(s_{sm,t}) \times \theta_{sm0,t-1} + \Gamma X_t + \epsilon_t \tag{28}$$

where the controls X_t include $\theta_{sm0,t}$ and the change in output, $log(Y_{sm1,t}) - log(Y_{sm0,t})$, in the secondary market. Under above calibrations, the sign of $\hat{\beta} = -42.6$ is positive and the sign of $\hat{\gamma} = 22.3$ is negative. While this regression does not replicate the panel regression with fixed effects presented in the empirical section, the finding that EM funds reduce the sensitivity of asset prices to shocks is consistent with the empirical results.

Table 6: Moments

Target	Description	Data	Model
$\mathbb{E}[SP]$	EM LC bond spreads— average	5.1%	5.2%
$\sigma(SP)$	EM LC bond spreads— volatility	2.0%	7.9%
$\mathbb{E}[B/Y]$	External LC debt to output	18.7%	18.4%
η	Index inclusion yield elasticity	-0.16	-0.27

Notes: The spread (SP) is the difference between annualized yields of Colombian and US Treasury bonds at a 5-year tenor. The data is from Du and Schreger (2016). η is defined in equation 27.

5.5 Mechanisms and their quantitative importance

Under the supply-demand framework in the model, the quantity of bonds affects the transmission of shocks in the secondary market. To see this, I extend the analysis of the impact of the risk factor z_0 on yield y_0 in Section 3.3.1 to the outstanding bond B:

$$\frac{\partial^2 \hat{y}_0}{\partial \hat{z}_0 \partial B} \propto \phi + (w_0 - (1-\mu)(1-\alpha)\epsilon^b)(1 - \frac{1}{(1+r^f)V_0})$$

where $w_0 - (1-\mu)(1-\alpha)\epsilon^b$, ϕ and V_0, as defined in Section 3.3.1, are positive. The quantity of bonds has ambiguous effects on the transmission, but when the exchange rate is very volatile (larger V_0), the relationship is likely to be positive. Benchmarking could affect the outstanding debt in the secondary market in equilibrium, and this motivates me to analyze mechanisms during the primary market.

I omit the secondary market and analyze model mechanisms in a two-period model, followed by their quantitative relevance based on previous calibration. At $t = 1$, the government inherits a legacy debt B_{01} and observes the state $\mathbf{S}_1$ and the index status h. It decides to repay the legacy debt or not. If it repays, it can borrow using a one-period bond B due at $t = 2$. After observing the new state $\mathbf{S}_2$ at $t = 2$, the government decides to repay or default. The value functions are $V_t(\mathbf{S}_t, h)$. I assume a log utility function $u(c) = ln(c)$.

5.5.1 Benchmarking changes prices and demand elasticity

I analyze when a country benefits from index inclusion. I conduct a hypothetical experiment. Suppose that a country with an open capital account is under index exclusion ($h = 0$). The direct effect of index inclusion is raising the price, that is $q(\mathbf{S}_1, h = 1) - q(\mathbf{S}_1, h = 0) > 0$[17]. Let B^* and q^* be the optimal borrowing and the bond price for a given state $\mathbf{S}_1$ under index exclusion. Welfare improvement requires the inequality below when prices and their derivatives are evaluated at B^*:

$$\begin{aligned}
&(Y_1 - B_{01})(q(\mathbf{S}_1, h=1) - q(\mathbf{S}_1, h=0)) \\
&+ (Y_1 - B_{01})B\left(\frac{\partial q(\mathbf{S}_1, h=1)}{\partial B}\bigg|_{q(\mathbf{S}_1, h=1)} - \frac{\partial q(\mathbf{S}_1, h=0)}{\partial B}\bigg|_{q(\mathbf{S}_1, h=0)}\right) \\
&+ B^2\left(q(\mathbf{S}_1, h=0)\frac{\partial q(\mathbf{S}_1, h=1)}{\partial B} - q(\mathbf{S}_1, h=1)\frac{\partial q(\mathbf{S}_1, h=0)}{\partial B}\right) > 0
\end{aligned}$$

Given that $q(\mathbf{S}_1, h = 1) - q(\mathbf{S}_1, h = 0) > 0$, the second line in the above inequality being greater than zero is a sufficient condition for a welfare improvement under index inclusion. The proposition below summarizes this.

[17] See equation 29 in Appendix C.1 for the derivation. Section 3.3 also discusses this effect.

Proposition 3 *A sufficient condition exists to ensure the following holds:*

$$V_1(\mathbf{S}_1, h = 1) \geq V_1(\mathbf{S}_1, h = 0), \textit{ and } B(\mathbf{S}_1, h = 1) > B(\mathbf{S}_1, h = 0)$$

Appendix C.4 shows the proof.

Under the assumption that rising debt levels depress bond prices, a country improves its welfare by joining the index because bond prices become less sensitive to increases in supply. The deeper reason is that benchmarking overcomes the private cost of asset managers. Underinvestment in risky assets occurs when managers are compensated only by their absolute returns. Awarding performance relative to benchmark returns reduces underinvestment and potentially improves the welfare of EM countries[18].

To further characterize this sufficient condition, I denote the index weight upon inclusion, if the government continues to borrow at the previous optimal B^* but the equilibrium price $q(\mathbf{S}_1, h = 1)$ under index inclusion, as $w^b \equiv w^b(s_1^{-1} q(\mathbf{S}_1, h = 1) B^*)$. The sufficient conditions are

$$1 > \frac{\partial w^b}{\partial (s_1^{-1} q B)} \frac{s_1^{-1} q(\mathbf{S}_1, h = 1) B^*}{w^b}$$
$$\frac{\partial \mathbb{E}_1(\frac{s_1(1-d_2)}{s_2})}{\partial B} \operatorname{Var}(\frac{s_1(1-d_2)}{s_2}) - \frac{\partial \operatorname{Var}(\frac{s_1(1-d_2)}{s_2})}{\partial B} \mathbb{E}_1(\frac{s_1(1-d_2)}{s_2}) > 0$$
$$\frac{s_1^{-1} q(\mathbf{S}_1, h = 1) B^*}{D_0} - (1-\mu)(1-\alpha) w^b > \frac{s_1^{-1} q(\mathbf{S}_1, h = 0) B^*}{D_0}$$

where $\tilde{\epsilon}(q, B) \equiv \frac{\partial w^b}{\partial (s_1^{-1} q B^*)} \times \frac{s_1^{-1} q(\mathbf{S}_1, h=1) B}{w^b}$ is the sensitivity of the index weight to changes in bond market values. One real world interpretation of the index weight is that it is a weighted average of the country's weight across all existing indices that include this country. I want to remind readers that $0 \leq \tilde{\varepsilon} < 1$ as long as indices use market-value weighting rules[19]. However, many indices deviate from simple market-value weighting. For example, the JP Morgan GBI-EM index imposes

[18] Credit ratings, another bond market institution, also operate through price-based disciplines (Yang, 2023). Unlike bond benchmarks, which support more borrowing, price-based disciplines often imply that a country borrows less.

[19] The proof is straightforward. The set of indices that include a country's LC sovereign bonds consists of N indices. Each index includes other assets, such as LC bonds from other EM countries, sovereign bonds from developed markets, or corporate bonds. b_{EM} denotes the market value of LC bonds from country c and $b_{o/t,n}$ denotes the market value of other assets in the index n. $w^b \equiv \sum a_n \frac{b_{EM}}{b_{EM} + b_{o/t,n}}$, where arbitrary weights a_n satisfy $\sum_{n=1}^{N} a_n = 1$. Then $\tilde{\epsilon} = (\sum a_n \frac{b_{o/t,n}}{(s^{-1} q^* B^* + b_{o/t,n})^2}) \times \frac{s^{-1} q^* B^*}{w^b}$. Given that real-world values $s^{-1} q^* B^* + b_{o/t,n} > 2 \ \forall n$, the following inequalities hold

$$\tilde{\epsilon} < (\sum a_n \cdot \frac{b_{o/t}^{max}}{(s^{-1} q^* B^* + b_{o/t}^{max})^2}) \times \frac{s^{-1} q^* B^*}{s^{-1} q^* B^* / (s^{-1} q^* B^* + b_{o/t}^{max})} < \frac{b_{o/t}^{max}}{s^{-1} q^* B^* + b_{o/t}^{max}} < 1$$

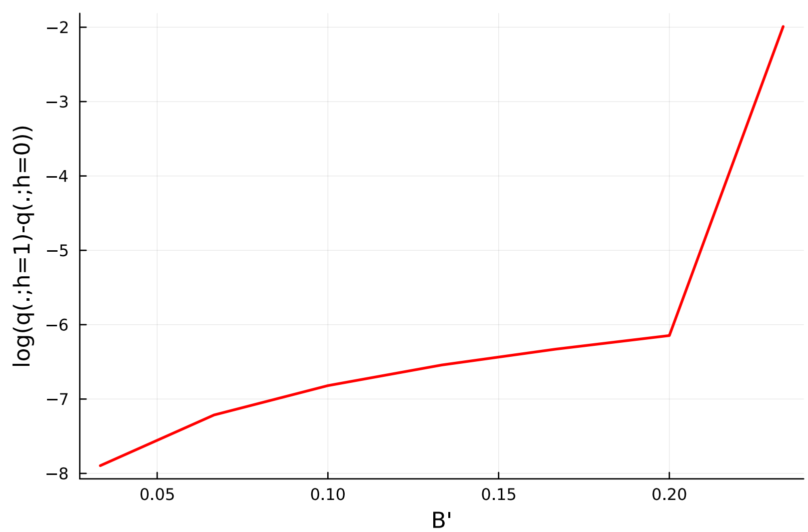
where $b_{o/t}^{max} \equiv \max\{b_{o/t,n} : n = 1 \ldots N\}$.

a ceiling on country weights to ensure diversification. A country's index inclusion may also come at the expense of lowering the weight of other countries.

In the second condition, borrowing increases the probability of default, so $\frac{\partial \mathbb{E}_1(\frac{s_1(1-d_2)}{s_2})}{\partial B} < 0$. Higher default probability implies that there are fewer states of the world where the country repays, and the variance of the exchange rate depreciation is also reduced, $\frac{\partial \operatorname{Var}(\frac{s_1(1-d_2)}{s_2})}{\partial B} < 0$. Therefore, the second condition says that changes of the variance of the exchange rate depreciation is more sensitive to changes of the expected value. In the third condition, $\frac{s_1^{-1} q(\mathbf{S}_1, h=1) B^*}{D_0} - (1-\mu)(1-\alpha) w^b$ can be interpreted as the "active portfolio". This condition says that the active portfolio weight rises under index inclusion.

Quantitative relevance. Figure 6 plots the difference in prices between index inclusion and exclusion for a "good" state (high y and low s). Higher prices under index inclusion align with the analytical results, but the difference is small. However, this difference increases as borrowing levels rise.

Figure 6: Difference in prices ($log(q(., h = 1) - q(., h = 0))$ under a "good" state



Notes: This graph plots the log difference in the pricing function with respect to future borrowing when the country has an output above average and an exchange rate below average (i.e., a "good" state for the country).

5.5.2 Benchmarking changes the probability of default

Let the set of states $\mathbb{S}$ be the combination of output and the exchange rate such that the country is indifferent between repaying and defaulting under index exclusion:

$$\mathbb{S} = \{\mathbf{S} : V_1(\mathbf{S}, B_{01}, h = 0) = V_1^d(\mathbf{S}, h = 0)\}$$

If the Proposition 3 holds for at least one element in the set $\mathbb{S}$, meaning that a country improves its welfare through index inclusion, then

$$\exists \mathbf{S} \in \mathbb{S} \text{ s.t. } V_1(\mathbf{S}, B_{01}, h = 1) > V_1^d(\mathbf{S}, h = 1)$$

In other words, the country does not default under index inclusion for states that would otherwise lead to default under index exclusion. Benchmarking reduces the set of states in which the country chooses to default.

