

Institutional Investment and International Risk-sharing

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Abstract

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JEL classification: G12, G15, F30, G20.

Keywords: international asset pricing, institutional investors, underdiversification.

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Abstract

This paper proposes a new channel for local risk premia in international stocks based on investor underdiversification. Holdings data reveal two facts. First, retail investors display near-perfect home bias while institutions are globally diversified. Second, institutions substantially underdiversify within each market. Institutions' imperfect cross-market hedging generates an institutional local premium, while their limited investment scope leaves residual risk, generating a retail local premium. These premia average 2.2% and 2.5% across 21 developed markets and 3.5% and 4.5% across 20 emerging markets. Unlike the investability channel, local risk premia persist even when markets are fully open to foreign investors.

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A long-standing question in international asset pricing is whether stocks are priced locally or globally (Karolyi and Stulz, 2003). The answer matters for international portfolio choice, the performance evaluation of global asset managers, and firms' cost of capital worldwide. If risk-sharing is perfect across equity markets, stocks should be priced only by global factors. Yet local asset-pricing models consistently outperform global ones (e.g., Lewis, 2011; Hou et al., 2011; Fama and French, 2012; Petzev et al., 2016; Hollstein, 2022).

The existing literature explains local pricing through differences in securities' investability.¹ A stock is classified as investable if its shares trade in a market without material barriers to foreign investment, either through its domestic listing or a cross-listing abroad (Karolyi and Wu, 2018). Classic theories predict that removing such barriers eliminates local risk premia (e.g., Stulz, 1981; Errunza and Losq, 1985; De Jong and De Roon, 2005). Yet local pricing persists in markets that are open to global investors. Moreover, regulatory openness and implicit barriers proxied by political risk and financial development do not fully explain cross-market valuation differentials (Bekaert et al., 2011), and globalization measures have limited explanatory power for cross-market return comovement (Bekaert et al., 2016). This paper instead focuses on the revealed investment scope of different investor types. Heterogeneous underdiversification across investor types, observed in holdings data, defines the *de facto* market structure, yielding a local premium for investable securities and a decomposition of the local risk premium by investor type, two implications absent from the investability channel.

Figure 1 reveals two striking new facts about investor diversification across countries and within markets as of December 2024. First, while households exhibit almost perfect home bias, institutional investors do not. Panel A reports the home bias of three investor groups, defined as the difference between each group's actual domestic portfolio weight and the market's weight in the world market portfolio. On average, the home bias of households is around 85% in developed markets and around 95% in emerging markets. In contrast, the home bias of *global institutional*

¹A notable exception is Lau et al. (2010), who study local pricing implied by the home bias of domestic investors. Other studies derive local pricing from a linear pricing kernel specification that loads on local factors; see e.g., Bekaert and Harvey (1995), Bekaert et al. (2007).

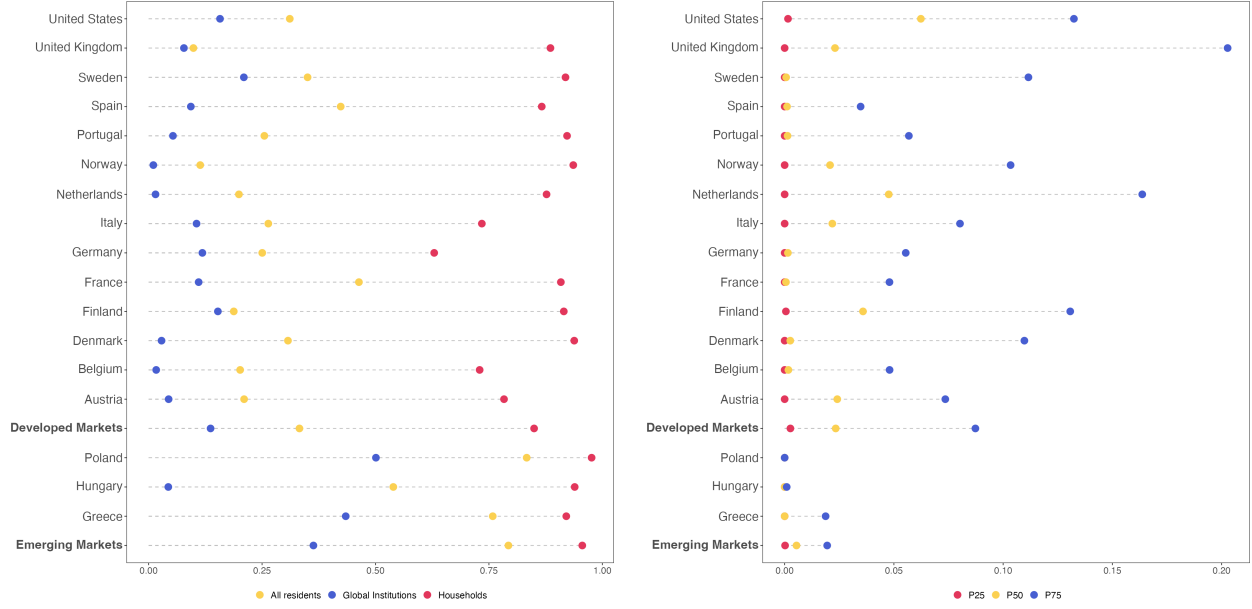
investors, defined as institutions with substantial international diversification (following Bartram et al. (2015); see Appendix A for details), is only 14% in developed markets and 36% in emerging markets.² Second, even though institutional investors diversify internationally, their holdings are concentrated among a limited set of stocks within each market, consistent with a limited investment scope that may be shaped by investment mandates and benchmarking incentives (Basak and Pavlova, 2013; Pavlova and Sikorskaya, 2023). Panel B reports the quartiles of stock-level ownership by global institutions, defined as the total number of shares held by global institutions divided by the total number of shares outstanding, in the same markets. The median institutional ownership is near zero in many markets, and is exactly zero in 17 of the 41 markets used in the asset-pricing tests. Together, these two findings reveal segmentation both across and within international stock markets.³