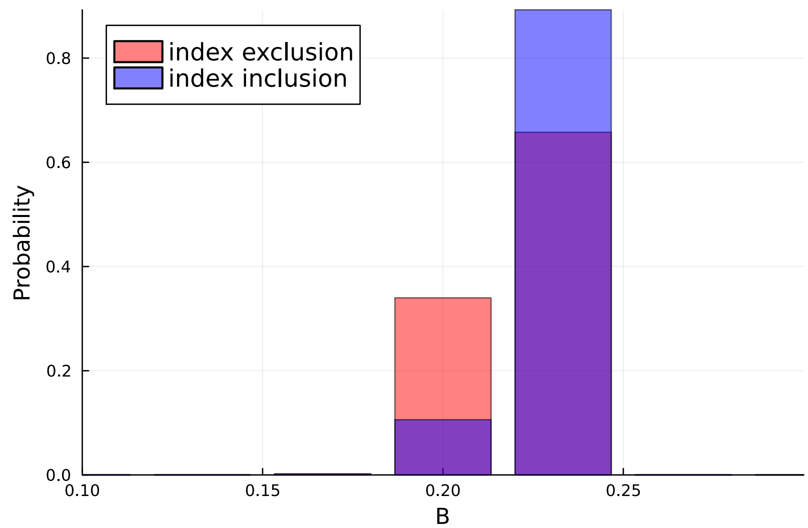
I provide a concrete example here. Suppose that value function under any index status is monotone decreasing in output for a given exchange rate $\tilde{s}_1$. Then, there is a unique cutoff output level $Y_1^{\min}$, below which the country defaults at $t = 1$ under index exclusion. When a country benefits from index inclusion under the state $(Y_1^{\min}, \tilde{s}_1)$, according to the proposition above, $V_1(Y_1^{\min}, \tilde{s}_1, B_{01}, h = 1) > V_1(Y_1, \tilde{s}_1, B_{01}, h = 0)$. The monotone relationship between the output and the value function implies that the cutoff output for default under $h = 1$ must be smaller than $Y_1^{\min}$. This relationship means that prior to the realization of Y_1, the country is expected to be less likely to default and could withstand a worse output shock.

Shrinking the set of states a country chooses to default increases the prices when creditors are risk-neutral. However, the higher expected return is offset by higher volatility of the return for risk-averse creditors here. As a result, the overall effect of lowered probability of default on prices is ambiguous.

The model of benchmarking offers a nuanced explanation for why countries are able to borrow despite their limited commitment to repaying: repaying debt is rewarded with potentially beneficial index inclusion, and the net cost of defaulting increases when this beneficial index inclusion is present.

Quantitative relevance. Conditional on a country choosing to default, Figure 7 shows the distribution of default probability across different levels of current debt for the country under index inclusion or exclusion. When the economy is under index inclusion, defaults are pushed to higher debt levels. In the current calibration, conditional on repayment, the value of index inclusion is higher in all possible states. Since the value when choosing to default is identical regardless of

Figure 7: Distribution of default probability



Notes: Conditional on a country choosing to default in the current period, this plot shows the distribution of the probability of default across three different levels of debt the country defaults on under two cases: the country is under index inclusion or exclusion status in the current period. To construct the plot, I simulate the economy over 10 million periods and calculate the distribution of debt for periods when the country chooses to default under different index statuses.

index status, the net cost of default is higher under index inclusion. Therefore, the country is more likely to choose to repay its existing debt.

5.6 Decomposing the transmission of shocks through benchmarking

I use the calibrated model to decompose the propagation of shocks into direct and equilibrium effects. Figures 8 and 9 plot the impulse response functions to one-standard-deviation shocks to the exchange rate and output. *Zero weight* refers to a counterfactual scenario where the country has a zero index weight while all else remains identical to the baseline case under index inclusion. The difference between this and the baseline case is the direct effect of benchmarking. *Index exclusion* refers to a counterfactual world where this country is excluded from the index. It is a counterfactual world without benchmarking for only one period in the primary market and the following secondary market, where the counterfactual experiment takes place. This case captures the equilibrium effect. It differs from *zero weight* by allowing the outstanding bonds in the secondary market to be optimally determined. When the outstanding bonds are identical under *index exclusion* and *zero weight*, the propagation of shocks is identical because they share the same pricing function.

Across all experiments, prices are lower under *zero weight*. This confirms the central message that benchmarking reduces the transmission of shocks to prices. However, benchmarking is more effective in insulating against output than exchange rate shocks, as exchange rate shocks directly reduce the benchmark weight. After 16 quarters, the price is about half of what it is prior to the shocks in the baseline, and benchmarking contributes to 8% of the total effect as the decline is 5 percentage points larger under *zero weight*. Figure 9 shows that benchmarking is effective at insulating against output shocks, as the price is 7.5% lower under the baseline compared with 15% under *zero weight*. Compared to exchange rate shocks, output shocks do not directly affect the index weight, and the stabilizing effect is preserved.

The experiment also reveals that the shock has a slightly smaller impact on prices from the second period onward in a world without benchmarking (*index exclusion*) relative to the baseline. The reason is that under *index exclusion*, the country reduces the amount of debt carried to the secondary market. The debt-to-output ratio is 18.7% in the baseline, whereas it is 4% lower under *index exclusion*.

Benchmarking provides the most stabilizing on the impact of shocks. For a large exchange rate shock, the secondary market of sovereign bonds unravels: the price approaches zero. At this point, the difference in price between the baseline and the two counterfactual scenarios is indistinguishable. For a output shock, benchmarking almost completely insulates the shock in the baseline, while the price falls by 12.5% in two counterfactual cases.

6 Conclusion

Benchmarks are market indices used to measure the performance of mutual funds. Many developing countries, such as China, Egypt, and India, tailor their market reforms to achieve inclusion in bond indices used as benchmarks. Capital inflows from global mutual funds have become an important driver of the internationalization of sovereign bond markets, but mutual funds, particularly ETFs and passive funds that closely follow indices, are often viewed as amplifying global shocks and increasing market volatility.

This paper shows that the index inclusion of emerging market local-currency bonds in global bond indices reduces the transmission of global shocks to their sovereign bond prices. I first propose a new binary classification of mutual funds based on whether or not their benchmark indices include local-currency bonds. I present a new stylized fact that mutual funds whose benchmark indices are unrelated to local-currency bonds actively invest in them. Using this new fact, my empirical strategy compares the shock transmission under different asset ownerships based on the new classification. I achieve this by assembling a bond-level dataset of sovereign bond holdings. I show that mutual funds whose benchmarks include local-currency bonds stabilize bond prices against shocks from global risk factors, as approximated by changes in the VIX. A sovereign default model with benchmarking provides additional insights into the impact of index inclusion under output and exchange rate shocks.

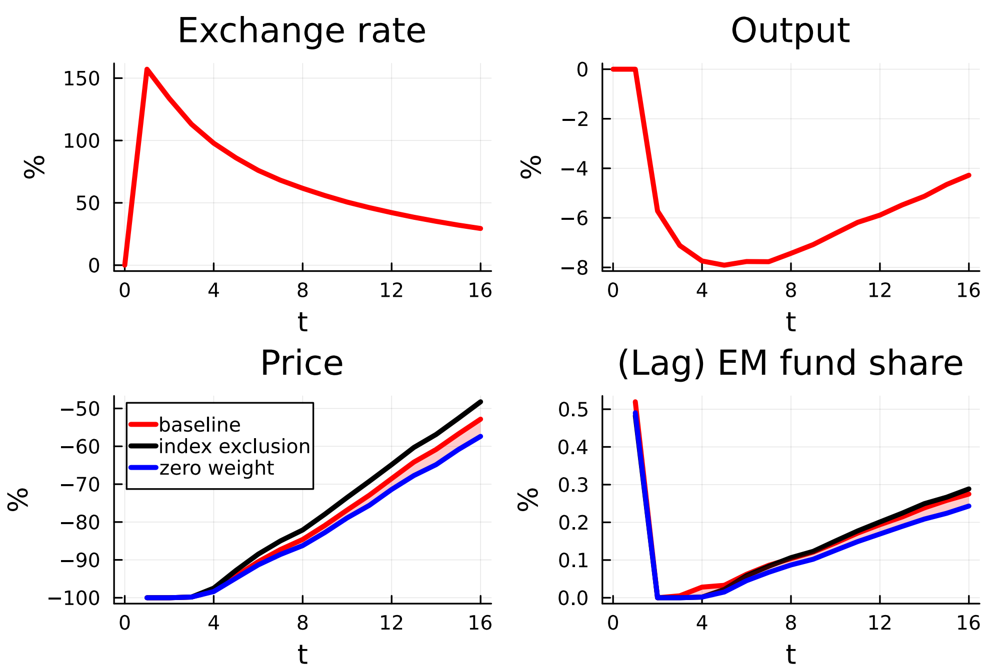
More significantly, if the creditors' heterogeneity in benchmarks directly affects asset prices, the determinants of benchmarks upon the entry of investment funds over the global financial cycle affect capital flows. Emerging economies may need to facilitate the entry of funds benchmarked to their assets or the creation of market indices that include their assets to further globalize their capital markets. Policies that influence the footprint in global indices become alternative forms of capital controls. For example, countries included in global bond indices have direct controls over the characteristics of bonds and hence the degree of presence in global indices, which ultimately influences capital account openness. In principle, the framework of benchmarking in this paper is qualitatively consistent with the pro-cyclical borrowing in bonds with characteristics that make them index eligible[20]. Future research could examine the imperfect substitution of two types of local-currency bonds—one with and one without index status—despite identical credit and currency risk.

More broadly, this paper examines the macroeconomic implications of financial frictions in asset management and their remedies through performance-based compensation using benchmarks. On the one hand, while models often invoke collateral constraints, as He and Krishnamurthy (2008)

[20] Appendix Section B.2 documents policies related to index presence. It also shows that countries are less likely to issue index eligible bonds during global downturns.

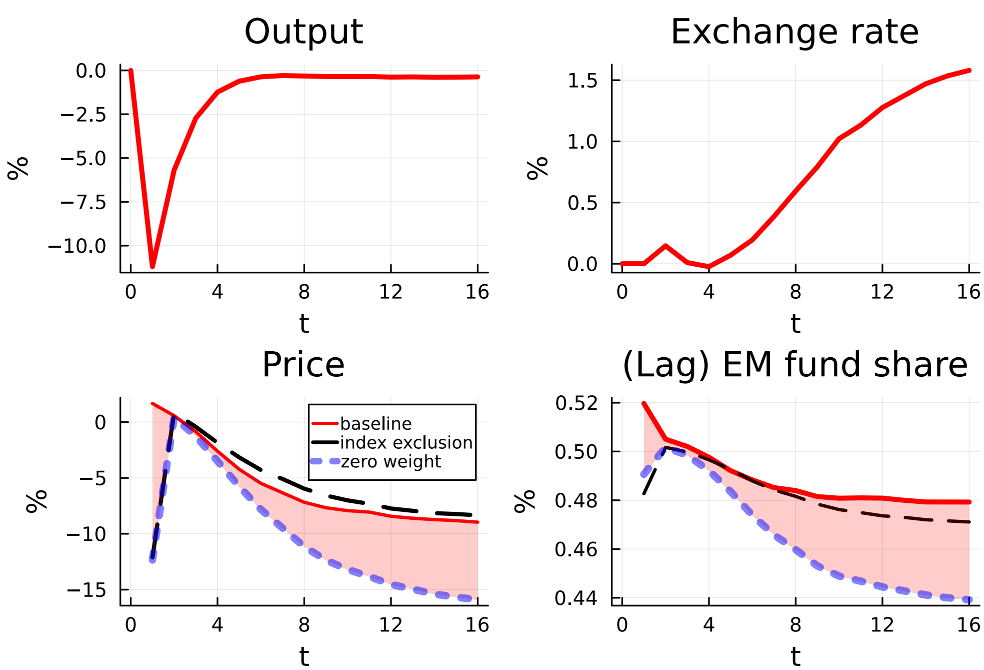
note, the micro-foundation of collateral constraints assumes the absence of benchmarks. Without this assumption, only a few works, such as Krishnamurthy (2003) and Di Tella (2017), explain financial amplification. On the other hand, benchmark compensation contracts—the micro-foundation behind the relative return in the utility function in this paper—are common in asset management. Assuming benchmark contracts does not preclude financial amplification due to the heterogeneity of benchmarks and deviations in investing in benchmark-related asset classes. The financial frictions in asset management may differ from those in other types of financial intermediaries, such as banks and hedge funds (He and Krishnamurthy, 2013), and may have broader macroeconomic implications for future exploration.

Figure 8: Decomposing of the transmission of exchange rate shocks



Notes: Impulse response functions to a one-standard-deviation real exchange rate shock in the secondary market for the economy under index inclusion, the baseline case, and two counterfactual cases. The unconditional standard deviation is 0.226. Before the shock, the stochastic real exchange rate and output follow their underlying Markov processes. Upon the EM country's entry into the secondary market, the exchange rate shock occurs while output remains constant. For the following 15 periods in the same secondary market, both the exchange rate and output follow their Markov processes. The response functions of the exchange rate, output, and the price of bonds in local currency are the deviations from the levels prior to the shock. The lagged EM fund share refers to the share of sovereign bonds' market value held by EM funds in the previous period. The graph plots the response of the bond price and the EM fund share starting from $t = 1$. The response functions consider two alternative scenarios. *Zero weight* refers to the case where the EM country has a zero index weight, while the debt level in the secondary market is identical to the baseline case. *Index exclusion* refers to a counterfactual case where the country has an index exclusion status throughout the primary and secondary markets. This scenario differs from *zero weight* by allowing the outstanding bonds to be optimally determined prior to the shock. Responses are the average of 10,000 simulations.

Figure 9: Decomposing of the transmission of output shocks



Notes: Impulse response functions to a one-standard-deviation output shock in the secondary market for the economy under index inclusion, the baseline case, and two counterfactual cases. The unconditional standard deviation is 0.0436. Before the shock, the stochastic real exchange rate and output follow their underlying Markov processes. When the EM country enters into the secondary market, the output shock occurs while the real exchange rate remains constant. For the next 15 periods in the same secondary market, both the exchange rate and output follow their Markov processes. See the notes in Figure 8 for other details.

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Appendix
Benchmarking and Sovereign Risk

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January 31, 2025

Table of Contents

A Data appendix

A.1 Bloomberg data

Identifying duplicate bonds. Commonly used bond identifiers, such as CUSIP and ISIN, do not uniquely identify bonds in Bloomberg bond data. I describe a procedure to identify duplicate bonds by considering STRIPS (Separate Trading of Registered Interest and Principal of Securities), global offerings under Regulation S or 144A, and GDN (Global Depository Notes). Bloomberg Terminal offers a built-in function to consolidate duplicate bonds resulting from STRIP or Regulation S/144A. I download two lists of sovereign bonds, restricting the variable "BCLASS II" to "Treasury" or "Sovereign", with and without consolidating duplicated bonds. The union of these two lists covers all CUSIP and ISIN identifiers that might be reported by mutual funds. Bloomberg also reports the status of bonds being "STRIP", "Reg S/144A", and "GDN". I retrieve a list of these types of bonds and include the variable "DES Notes", which provides additional descriptive information about a bond. From these notes, I obtain parent bond identifiers. I use ISIN to identify bonds whenever available, and use CUSIP when the corresponding ISIN cannot be found. Additionally, I refer to the GDN list from Citi Bank[21].

Bond yields. I download bond bid yield to maturity data from the Bloomberg Terminal. I select BVAL as the pricing source whenever available[22]. The pricing source is mainly BGN for bond yields before 2010.

A.2 Form N-PORT-P data

I create a dataset of U.S. mutual fund and ETF holdings of EM LC bonds using Form N-PORT-P from the U.S. Security Exchange Commission (SEC). The advantage of this dataset is that it is supposed to contain the universe of mutual funds in the U.S. Investment firms in the U.S. disclose fund-level portfolios to the SEC. SEC only releases filings to the public in the third month of the filers' fiscal quarters. The disadvantage of this dataset is that the sample only covers 12 quarters, from Q2 2019 to Q2 2022.

To create this data, I compile a list of share class IDs of U.S. mutual funds from Morningstar and identify corresponding ones in the SEC Investment Firm directory. Then, I link CIKs between the SEC Investment Firm directory and SEC Filing Index, which allows me to identify the ID numbers of filings to download. Appendix Table A.1 shows the identifiers in each dataset.

[21] See https://depositaryreceipts.citi.com/adr/guides/uig.aspx?pageId=8&subpageid=34. I also cross-check the GDN cases from Deutsche Bank (https://www.adr.db.com/drwebrebrand/dr-universe/dr_universe_type_e.html)

[22] See https://data.bloomberglp.com/professional/sites/10/Fixed-Income-Cash-Pricing-Sources.pdf for details of different pricing sources.

Table A.1: Identifiers in Morningstar and SEC data

Source	Morningstar	SEC Investment Firms	SEC Filing Index	Form N-PORT-P
Usage	Identify tickers of relevant funds and managing firms	Using tickers to get CIKs	Getting directory of Form N-PORT	Getting holdings
Parent Firm ID (CIK)		✓	✓	✓
SEC Fund Series ID		✓		✓
Share Class ID (Ticker)	✓	✓		

Note: SEC filing index: https://www.sec.gov/Archives/edgar/full-index/; SEC Investment firms: https://www.sec.gov/open/datasets-investment_company. A sample of filing is https://www.sec.gov/Archives/edgar/data/741350/000175272422215831/xslFormNPORT-P_X01/primary_doc.xml

I validate the Form N-PORT-P data against official data. The credit supply of $43.97 billion to EM local currency bond markets in Q2 2022 in the N-PORT-P data is comparable to $40 billion reported by mutual funds in the U.S. Portfolio Holdings of Foreign Securities as of December 2022[23]. Similarly, the credit supply of $65.66 billion in Q4 2020 and $65.95 billion in Q4 2021 from the Form N-PORT-P data is comparable to $69 billion in December 2020 and $67 billion in December 2021 from the official data.

A.3 Colombian sovereign bond issuance

I search for Colombian Peso denominated sovereign bonds from Bloomberg Terminal. The search criteria include asset classes categorized as "Government" (with consolidated duplicate bonds), BCLASS classification as "Treasury" and "Sovereign," maturity dates after January 1, 2005, and issuance dates prior to December 31, 2023. After removing duplicate bonds, such as those from Global Depository Notes (as described in Section A.1), and excluding bonds without information on the amount issued, this yields 309 individual bonds. The earliest bond in the sample was issued in 1997, and the latest maturity is in 2050. There is no coupon information available for floating-rate bonds, so their coupon payments are excluded from the analysis.

The discount rate used is the average annual short-term rate in Colombia from 2005 to 2023,

[23]See https://ticdata.treasury.gov/resource-center/data-chart-center/tic/Documents/shca2022_report.pdf

Table A.2: EM LC credit supplies (USD bn) in Q4, 2019 by US fund categories

Fund Category	Benchmarks Include EM LC bonds	Benchmarks Exclude EM LC bonds
Allocation	1.14	0.77
Alternative	0.01	0.26
Taxable Bond	46.95	33.67
Commodities		0.05
International Equity		0.62
Miscellaneous		0.01
Nontraditional Equity		0.00
Sector Equity		0.00
U.S. Equity		0.01

Notes: This table presents the market value (in USD billion) of EM local-currency bonds held by different categories of U.S. mutual funds, as reported in portfolios from Form N-PORT-P between December 2019 and February 2020. Fund categories are provided by Morningstar. The identification of EM local-currency sovereign bonds is based on reported currency denominations, issuer sectors (i.e., corporate versus treasury), and issuer countries. Benchmark information from Morningstar is also used. A fund is considered to include EM local-currency bonds if either its primary or secondary benchmark includes EM LC bonds.

sourced from the FRED series COLIR3TCD01STQ. For inflation-linked bonds, of which there are 57 in my sample, I use the Colombia Peso to UVR exchange rate published by the central bank[24]. I use the exchange rate on the last day of each year from 2005 to 2023 to calculate an average rate of exchange rate depreciation, and then use this rate to interpolate the level of exchange rates from 1997 up to 2050 to adjust the principal payment of inflation-linked bonds.

A.4 Morningstar holdings data

Once I have a list of funds through the two-step procedure described in Section 4.1, I then download their portfolio reporting dates from Morningstar Direct. At this stage, I further restrict the sample to funds with consistent portfolio reporting. I then use the API from Morningstar Direct to directly download holdings.

A large number of holdings lack CUSIP or ISIN identifiers. For the purposes of this study, I extract issuer names, bond coupons, and bond maturities from the security description field in the

[24] https://www.banrep.gov.co/es/estadisticas/unidad-valor-real-uvr

holding data. These three pieces of information are the minimum required for matching a bond identifier. However, out of millions of positions I have downloaded, approximately 10,000 do not have reported bond identifiers but contain enough information in the bond description to assign an identifier using an external database. Since this small number is unlikely to affect the results, I delete these positions without bond identifiers.

I then clean the fund holding data. First, I sum the weights of the bonds in each portfolio. I delete funds whose total reported weight in any quarter exceeds 100%. There is inconsistency and a lack of uniformity in how funds report the quantity of bonds in the "shares" field. Ideally, they should report the face value or the number of bonds in multiples of the par amount—the minimum face value of bonds auctioned in the primary market. I retrieve the exchange rates and calculate the implied market price of the bonds. Based on this price, I determine whether the quantity of the bond is reported in multiples of tens.

I illustrate the solution for cleaning reported face values using the SPDR Bloomberg Emerging Markets Local Bond UCITS ETF. I retrieve its portfolio on July 4th, 2024 from Morningstar Direct and validated using its public disclosure. It held the Brazilian bond—LETRA TESOURO NACIONAL 0 07/01/2026 (BRSTNCLTN848)— with reported shares of 89,300, a bond price of 0.8057, and a market value of $13,096,800.92. Here, the fund reported the share in units of 1,000. The adjusted share should be 89,300,000. I account for this possibility in the reported shares in the data.

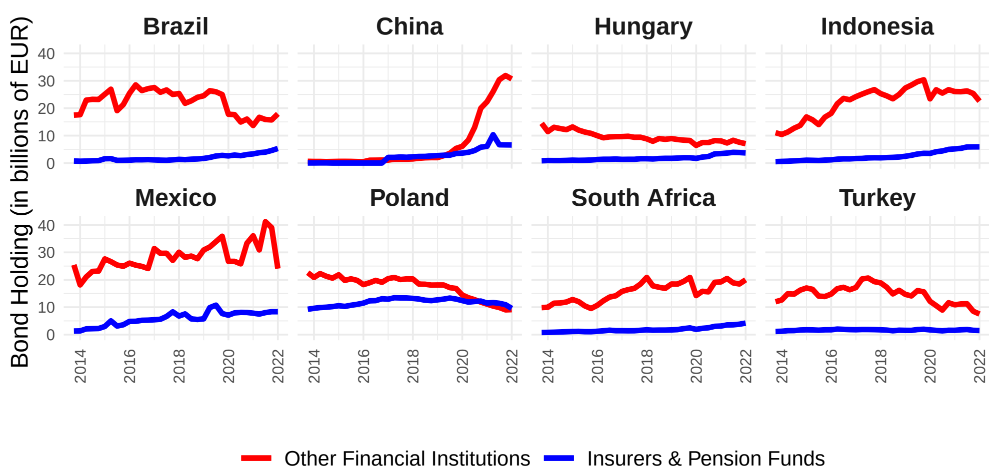
B Empirical result appendix

B.1 Investment funds and EM LC globalization

Many developing countries have globalized their domestic sovereign bond markets over the past two decades. Unlike hard-currency denominated debts, investment funds hold more local-currency denominated sovereign bonds than other types of non-official creditors who are crucial for their hard-currency borrowing, such as commercial banks, insurance companies, and pension funds.