To study the implications of this market structure, I develop and estimate a new asset-pricing model with two investor types. Global investors, referred to as *institutional investors*, diversify internationally but invest only in securities that fall within their investment scope. Local investors, referred to as *retail investors*, invest exclusively domestically but hold the full set of available local assets. Under this setup, local pricing reflects heterogeneous underdiversification across investor types, not security-level investability.

This market structure yields two new asset-pricing implications. First, when cross-market hedging is imperfect, securities held by global institutions earn an institutional local premium despite being internationally investable. Second, the market-wide local risk premium can be decomposed into two components with distinct economic origins. The *institutional local (market) risk premium* is governed by institutional investors' risk-bearing capacity and their ability to hedge across markets. The *retail local (market) risk premium* is governed by retail investors' risk-bearing capacity, with the set of securities earning this premium determined by institutional investors' investment scope. This decomposition implies that expanding investability can shift the

²Internet Appendix Table IA.1 reports the detailed values.

³These two findings are stable over time; Internet Appendix Figures IA.2 and IA.3 report the time series of home bias and global institutional ownership quartiles from 2001 to 2024.



A. Home bias by investor group

B. Global institutional ownership quartiles

Figure 1. Home Bias and Global Institutional Ownership, December 2024.

Panel A plots the home bias of three investor groups—all residents, global institutions, and households—in developed and emerging markets as of December 2024. Home bias is defined as the difference between each group’s actual domestic portfolio weight and the country’s weight in the world market portfolio, using portfolio holdings from the IMF Coordinated Portfolio Investment Survey (CPIS), OECD National Accounts, and FactSet. Panel B reports the quartiles of global institutional ownership among stocks in the CRSP/Compustat Global universe as of December 2024. Stock-level ownership is calculated as the shares held by global institutions, divided by the total number of shares outstanding, using FactSet Ownership, following Ferreira and Matos (2008). The figure includes economies with complete data coverage and reports averages for developed (excluding the United States, which is reported separately) and emerging markets. Appendix A provides details on variable construction and data sources.

composition of local premia by enabling institutional participation but need not eliminate them. In contrast, under the investability channel, local premia vanish once a security becomes investable. Empirically, unconditional estimates of both components are economically meaningful, averaging 2.2% and 2.5% in developed markets and 3.5% and 4.5% in emerging markets, and positive and statistically significant at the 5% level (one-sided) in 15 and 18 of the 41 economies in the sample.

The mechanism is intuitive. Heterogeneity in investment scope across investor groups generates two layers of partial segmentation. Home bias segments markets across countries, while institutions’ limited investment scope segments securities within countries. Institutional investors facilitate cross-border risk-sharing only through securities within their investment scope, defined

as *institutional securities*. Consequently, the degree of international risk-sharing depends on three aspects of institutional investment: the *coverage* of the institutional investment scope (how fully institutional securities span local investment opportunities), the *correlation* between institutional security returns across countries (how well risks can be hedged across markets), and institutional investors' risk-bearing *capacity*.