To gauge the size of credit from investment funds, I use credit supply data by investor sectors from the ECB. I select non-developed countries which experienced local currency market development. As Figure B.1 shows, the investor sector "other financial institutions" which include European investment funds generally supply more credit than insurance and pension funds. Depository banks, extensively analyzed, supplied little credit to these countries. The visual inspection also reveals that the credit from "other financial institutions" are volatile. This fact is consistent with the discussion of flighty credit from investment funds.

Figure B.1: Credit supplied to emerging markets by investor sectors



Notes: Data comes from the European Security Holding Database. "Other financial intuitions" include mutual funds. It is referred to "Financial corporations other than money-issuing institutions, insurance corporations and pension funds" (sector code S12P) in the ECB data.

ECB data does not reveal details on finer divisions by currency and creditor types such as the credit supplied by hedge funds. For hedge funds, U.S. Treasury International Capital (TIC) System data indicates that the holdings by other financial firms which include hedge funds are $8 billion, $7 billion and $16 billion in 2020, 2021 and 2022 respectively, comparing to an average of $59 billion by mutual funds in these years.

B.2 Additional evidence on borrowers' policy decisions on benchmark index presence

I present the following facts:

1. Countries have policy tools to influence their index presence;
2. Countries without explicit capital controls and that are persistently included in global indices tilt their LC borrowing away from index-eligible bonds during global downturns.

The model in the main text assumes the country only issues index eligible bonds. That model is consistent with these facts qualitatively, and the quantitative importance depends on the specific weighting function.

B.2.1 Narrative evidence

China relaxed its capital controls specific to mutual funds in 2016 and subsequently gained index inclusion in the JP Morgan GBI-EM index (Clayton et al., 2022). Like China, India imposed discriminatory capital control policies across different types of creditors (Patnaik et al., 2013). Policies that removed capital controls and developed clearing technologies promoted India's inclusion in the JP Morgan GBI-EM index in 2023. Argentina and Nigeria are two cases where countries prioritized managing exchange rates and controlling capital flows. Facing economic downturns, these two countries restricted currency exchanges to stabilize the value of their currencies. Global index providers deemed this policy as reducing market liquidity and currency convertibility despite potential benefits. As a result, JP Morgan issued warnings but both countries responded by heightening currency exchange controls. Consequently, JP Morgan excluded Nigeria in 2013 and Argentina in 2019, just two years after the inclusion in 2017, from the GBI-EM Broad index.

Egypt is another recent example. It is a country with no explicit capital control to foreign investors. However, it imposes policies to borrow in index eligible bonds and fulfill index requirement to gain inclusion in JP Morgan's indices (Ministry of Finance Arab Republic of Egypt, 2022). Shortly after its 2022 inclusion, Egypt suffered an economic crisis in 2023, devalued its currency, and depleted its dollar reserves. JP Morgan excluded Egyptian domestic bonds in 2024. What complicates the Egyptian case are the divided views on index inclusion and exclusion from the perspective of emerging markets. Despite the expected small index weights, the finance minister tailored reforms to appeal to index providers and seek inclusion. The opposing view emerged when the exclusion occurs— arguing that Egypt's index weights are too small for any significant negative consequence from the exclusion (Abdel-Razek, 2024).

B.2.2 A case study of Colombia

I further use Colombia to illustrate how the characteristics of bonds may influence index presence. Colombian global LC bonds were already in the index prior to the full inclusion, representing about 0.3% in the JP Morgan GBI-EM Global Diversified index, JP Morgan rationalized the inclusion in domestically issued bonds as the "result of improved transparency and accessibility for international investors in the local TES market" (Broner et al., 2021). Colombia did not have any significant bond market or capital account policy changes prior to inclusion, but borrowing in index-eligible bonds may have been a precursor.

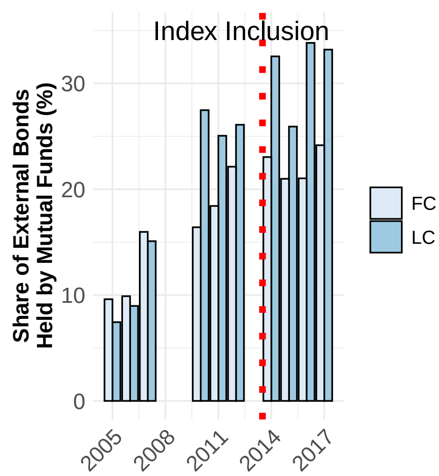
For each unique COP-denominated bond issued by the Colombian Treasury, I classify them as index-eligible or ineligible according to the criteria outlined in JP Morgan documentation. Given the different characteristics of bonds and varying maturities, I calculate the present value of expected interest and principal payments in Colombian Peso. This represents the debt level of the Colombian government. Historical coupon payment data is not available from commercial sources, so the estimates of the debt level omit interest payments for floating-rate and variable-rate bonds. The Colombian government experienced a significant increase in the debt level of eligible bonds prior to index inclusion. However, after inclusion, the government shifted away from issuing eligible bonds. I also decompose the debt level by other bond characteristics, such as coupon types, maturity, and issuance size, in Appendix A.3. Following the inclusion, the government borrowed more through short-term bonds.

Furthermore, I directly estimate the share of external local-currency bonds held by mutual funds using their aggregate positions up to 2017 from the Global Capital Allocation Project (Maggiori et al., 2027; Coppola et al., 2021). I construct the share of external local- and hard-currency sovereign bonds held by mutual funds. Figure B.2 shows that mutual fund shares reached 35% for local-currency denominated debt in 2017. More importantly, mutual fund shares of local-currency debt are consistently higher than those of dollar-denominated debt. Colombia was able to attract capital from global funds even with minimal index presence. This fact is consistent with the model.

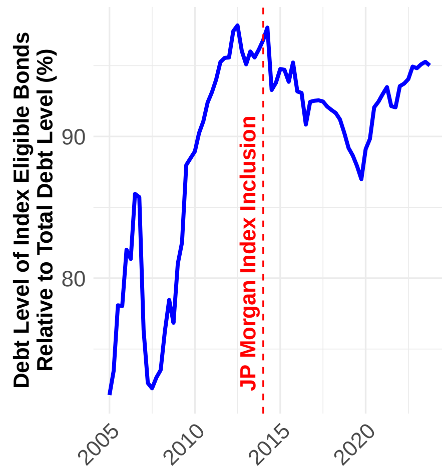
B.2.3 Issuing index eligible bonds over the global financial cycle

The case study of Colombia shows that it borrows in both index-eligible and ineligible bonds and that the debt level of index-eligible bonds exhibits cyclical patterns. I further explore the cyclicality of issuing index-eligible bonds. I collect data from Bloomberg on local-currency sovereign bonds issued by common emerging market countries from 2005 to 2023. Similar to the case study, I use the bond eligibility criteria of the JP Morgan GBI-EM Diversified index to identify index

Figure B.2: A case study of Colombia's index inclusion: creditor and debt instrument compositions



(a) Colombian external debts held by mutual funds



(b) Colombian debt level of index eligible bonds

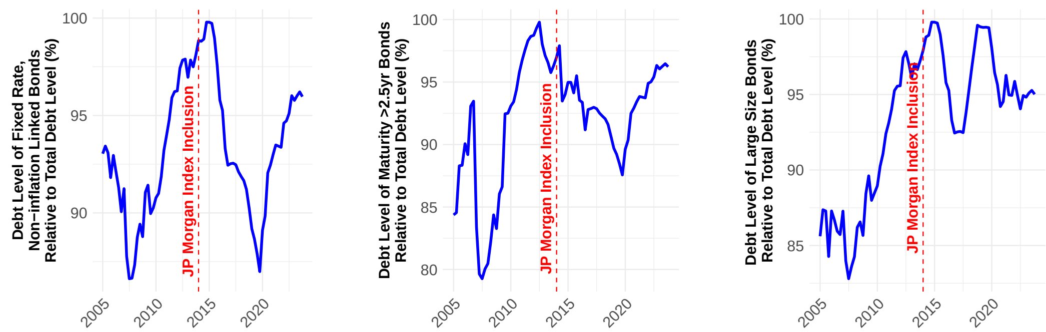
Notes: Aggregate positions of mutual funds domiciled in Ireland, Luxembourg, and the U.S. are from: www.globalcapitalallocation.com, based on the work of Maggiori et al. (2027) and Coppola et al. (2021). Foreign creditor holdings of domestic debts by currency are from Onen et al. (2022). Data for years 2008, 2009, and 2013 are missing because Colombian represented less than 1% of the total sovereign fixed income portfolios for Ireland, Luxembourg, and the US. As a result, the data omits Colombia. I estimate the present value of debt service payments. The discount rate is the average of Colombian short-term interest rate from 2005 and 2023. I retrieve the end-of-year Colombian Peso to UVR (the real value unit for inflation index bonds) exchange rate from the central bank. I use the average growth between 2005 and 2023 to interpolate for other years. When coupon payments are not available, the principal payment is used. See Appendix A.3 for more details on the construction.

eligible bonds[25]. After extracting the cyclical components of the VIX and the share of issuing index-eligible debt relative to total LC bonds at the annual frequency, I calculate their correlations.

Table B.3 Panel A presents the cyclical pattern of these currencies, while currencies that experienced inclusion/exclusion events are shown in Panel B. Countries persistently included in GBI-EM index tend to issue fewer index- eligible bonds when the VIX is high during global downturns. This negative correlation between the VIX and the issuance share of index-eligible bonds is present in nine countries, excluding Thailand and Poland.

[25]Note that JP Morgan has discretion to decide which bonds are eligible. As mentioned above, liquidity could be a concern. To the best of my knowledge, these isn't any study of JP Morgan GBI-EM that analyzes exactly which bonds are eligible. All existing approaches focus on the country weight in the GBI-EM index. Using a snapshot of the publicly available index weights of individual bonds from the GBI-EM website, I verified using eligibility criteria and identified all but a few bonds in July 2023. The bond-level index weights can be assessed here: https://www.jpmorgan.com/insights/global-research/index-research/composition. However, historical index weights are not publicly available without a paid subscription.

Figure B.3: Colombian debt level



(a) Debt level of Fixed-rate non-indexed bonds

(b) Debt level of bond with maturities ≥ 2.5yr

(c) Debt level of large size bonds

Even though two bonds differ in their index status, they face identical credit risk. The supply of index-eligible $B^{e\prime}$ and ineligible $B^{o\prime}$ bonds issued by the same country included the index should have identical prices. Let the $B_d^{e\prime}$ and $B_d^{o\prime}$ be the bond demand by global mutual funds, and the demand depends on

$$q = \frac{\mathbb{E}(\frac{s(1-d')}{s'})/(1+r^f)}{(1+r^f)D_0(\frac{1-\mu}{\gamma_L}+\frac{\mu}{\gamma_H})\operatorname{Var}(\frac{s(1-d')}{s'})\Big(\frac{s^{-1}q\times(B_d^{e\prime}+B_d^{o\prime})}{D_0}-(1-\mu)(1-\alpha)w^b(s^{-1}qB^{e\prime})\Big)+1}$$

Index-eligible and ineligible bonds have an identical demand elasticity because $\frac{\partial q}{\partial B_d^{e\prime}} = \frac{\partial q}{\partial B_d^{o\prime}}$ whenever the demand for index-eligible and index-ineligible bonds $B_d^{e\prime} = B_d^{o\prime}$. However, their supply elasticities are different. Imposing the market clearing conditions, $B_d^{e\prime} = B^{e\prime}$ and $B_d^{o\prime} = B^{o\prime}$, yields

$$q = \frac{\mathbb{E}(\frac{s(1-d')}{s'})/(1+r^f)}{(1+r^f)D_0(\frac{1-\mu}{\gamma_L}+\frac{\mu}{\gamma_H})\operatorname{Var}(\frac{s(1-d')}{s'})\Big(\frac{s^{-1}q\times(B^{e\prime}+B^{o\prime})}{D_0}-(1-\mu)(1-\alpha)w^b(s^{-1}qB^{e\prime})\Big)+1}$$

For any identical supplies $B^{e\prime} = B^{o\prime}$, the supply elasticities are different because the country internalizes the additional effect of issuing index-eligible bonds in the benchmark weight. If larger borrowing depresses bond prices in equilibrium, as in standard sovereign default models, then $\frac{\partial q}{\partial B^{e\prime}} < 0$. Using the Implicit Function Theorem, I obtain:

$$\frac{\partial q}{\partial B^{e\prime}} - \frac{\partial q}{\partial B^{o\prime}}\Big|_{B^{e\prime}=B^{o\prime}} \propto \frac{\partial w^b}{\partial(s^{-1}qB^{e\prime})}s^{-1}q \geq 0$$

Therefore, the price is less responsive to the supply of index-eligible bonds. Above patterns of issuing fewer index eligible bonds in global downturns are qualitatively consistent with a model where the supply elasticities of index-eligible and ineligible bonds are different.

Table B.3: Summary of the cyclical pattern of the debt issuance by currencies eligible for the JP Morgan GBI-EM index

Currency HP filter detrended	Avg. index eligible share (%) Y	Corr. with VIX (%) Y	Corr. with VIX (%) N
Panel a: Currencies included since the index inception			
BRL	7.22	-31.30	-24.04
CLP	4.11	-2.60	-10.85
HUF	16.08	-8.71	-18.66
IDR	29.79	-18.99	-18.72
MXN	12.98	-4.31	5.09
MYR	47.43	-13.25	-22.70
PEN	23.64	-20.72	-15.24
PLN	44.01	3.03	-2.30
THB	8.71	14.67	12.50
TRY	29.09	-11.20	-18.57
ZAR	15.68	-46.22	-38.53
Average	21.70	-12.69	-13.82
Median	16.08	-11.20	-18.57
Std.Dev	14.54	16.62	14.34
Panel B: Currencies with inclusion/exclusion events			
ARS	0.23	5.62	-1.09
CNY	54.52	-36.91	-27.79
COP	53.40	48.85	53.52
CZK	10.19	10.96	10.29
DOP	2.16	21.33	16.21
EGP	9.17	14.11	11.39
ILS	22.31	-4.37	-6.31
INR	27.87	10.25	14.05
PHP	35.98	-41.96	-38.04
RON	55.65	7.30	-15.04
RUB	18.58	32.22	27.27
Average	26.37	6.13	4.04
Median	22.31	10.25	10.29
Std.Dev	20.93	26.68	25.71

Notes: The index-eligible share refers to the proportion of index eligible bonds relative to the total LC bonds issued annually. The correlation between the VIX and the index eligible LC bond share refers to the relationship between the cyclical components of both the VIX and the share after 2005. The currencies in Panel A represent countries that have been included in the GBI-EM index since its inception and have not experienced any inclusion or exclusion events.

C Analytical appendix

C.1 Deriving portfolios

The portfolio weight w_{jt} for $j = \{\text{EM fund}, \text{Dollar fund}\}$ is

$$w_{jt} = \frac{\frac{1}{q_t}\mathbb{E}_t\left(\frac{s_t(1-d_{t+1})}{s_{t+1}(1+r^f)}\right) - 1}{(1+r^f)D_0\gamma_j \operatorname{Var}\left(\frac{s_t(1-d_{t+1})(1+r_t)}{s_{t+1}(1+r^f)} - 1\right)} + (1-\alpha)w^b_{jt} \approx \frac{\frac{1}{q_t}\mathbb{E}_t\left(\frac{s_t(1-d_{t+1})}{s_{t+1}(1+r^f)}\right) - 1}{(1+r^f)D_0\gamma_j \operatorname{Var}\left(\frac{s_t(1-d_{t+1})}{s_{t+1}}\right)} + (1-\alpha)w^b_{jt}$$

where the approximation is under a state without any currency and credit risk, that is $1/q_t = 1 + r_t \approx 1 + r^f$. The asset demand B^d_{t+1} by global mutual funds is

$$(1-\mu)D_0 \cdot w_{\text{EM Fund},t} + \mu \cdot D_0 \cdot w_{\text{Dollar Fund},t} = s_t^{-1} q_t B^d_{t+1}$$

Using the expressions of portfolio weights, the asset demand is

$$q_t = \frac{\mathbb{E}_t\left(\frac{s_t(1-d_{t+1})}{s_{t+1}(1+r^f)}\right)}{(1+r^f)D_0\left(\frac{1-\mu}{\gamma_L} + \frac{\mu}{\gamma_H}\right)^{-1} \cdot \operatorname{Var}\left(\frac{s_t(1-d_{t+1})}{s_{t+1}}\right) \cdot \left(\frac{s_t^{-1}q_t B^d_{t+1}}{D_0} - (1-\mu)(1-\alpha)w^b_t(s_t^{-1}q_t B_{t+1})\right) + 1}$$

where the supply of local currency bond B_{t+1} instead of the demand B^d_{t+1} determines the benchmark weight. I omit the time subscript and express the asset demand for B^d as

$$K(q, B^d) = q - \frac{a_0}{a_1(\frac{s^{-1}B^d}{D_0}q - a_2 w^b_t(s^{-1}qB)) + 1} = 0$$

where $a_0 = \mathbb{E}(\frac{s(1-d')}{s'(1+r^f)})$, $a_1 = (1+r^f)D_0(\frac{1-\mu}{\gamma_L} + \frac{\mu}{\gamma_H})^{-1} \cdot \operatorname{Var}(\frac{s(1-d')}{s'})$, and $a_2 = (1-\mu)(1-\alpha)$. Using the Implicit Function Theorem,

$$\frac{\partial q}{\partial w^b}\Big|_{w^b=0} = -\frac{\partial K/\partial w^b}{\partial K/\partial q} = \frac{a_0 a_1 a_2}{(a_1\frac{s^{-1}B^d}{D_0}q + 1)^2 + a_0 a_1 \frac{s^{-1}B^d}{D_0}} > 0 \tag{29}$$

which implies that index inclusion raises the price.

The slope of the demand curve

$$\frac{\partial B^d}{\partial q} = -\frac{\partial K/\partial q}{\partial K/\partial B^d} = -\frac{(a_1(\frac{s^{-1}B^d}{D_0}q - a_2 w^b) + 1)^2 + a_0 a_1(\frac{s^{-1}B^d}{D_0} - a_2 q^{-1}\epsilon^b)}{a_0 a_1 \frac{s^{-1}q}{D_0}} \tag{30}$$

where the partial elasticity of benchmark weights with respect to the market value of assets $\epsilon^b =$

$\frac{\partial w^b}{\partial(s^{-1}qB)}(s^{-1}Bq)$. Moreover,

$$\frac{\partial^2 B^d}{\partial q \partial w^b} > 0 \tag{31}$$

which means that a rising benchmark weight makes demand more inelastic.

Relationship to CRRA preference. The manager's problem under a CRRA preference is

$$\begin{aligned} \max_{b^{DM}_{j,t+1}, b_{j,t+1}} \mathbb{E}_t & \frac{(\alpha W_{j,t+1} + (1-\alpha)(W_{j,t+1} - W^b_{j,t+1}))^{1-\gamma_k}}{1-\gamma_k} \\ \text{s.t. } & D_0 = q^{DM}_t b^{DM}_{j,t+1} + s_t^{-1} q_t b_{j,t+1} \\ & W_{j,t+1} = b^{DM}_{j,t+1} + s^{-1}_{t+1}(1-d_{t+1}) b_{j,t+1} \end{aligned} \tag{32}$$

The first order condition of the portfolio choice for type j fund is:

$$\mathbb{E}_t \left[\left(\frac{s_t(1-d_{t+1})\,(1+r_t)}{s_{t+1}\,(1+r^f)} - 1 \right) \left(\alpha + \left(\frac{s_t^{-1} q_t B_{t+1}}{D_0} - (1-\alpha) w^b_t \right) \left(\frac{s_t\,(1-d_{t+1})\,(1+r_t)}{s_{t+1}\,(1+r^f)} - 1 \right) \right)^{-\gamma_j} \right] = 0$$

To derive the portfolio share, I expand $\frac{s_t(1-d_{t+1})(1+r_t)}{s_{t+1}}$ around the risk-free rate $1+r^f$ up to the second order. In other words, I solve the portfolio problem using the Taylor expansion around a state without any risk. The result for EM funds is below, and the expression of a similar form applies for dollar funds.

$$\mathbb{E}_t \left(\frac{s_t(1-d_{t+1})\,(1+r_t)}{s_{t+1}\,(1+r^f)} - 1 \right) - \frac{\gamma_L}{\alpha} \left(\frac{s_t^{-1} q_t B_{t+1}}{D_0} - (1-\alpha) w^b_t \right) \mathbb{E}_t \left(\frac{s_t(1-d_{t+1})\,(1+r_t)}{s_{t+1}(1+r^f)} - 1 \right)^2 = 0$$

The portfolio weight w_{jt} for $j = \{\text{EM fund}, \text{Dollar fund}\}$ is

$$w_{jt} = \frac{\frac{1}{q_t} \mathbb{E}_t \left(\frac{s_t(1-d_{t+1})}{s_{t+1}(1+r^f)} \right) - 1}{\frac{\gamma_L}{\alpha} \operatorname{Var} \left(\frac{s_t(1-d_{t+1})}{s_{t+1}} \right)} + (1-\alpha) w^b_{jt}$$

The asset market clearing condition is

$$(1-\mu) D_0 \cdot w_{\text{EM Fund},t} + \mu \cdot D_0 \cdot w_{\text{Dollar Fund},t} = s_t^{-1} q_t B_{t+1}$$

The expression of bond price q_t is

$$q_t = \frac{\mathbb{E}_t \left(\frac{s_t(1-d_{t+1})}{s_{t+1}(1+r^f)} \right)}{\frac{\left(\frac{1-\mu}{\gamma_L} + \frac{\mu}{\gamma_H} \right)^{-1}}{\alpha} \cdot \operatorname{Var} \left(\frac{s_t(1-d_{t+1})}{s_{t+1}} \right) \cdot \left(\frac{s_t^{-1} q_t B_{t+1}}{D_0} - (1-\mu)(1-\alpha) w^b_t \right) + 1}$$

I explain why bond prices do not enter in $\mathrm{Var}(.)$, as $\mathbb{E}_t\left(\frac{s_t(1-d_{t+1})(1+r_t)}{s_{t+1}(1+r^f)}-1\right)^2 \approx \mathrm{Var}\left(\frac{s_t(1-d_{t+1})}{s_{t+1}}\right)$. Define $\frac{1+r^f+r_{t+1}-r^f}{1+r^f} = 1+\frac{r_{t+1}-r^f}{1+r^f} \equiv 1+\Delta_t$ and $\frac{s_t(1-d_{t+1})}{s_{t+1}} \equiv x_{t+1}$. The original term can be written as

$$\begin{aligned}\mathbb{E}((1+\Delta_t)x_{t+1}-1)^2 &= \mathbb{E}(1+\Delta_t)^2x_{t+1}^2+1-2\mathbb{E}(x_{t+1})(1+\Delta_t)\\ &= \mathrm{Var}(x_{t+1})+\left(\Delta_t^2+2\Delta_t\right)\mathbb{E}\left(x_{t+1}^2\right)+\mathbb{E}^2(x_{t+1})+1-2\mathbb{E}(x_{t+1})(1+\Delta_t)\end{aligned}$$

When $\Delta_t \to 0$, $\mathbb{E}((1+\Delta_t)x_{t+1}-1)^2 = \mathrm{Var}(x_{t+1})+\mathbb{E}^2(x_{t+1})+1-2\mathbb{E}(x_{t+1})$. When $\mathbb{E}_t(d_{t+1}) \to 0$ (i.e. the default probability is approaching zero) and the exchange rate is not expected to depreciate (i.e. $\mathbb{E}\left(\frac{s_t}{s_{t+1}}\right) \to 1$), the above expression becomes $\mathbb{E}((1+\Delta_t)x_{t+1}-1)^2 = \mathrm{Var}(x_{t+1})+1+1-2 = \mathrm{Var}(x_{t+1})$

C.2 Proof of Proposition 1.

From the market clearing condition:

$$y_0 = (1+r^f)\left(\mathbb{E}(\frac{s_0(1-d_{t+1})}{s_{t+1}})\right)^{-1}\left(\frac{(1+r^f)D_0\,\mathrm{Var}(\frac{s_0(1-d_{t+1})}{s_{t+1}})}{E_\gamma}(\frac{s_0^{-1}B}{y_0D_0}-(1-\mu)(1-\alpha)w^b(\frac{s_0^{-1}B}{y_0}))+1\right)$$

Denote $R_0 \equiv \mathbb{E}_0\left(\frac{s_0(1-d_1)}{s_1}\right)$, $w_0 \equiv \frac{s_0^{-1}B}{y_0D_0}$, $\varepsilon^b \equiv \frac{\partial w^b}{\partial(s^{-1}B/y_0)}\frac{s_0^{-1}B}{y_0}$, $V_0 \equiv (1+r^f)D_0\,\mathrm{Var}(\frac{s_0(1-d_1)}{s_1})$. I rewrite the market clearing condition as

$$y_0 = (1+r^f)R_0^{-1}\left(\frac{V_0}{E_\gamma}\left(w_0-(1-\mu)(1-\alpha)w^b\right)+1\right)$$

I perform the first-order Taylor expansion of y_0, R_0, V_0 and s_0. $\wedge$ denotes the log-deviation from the initial value.

$$\begin{aligned}(1+\frac{(1+r^f)V_0}{R_0y_0E_\gamma}(w_0-(1-\mu)(1-\alpha)\epsilon^b)\hat{y}_0 = -\hat{R}_0+&\\ \hat{s}_0\frac{(1+r^f)V_0}{R_0y_0E_\gamma}(-w_0+(1-\mu)(1-\alpha)\epsilon^b)+&\\ \hat{V}_0\frac{(1+r^f)V_0}{R_0y_0E_\gamma}(w_0-(1\mu)(1-\alpha)w^b)&\end{aligned} \tag{33}$$

Define the creditor composition θ_0 as the market value of the bond held by EM funds relative to the market value of the bond held among global mutual funds.

$$\theta_0 = \frac{\left[\left(\frac{y_0R_0}{1+r^f}-1\right)\frac{1-\mu}{\gamma_LV_0}+(1-\mu)(1-\alpha)w^b\right]D_0}{\frac{s_0^{-1}B}{y_0}} \Rightarrow (1-\mu)(1-\alpha)w^b = \theta_0b_0-\left(\frac{y_0R_0}{1+r^f}-1\right)\frac{1-\mu}{\gamma_LV_0}$$

Using the definition of dollar fund ownership,

$$1-\theta_0=\frac{\left(\frac{y_0R_0}{1+r^f}-1\right)\frac{\mu}{\gamma_H V_0}D_0}{\frac{s_0^{-1}B}{y_0}}\Rightarrow\frac{y_0R_0}{1+r^f}-1=(1-\theta_0)\frac{s_0^{-1}B}{y_0D_0}\cdot\left(\frac{\mu}{\gamma_H V_0}\right)^{-1}$$

Using the above expressions, I rewrite terms with w^b.

$$w_0-(1-\mu)(1-\alpha)w^b=(1-\theta_0)w_0(1+\frac{1-\mu}{\mu}\frac{\gamma_H}{\gamma_L})$$

Define the following

$$\begin{aligned}\hat{R}_0&=\frac{\partial R_0}{\partial z_0}\frac{z_0}{R_0}\hat{z}_0\equiv\Phi_R\hat{z}_0\\ \hat{s}_0&=\frac{\partial s_0}{\partial z_0}\frac{z_0}{s_0}\hat{z}_0\equiv\Phi_s\hat{z}_0\\ \hat{V}_0&=\frac{\partial V_0}{\partial z_0}\frac{z_0}{V_0}\hat{z}_0\equiv\Phi_v\hat{z}_0\end{aligned}$$

Finally, using above results, I rewrite equation 33.

$$\begin{aligned}\hat{y}_0&=-\Phi_R\hat{z}_0+\Phi_1\left(-\Phi_R+\Phi_s+\Phi_v\Phi_2\right)\hat{z}_0,\\ \text{where }\Phi_1&=\left(1+\frac{w_0-\epsilon^b(1-\mu)(1-\alpha)}{(1-\theta_0)w_0(1+\frac{(1-\mu)\gamma_H}{\mu\gamma_L})+\frac{E_\gamma}{(1+r^f)V_0}}\right)^{-1},\\ \Phi_2&=\frac{(1-\theta_0)w_0(1+\frac{(1-\mu)\gamma_H}{\mu\gamma_L})}{(1-\theta_0)w_0(1+\frac{(1-\mu)\gamma_H}{\mu\gamma_L})+\frac{E_\gamma}{(1+r^f)V_0}}\end{aligned}$$

QED.

C.3 Proof of Proposition 2.

To show Proposition 2, when the asset is under index exclusion, $w^b=0$ and by definition $\theta_0=0$. When the index status changes from inclusion to exclusion, according to the above definition, Φ_1 and Φ_2 become larger. Therefore, $\frac{\partial\hat{y}_0}{\partial\hat{z}_0}|_{\text{exclusion}}-\frac{\partial\hat{y}_0}{\partial\hat{z}_0}|_{\text{inclusion}}>0$. Moreover, to the first order approximation around $w_0=\bar{w}$ and $\theta_0=0$,

$$\frac{\partial\hat{y}_0}{\partial\hat{z}_0}|_{\text{exclusion}}-\frac{\partial\hat{y}_0}{\partial\hat{z}_0}|_{\text{inclusion}}=-\left[(-\Phi_R+\Phi_s)\frac{\partial\Phi_1}{\partial(w_0\theta_0)}|_{\bar{w},0}w_0\theta_0+\Phi_v\frac{\partial(\Phi_1\Phi_2)}{\partial(w_0\theta_0)}|_{\bar{w},0}w_0\theta_0\right]$$

The first order approximation of $\hat{y}_0 = -\Phi_R \hat{z}_0 + \Phi_1 \left(-\Phi_R + \Phi_s + \Phi_v \Phi_2\right) \hat{z}_0$ is

$$\hat{y}_0 = \alpha \hat{z}_0 + \beta \hat{z} \times w_0 + \gamma \hat{z}_0 \times w_0 \times \theta_0$$

where $\gamma = (-\Phi_R + \Phi_s) \frac{\partial \Phi_1}{\partial (w_0 \theta_0)}|_{\bar{w},0} w_0 \theta_0 + \Phi_v \frac{\partial (\Phi_1 \Phi_2)}{\partial (w_0 \theta_0)}|_{\bar{w},0} w_0 \theta_0$.

QED.

C.4 Proof of Proposition 3.

To show Proposition 3, I first use the first-order condition at $t = 1$ when $h = 0$. The optimal choice of bonds $B^* > 0$ satisfies the following.

$$\frac{1}{Y_1 + q(h = 0, B^*) B^* - B_{01}} \left(\frac{\partial q\,(h = 0)}{\partial B} \cdot B^* + q\,(h = 0) \right) + \beta_{EM} \frac{\partial E_1 (V_2(S_2, B))}{\partial B} = 0$$

I first show when $V_1\,(h = 1, B^*) > V_1\,(h = 0, B^*)$ could happen. I analyze if the first order condition becomes positive when the index status switches to inclusion, $h = 1$, when the amount of borrowing remains the same at B^*. Given that the value function is concave in B, the first order condition being positive at B^* is a necessary condition for a higher debt level and a higher value of the value function under index inclusion. Let $q^0 \equiv q(h = 0, B^*)$ and $q^1 \equiv q\,(h = 1, B^*)$. Given that the price is higher under index inclusion, a sufficient condition is:

$$\left.\frac{\partial q^1}{\partial B}\right|_{B=B^*,q^1} - \left.\frac{\partial q^0}{\partial B}\right|_{B=B^*,q^0} > 0 \quad \text{and} \quad \left.\frac{\partial q^1}{\partial B}\right|_{B=B^*,q^1} \cdot q^0 - \left.\frac{\partial q^0}{\partial B}\right|_{B=B^*,q^0} \cdot q^1 > 0$$

I assume that above partial derivatives are negative. Quantitative sovereign default models typically imply this negative relationship because rising debt increases the default risk and depresses the price. With this assumption, a sufficient condition for more debt under index inclusion is $\left.\frac{\partial q^1}{\partial B}\right|_{B=B^*,q^1} - \left.\frac{\partial q^0}{\partial B}\right|_{B=B^*,q^0} > 0$. I characterize conditions under which this inequality holds. Let $a_0(B) \equiv \mathbb{E}_1(\frac{s_1(1-d_2)}{s_2})$, $a_1(B) = (1 + r^f) D_0 \operatorname{Var}(\frac{s_1(1-d_2)}{s_2}) \frac{1}{E_\gamma}$, $a_2 = (1 - \mu)(1 - \alpha)$, $w_0 = \frac{s_0^{-1} q^0 B^*}{D_0}$, $w_1 = \frac{s_0^{-1} q^1 B^*}{D_0}$, $\phi_0 = \frac{\partial a_0(B)}{\partial B}$, $\phi_1 = \frac{\partial a_0(B)}{\partial B} a_1 - \frac{\partial a_1(B)}{\partial B} a_0$, and $\epsilon^b = \frac{\partial w^b}{\partial (s_1^{-1} q B)} (s_1^{-1} q^1 B^*)$.

$$\left.\frac{\partial q^1}{\partial B}\right|_{B^*,q^1} - \left.\frac{\partial q^0}{\partial B}\right|_{B^*,q^0} > 0 \Leftrightarrow \begin{array}{l} (\phi_0 + \phi_1(w_1 - a_2 w^b) + a_1(w_1 - a_2\epsilon^b) B^{-1})((w_0 a_1 + 1)^2 + a_0 a_1 w_0 \dfrac{1}{q^0}) \\ -\, (\phi_0 + \phi_1 w_0 + a_1 w_0 B^{-1})(((w_1 - a_2 w^b) a_1 + 1)^2 + a_0 a_1 (w_1 - a_2 \epsilon^b) \dfrac{1}{q^1}) > 0 \end{array}$$

Under the above assumption of a lower price under higher debt, $(\phi_0 + \phi_1(w_1 - a_2 w^b) + a_1(w_1 - a_2\epsilon^b) B^{-1}) < 0$ and $(\phi_0 + \phi_1 w_0 + a_1 w_0 B^{-1}) < 0$. A sufficient condition to ensure the above inequality is that inequalities below hold at the same time

$$(\phi_0 + \phi_1(w_1 - a_2 w^b) + a_1(w_1 - a_2\epsilon^b)B^{-1}) - (\phi_0 + \phi_1 w_0 + a_1 w_0 B^{-1}) > 0$$

$$(((w_1 - a_2 w^b)a_1 + 1)^2 + a_0 a_1 (w_1 - a_2\epsilon^b)\frac{1}{q^1}) - ((w_0 a_1 + 1)^2 + a_0 a_1 w_0 \frac{1}{q^0}) > 0$$

Relationships below are sufficient conditions.

$$1 > \frac{\partial w^b}{\partial(s_1^{-1}qB)}\frac{s_1^{-1}q^1 B}{w^b}$$
$$\frac{\partial \mathbb{E}_1(\frac{s_1(1-d_2)}{s_2})}{\partial B}\operatorname{Var}(\frac{s_1(1-d_2)}{s_2}) - \frac{\partial \operatorname{Var}(\frac{s_1(1-d_2)}{s_2})}{\partial B}\mathbb{E}_1(\frac{s_1(1-d_2)}{s_2}) > 0$$
$$\frac{s_1^{-1}q^1 B^*}{D_0} - (1-\mu)(1-\alpha)w^b > \frac{s_1^{-1}q^0 B^*}{D_0}$$

QED.

C.4.1 Bond supply elasticity and benchmark weight

How does the change of benchmark weight affect the bond supply elasticity in equilibrium? Let $a_0(B) = \mathbb{E}_1(\mathcal{E}_2)$, $a_1(B) = (1+r^f)D_0\operatorname{Var}(\mathcal{E}_2)\frac{1}{E_\gamma}$, $a_2 = (1-\mu)(1-\alpha)$, $f^1(q(h=1,B),B) = \frac{s_1^{-1}qB}{D_0} - a_2 w^b$, and $f^0(q(h=0,B),B) = \frac{s_1^{-1}qB}{D_0}$.

$$\frac{\partial q}{\partial B} = \frac{\frac{\partial a_0(B)}{\partial B} + (\frac{\partial a_0(B)}{\partial B}a_1 - \frac{\partial a_1(B)}{\partial B}a_0)(\frac{s^{-1}Bq^1}{D_0} - a_2 w^b) + a_1(\frac{s^{-1}B}{D_0} - a_2 B^{-1}\epsilon^b)}{(a_1(\frac{s^{-1}Bq^1}{D_0} - a_2 w^b) + 1)^2 + a_0 a_1(\frac{s^{-1}B}{D_0} - a_2 q^{-1}\epsilon^b)} \tag{34}$$

The supply elasticity depends on the expected default probability and the variance of the expected repayment. If $\frac{\partial a_0(B)}{\partial B}a_1 - \frac{\partial a_1(B)}{\partial B}a_0 < 0$, then $\frac{\partial^2 q}{\partial B \partial w^b} < 0$, implying a rising benchmark weight makes the equilibrium price more sensitive to changes in bond supply. If it is greater than zero, the sign is ambiguous. The implication of this result is consistent with Moretti et al. (2024) (p.25), which argues that the inelastic demand is a disciplinary mechanism that limits government debt.

C.5 The relationship to the literature of benchmark contracts

This section outlines the micro-foundation of the reduced-form utility function, in which asset managers invest in an asset class that is entirely different from the asset class specified in the performance benchmark. I extend the model in Kashyap et al. (2020) by introducing heterogeneity in asset managers' private costs. Furthermore, I also account for the (non-pecuniary) cost that investors derive from holding specific assets.[26]

I consider a two-period CARA-normal model with two risky assets. The risky asset pays $\tilde{D}_i \sim N(\mu_i, \sigma_i^2)$ per unit at date 1 for $i = 1, 2$. The exogenous gross interest rate of a risk free asset is normalized to 0. The aggregate supply of the risky asset is $\bar{x}_i$. I assume the pay-off depends on a common and an idiosyncratic component, that is $\tilde{D}_i = \nu + \tilde{D}_{\epsilon,i}$ where the common shock $\nu \sim N(0, \sigma_A^2)$ and specific shocks $\tilde{D}_{\epsilon,i} \sim N(0, \sigma_{\epsilon,i}^2)$. This implies a covariance $cov(\tilde{D}_1, \tilde{D}_2) = \sigma_A^2$.

There is a unit mass of two types of agents: asset managers (indicated by M) and asset management firms (indicated by F). Both agents have an identical CARA preference over their wealth at date 1. I denote the risk tolerance by γ. The objective of the agents is to maximize the expected utility $\mathbb{E}(-exp(-\gamma W_j))$ for $j = F, M$.

Asset management firms assign the asset manager the task of managing a risky asset 2. This is reflected by a wealth loss ϕ_p of the firm per unit of asset 1 the manager invests. The manager has an lump-sum value added $\epsilon \sim N(0, \sigma_\epsilon)$. Managing assets incurs a private management cost (or negative values can be thought of as private benefits) ϕ_i per unit for $i = 1, 2$. The manager may not commit to the task and may decide to invest in risky asset 1.

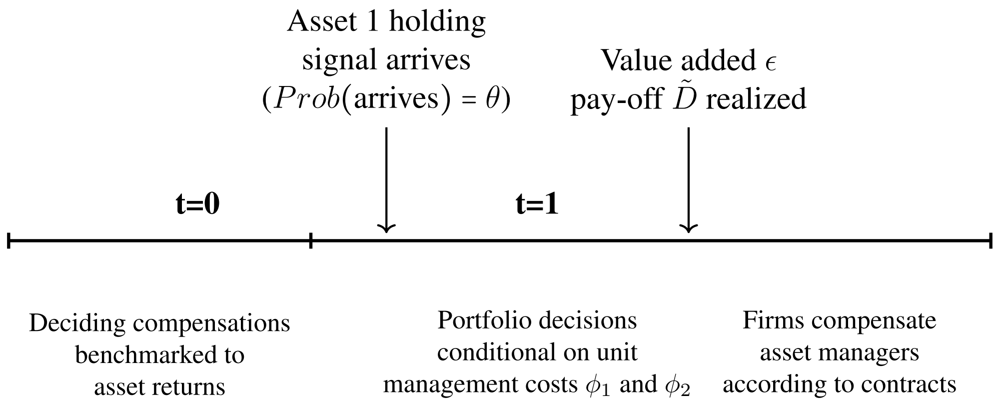
I divide the timeline into three stages and summarize it below. Figure C.4 also describes the timeline graphically. In the first stage, the asset management firm and the manager sign a compensation contract. The firm evaluates the manager's absolute return, and the return relative to assets 1 and 2. The contract (a, b, c) maximizes the welfare between the firm and the manager. Their expected utilities are U^F and U^M. Compensations depend on a fund's absolute return r_x and the return relative to benchmark indices with weights (a_0, a_1, a_2), that is $w = a_0 r_x + a_1(r_x - r_1) + a_c(r_x - r_2)$. Alternatively, the compensation can be written as $w = ar_x - br_1 - cr_2$. In the second stage, the signal of the manager investing in asset 1 arrives with a probability of θ. The manager only invest in risky asset 1 when the signal arrives. Conditional on the arrival of the signal, the manager decides to hold x_1 shares of asset 1 and x_2 shares of asset 2. Asset prices clear the market. In the last stage, values added to the management ϵ are realized. The pay-offs of assets are realized.

I solve the optimal compensation scheme when there aren't any asset 2 specific shocks (i.e. $\sigma_{\epsilon 2}^2 = 0$). I focus on parameters satisfying $1 > a^* > 0$. I summarize key insights below. Moreover, following Kashyap et al. (2020), the compensation should be benchmarked to the indices if the optimal benchmark contract contains $b > 0$ or $c > 0$.

The first result from the extension of the benchmark contract literature is that the optimal compensation may involve $b^* < 0$ and $c^* > 0$. $b^* < 0$ means that the manager is penalized even if he generates positive returns relative to the index of risk asset 1. If this penalty is not allowed in a compensation contract, then b^* should be 0. In other words, the practice of benchmarking to dollar

[26] Kashyap et al. (2024) also examines the optimal benchmark contract with asset-specific costs faced by investors.

Figure C.4: Model Timeline



indices despite of their actions of investing EM LC bonds may be optimal.

Moreover, I analyze how private costs affect the transmission of shocks. In this environment, I conduct an experiment on the how rising volatility affects the equilibrium price.

Proposition 4 *Let* $\rho_1 = \gamma(b^* \sigma_{\epsilon 1}^2 + c^* \sigma_A^2) + \theta(\phi_p + \phi_1)$ *denote the price of risk of asset 1.* $\mathbb{E}_0(r_1) = \gamma\theta(1-a^*)(\sigma_{\epsilon 1}^2 + 2\sigma_A^2)\bar{x}_1 + \rho_1$, *where the first term represents the quantity of risk. The following holds.*

1. $\frac{\partial \rho}{\partial \sigma_{\epsilon 1}^2} > 0$.
2. $\frac{\partial \rho}{\partial \sigma_A^2} > 0$ *or* < 0. *Specifically, when* $\frac{1-\theta}{\gamma a^*}(\frac{1}{a^*} - 1)\phi_1 > \phi_p + (3+\theta)\sigma_{\epsilon 1}^2(2 - \frac{1}{a^*})\bar{x}_1$, $\frac{\partial \rho}{\partial \sigma_A^2} < 0$.
3. *If* $\frac{\partial \rho}{\partial \sigma_A^2} < 0$ *for any* $(\tilde{\phi}_1, \tilde{\phi}_2, \tilde{\phi}_p)$, *then* $\frac{\partial \rho}{\partial \sigma_A^2} < 0 \forall \phi_2 > \tilde{\phi}_2$.

This proposition summarizes how specific volatility ($\sigma_{\epsilon 1}^2$) and common shocks (σ_A^2) affect the price of risk. Amplification through balance sheet channels implies that financial frictions increase the risk premia regardless of whether the shocks are aggregate or idiosyncratic. The first result describes an amplification mechanism, showing that the price of risk increases when asset 1 specific volatility rises. The second result reflects a key departure from the traditional channels: there is an ambiguous effect on the price of risk when common volatility of common shocks.

Surprisingly, common shocks may reduce the price of risk. To induce the manager to commit to investing in asset 2, the firm would share the risk of asset 2 and reduce the incentives of investing asset 1. When a common shock drives asset returns, this compensation scheme effectively reduces the impact of common shocks on the asset manager's wealth.

The third result highlights how stabilization could happen. These conditions occur when the private cost of investing asset 1, ϕ_1, is small or even negative (i.e. there may be a private benefit.). Moreover, increasing the value of ϕ_2 may also produce this result.

Summary. 1) What dollar funds do in practice—investing in an out-of-benchmark asset type—may be optimal, given that the structure of private costs cannot be internalized in the compensation contract. 2) Benchmarking changes the composition of risk when it is micro-founded. Benchmarking increases the impact of idiosyncratic volatility on expected returns, but it may mitigate the impact of common shocks. For the above result to hold, the manager must have different private benefits/costs when investing in different assets.

Proof of Proposition 4.

I solve the optimal compensation scheme (a^*, b^*, c^*) when $\sigma_{\epsilon 2}^2 = 0$. a^* satisfies

$$\gamma a^{*2}(2a^* - 1)\sigma_{\epsilon 1}^2 = \frac{1}{\sigma_A^2}(\frac{1}{a^*} - 1)\phi_1\phi_2 + \frac{1}{\sigma_{\epsilon 1}^2}\theta(\phi_1 - \phi_2)\big(\frac{\phi_p}{a^{*2}} + (\frac{1}{a^*} - 1)(\phi_1 + \phi_p)\big) \tag{35}$$

The expected return of asset 1 is

$$E_0(r_1) = \gamma\theta(1 - a^{*2})(\sigma_{\epsilon 1}^2 + 2\sigma_A^2)\bar{x}_1 + \gamma(b^*\sigma_{\epsilon 1}^2 + c^*\sigma_A^2) + \theta(\phi_p + \phi_1) \tag{36}$$

The term $\gamma(b^*\sigma_{\epsilon 1}^2 + c^*\sigma_A^2) + \theta(\phi_p + \phi_1)$ can be thought as the price of risk. It depends on the returns relative to the benchmarks of asset 1 and 2 given the weights b^* and c^*. Under this special parameter case, the volatility of the common shock does not affect the value of a^*. Furthermore, the price of the risk is

$$\theta(\phi_p + \phi_1) + \theta a^*\gamma\frac{\sigma_{\epsilon 1}^2}{2\sigma_{\epsilon 1}^2 + (1-\theta)\sigma_A^2}\left(\bar{x}_1(2 - \frac{1}{a^*})(\sigma_{\epsilon 1}^2 + 2\sigma_A^2) + \frac{1}{\gamma a^*}\big(\phi_1(\frac{1}{a} - 1) - \phi_p\big)\right) \tag{37}$$

I can derive the Proposition 4 using equations 35 and 37.

QED.

D Additional tables

Table D.4: Summary Statistics by Currency

Currency	Number of Bonds	Amount Issued ($B)	Coverage (%)	YTM(%)	Residual Maturity
BRL	18	27.39	4.60	10.95	7.06
CLP	14	4.42	5.80	4.82	13.32
COP	29	7.58	9.09	7.52	8.21
CZK	48	4.55	2.46	2.54	10.97
HUF	50	4.29	6.95	4.92	7.06
IDR	99	4.80	8.28	7.68	11.62
ILS	37	5.52	1.14	2.27	8.71
MXN	36	15.81	6.97	7.05	11.65
MYR	133	3.83	3.80	3.75	9.17
PEN	8	4.45	4.81	6.03	15.75
PLN	38	10.23	5.33	3.56	7.02
RON	56	2.90	3.58	4.91	5.86
THB	67	4.54	2.83	2.89	13.58
TRY	64	7.25	5.37	12.73	5.18
ZAR	16	20.99	5.97	9.27	16.42

Table D.5: Summary of reported defaulted local-currency bonds by country (06/2019-12/2021)

country	# of defaulted bonds	# of funds reported	average market value (USD M)
AR	7	16	0.19
BR	2	3	29.83†
CL	1	1	2.48
CN	1	1	1.41
CO	1	1	0.58
DO	1	1	1.17
TH	1	3	1.65
UY	2	3	1.36
ZA	2	2	1.85

Notes: This table tabulates the statistics of defaulted bond reported by US mutual funds from Form N-PORT-P filings.†:The extreme value comes from filings of two funds managed by Carillon Reams. For example, Carillon Reams Unconstrained Bond Fund reported that a Brazilian sovereign bond in BRL was in default in 04/2021.

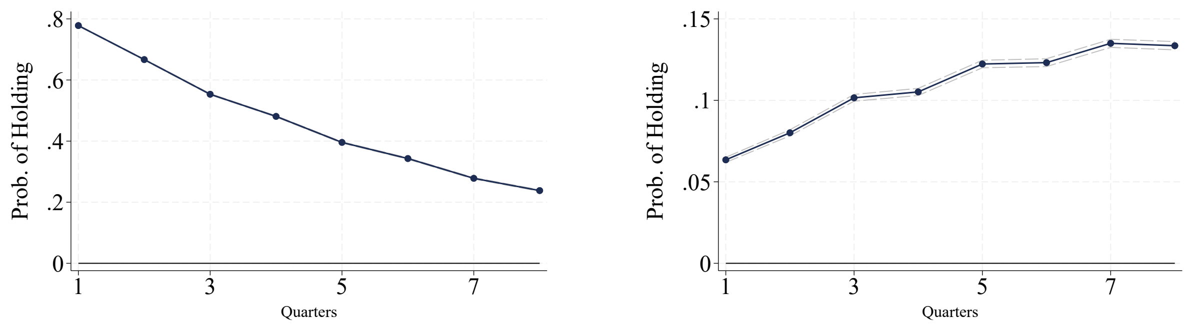
Table D.6: Benchmark index inclusion and exclusion events after index inceptions

Currency	Bloomberg EM LC	Bloomberg Global Aggregate	Bloomberg Pan-European Aggregate	FTSE International ILSI	FTSE WBIG	FTSE World ILSI	JPM GBI-EM	JPM GBI-EM Broad	JPM GBI-EM Global	Morningstar Global Treasury
COP		09/2020 in					03/2014 in†	03/2014 in†	03/2014 in†	
CZK			01/2005 in					02/2017 in	02/2017 in	
HUF		11/2013 ex 04/2017 in	11/2013 ex 04/2017 in							
IDR		06/2018 in								
ILS		01/2012 in			05/2020 in	05/2020 in				10/2022 in
MXN				12/2013 in	10/2010 in	12/2013 in				10/2022 in
MYR		01/2006 in			07/2007 in					10/2022 in
PEN		09/2020 in								
PLN				12/2013 in 02/2014 ex		12/2013 in 07/2021 ex				10/2022 in
RON	04/2013 in	09/2020 in	09/2020 in					01/2013 in	01/2013 in	
RUB	03/2022 ex	04/2014 in 03/2022 ex	04/2014 in 03/2022 ex							
THB		03/2007 ex 07/2008 in								
TRY		04/2014 in 10/2016 ex								
ZAR		05/2018 ex		12/2013 in	10/2012 in 05/2020 ex	12/2013 in 05/2020 ex				

Notes: This table outlines the inclusion and exclusion events of countries following index inceptions. The event dates are sourced from publicly available index methodologies provided by index providers. Brazil has intermittently imposed explicit capital controls. According to He and Beltran (2022), Brazil experienced a brief period of exclusion from the JPM GBI-EM Global index. However, to the best of my knowledge, without a paid subscription to index providers, public news sources have not detailed Brazil's index exclusions. The symbol † denotes COP bonds (Colombian TES) issued in Colombian local markets.

E Additional figures

Figure E.5: Extensive margin of fund holding

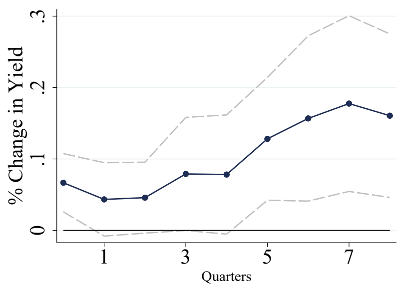


(a) Extensive margin, fund holding effect

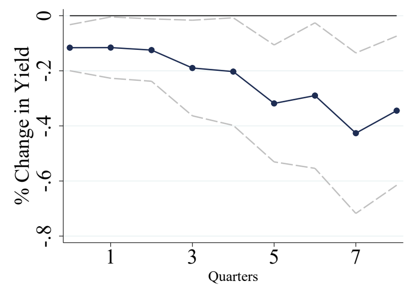
(b) Extensive margin, benchmarking effect

Notes: For a bond currently in a funds' portfolio, this two graphs show the hazard rate that a fund continues holding this bond in two years. A fund stops holding a bond may due to a fund doesn't report a bond in its holding or due to a fund's liquidation. The estimation controls bond and country-time fixed effects. The first graphs shows the hazard rate of dollar funds, and the second shows the difference in hazard rates between dollar and EM funds.

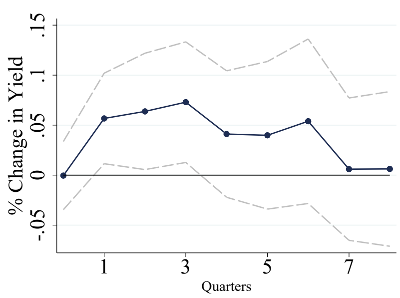
Figure E.6: Effects of benchmarking under alternative specifications



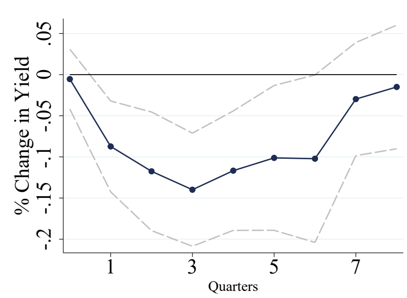
(a) VIX, fund holding effect (global LC bonds)



(b) VIX, benchmarking effect (global LC bonds)



(c) VIX, fund holding effect (additional controls)



(d) VIX, bechmarking effect (additional controls)

Notes: These figures show the direct effect of bond holdings by global funds and the interaction effect from the creditor composition of funds. The sample of bonds is restricted to bonds with issuance greater than one billion dollars and residual maturities greater than two years. Currencies are BRL, CLP, COP, CZK, HUF, IDR, ILS, MXN, MYR, PEN, PLN, RON, THB, TRY, ZAR. Dotted lines plot the 90% robust standard error bands.

F Quantitative model appendix

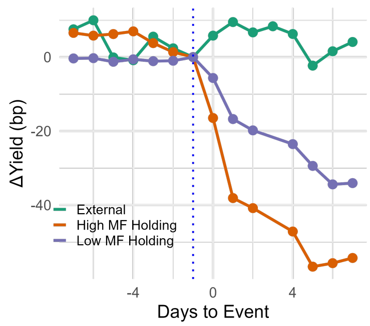
F.1 Index inclusion events

I provide further evidence on how a country's benchmark inclusion affects bond prices. I focus on EM's inclusion in JP Morgan's GBI-EM Index and Citi's WBI local currency bond index. Broner et al. (2021) analyzed these index inclusion events.

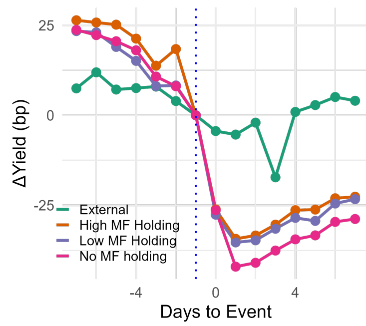
Using Asset Ownership Database in Bloomberg Terminal, I count the number of investment companies domiciled outside these countries reported holding in the quarter prior to the announcement. I classify local currency bonds into three categories based on how many different investment companies hold a bond. Low mutual fund holdings mean that fewer than five global investment companies report holdings. External bonds are dollar-denominated bonds issued in the global markets.

Figure F.7 shows that index inclusion has an impact on bonds already held by mutual funds. It is surprising that for bonds not held by global mutual funds in Romania and South Africa, they experienced larger changes in yields than sovereign bonds directly affected by the index inclusion shock. I have to note that standard theories cannot generate an overreaction if the inclusion shock only affects the quantity of bonds in the market. The co-movement of prices across bonds with different degrees of mutual fund ownership indicates a high substitution across bonds. The direct effect of index inclusion in the index-eligible bonds is attenuated. For this reason, examining the inclusion episode of Romania may underestimate the direct effect of inclusion.

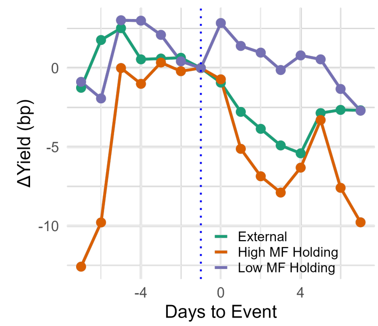
Figure F.7: Effects of index inclusion on sovereign bond yields



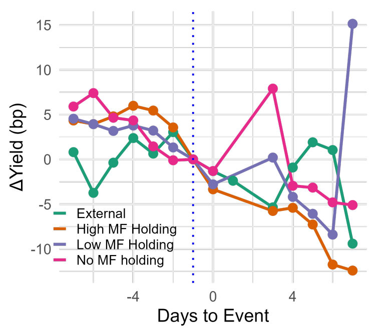
(a) Colombia



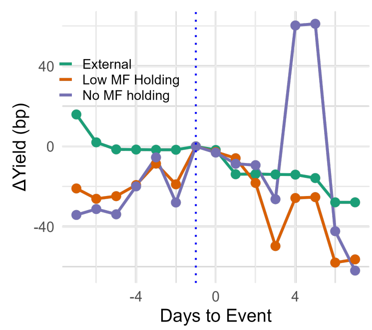
(b) Romania



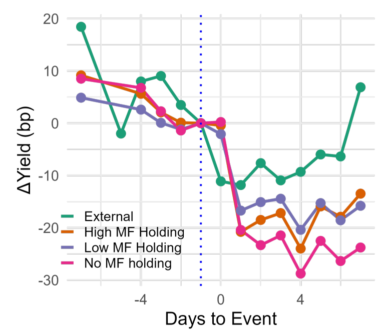
(c) Czech Republic



(d) Mexico



(e) Nigeria



(f) South Africa

Notes: This set of figures aims to replicate the analysis of the index inclusion events in Broner et al. (2021). I extend their study by including dollar-denominated bonds. Moreover, I use an alternative approach to measure the influence of the announcement of inclusion: index-eligible bonds are often widely held by global mutual funds even prior to the index inclusion events. The Security Ownership on Bloomberg Terminal, a database comparable to Morningstar's mutual fund database, directly reports for each bond the number of global investment firms, the ultimate owners of mutual funds, holding. "Low MF Holding" refers to bonds invested by fewer than five investment firms domiciled outside of respective countries in the quarter prior to the announcement.