To illustrate, consider risk-sharing between two stock markets, labeled domestic (the United States) and foreign (China). Global institutional investors, such as AustralianSuper, invest internationally but only in securities within their investment scope (institutional securities). They therefore obtain exposure to the foreign market through return components spanned by foreign institutional securities, defined as *attainable returns*, which enter international risk-sharing. The remaining unattainable risk is borne by foreign retail investors, who demand a *retail local risk premium*. Because retail investors do not hold assets from the other country, institutional investors are the sole mediators of cross-border risk-sharing, hedging foreign positions with domestic institutional securities. This intermediation generates an *attainable world market risk premium* for covariance with the attainable component of the world market portfolio. When institutional securities are imperfect substitutes across countries,⁴ institutional investors and foreign retail investors must bear residual attainable risk that cannot be fully hedged, giving rise to an *institutional local risk premium*, which declines with institutional risk-bearing *capacity*. In summary, all securities earn an attainable world market risk premium and an institutional local risk premium, while securities outside institutional investors' investment scope additionally earn a retail local risk premium.

The model provides a framework for assessing how institutional investment shapes the composition of local risk premia through three empirically relevant channels: capacity, correlation, and coverage. Specifically, the institutional local premium relative to the attainable world premium is proportional to inverse capacity times the unspanned variance of the attainable domestic market. Both components of this decomposition can be measured directly from the data. Capacity

⁴Substitutability refers to the extent to which a security can be replicated by other securities; see also Greenwood et al. (2018).

corresponds to the ratio of the total assets under management (AUM) of global institutional investors participating in a market to its total market capitalization. The proportion of unspanned local risk is negatively related to the correlation between the attainable domestic market and its substitute portfolio spanned by foreign institutional securities. Coverage, measured at the stock level, determines individual securities' exposure to unattainable local risk. Over the past two decades, institutional investors have substantially expanded both their ownership stakes in public firms and their international diversification (Faia and Ferreira, 2017); the framework can thus be used to understand the implications of these trends.

Mapping the model to the data requires classifying stocks into institutional or retail segments. To identify the boundary of global institutions' investment scope, I construct a stock-level coverage measure based on the revealed choice sets of active global open-end mutual funds, using quarterly holdings from FactSet Ownership, and show that this measure predicts future entry by global funds and global institutional ownership better than alternative indicators of scope, such as index membership and investability.

I estimate the model using monthly returns of 66,425 individual stocks from the CRSP/Compustat Global universe, covering the period from January 2003 to December 2024 across 41 economies. Individual stock returns preserve the within-market variation in global institutional coverage that identifies the local premia. I adopt the conditional two-pass regression approach of Gagliardini et al. (2016), which allows for the estimation of both unconditional and time-varying risk premia at the individual stock level. Test assets exclude stocks with secondary cross-listings on open exchanges as defined in Karolyi and Wu (2018), isolating the investor underdiversification channel from the investability channel. The estimation imposes the model-implied restriction that the attainable world premium is the same across all markets.

The empirical analysis proceeds in three steps. I first establish unconditional evidence on the magnitude and statistical significance of the institutional and retail local premia, and relate the cross-market variation in the institutional local premium to the capacity and correlation channels through the model's decomposition. I then use the conditional specification to study

the time-series variation in these risk premia. Finally, I examine how institutional ownership is associated with factor exposures and risk premia in the cross-section of individual stocks, using the conditional estimates.

In unconditional estimation, the institutional local premium is positive and significant in 8 out of 21 developed and 7 out of 20 emerging markets, and the retail local premium in 8 developed and 10 emerging markets. The levels and significance of the two local premia remain similar after controlling for additional local factors. In particular, in the presence of the two local factors, local size and value factors are priced in only 4 and 2 of the 41 markets, respectively, although local momentum remains significant in most markets. The main prediction of the model, however, is not the universal importance of the local premia but their cross-market variation. Consistent with the model-implied decomposition, the institutional local premium is lower in markets with greater global institutional capacity and higher in markets where a larger share of local risk is unspanned.

Conditional estimation serves two purposes: to examine the coverage channel by estimating how a stock's factor exposures vary with its global institutional ownership, and to characterize the dynamics of the risk premium components descriptively. Time variation in risk premia and factor exposures is modeled through instruments: dividend yields and dividend-yield spreads are common instruments for time-varying risk premia, and the within-market percentile rank of global institutional ownership is the stock-specific instrument for factor exposures. Consistent with the coverage channel, a stock's exposure to the retail local factor declines as its global institutional ownership rises, with a negative median coefficient in 29 of the 40 markets in the conditional sample. The conditional estimates also reveal distinct dynamics across risk premium components. The attainable world and the emerging-market institutional local premia are countercyclical, rising during major distress episodes such as the Global Financial Crisis and the Covid-19 crisis. By contrast, the developed-market retail local premium is procyclical, declining during these episodes.

The negative coefficient of the retail local exposure on the ownership rank captures a within-stock time-series effect but is silent about how institutional ownership relates to the composition of risk premia in the cross-section. To quantify the within-market cross-sectional variation, I run

panel regressions at the individual stock level of estimated time-varying factor exposures and risk premia on the ownership rank, controlling for stock characteristics and country-time fixed effects. Higher institutional ownership is negatively and significantly associated with exposure to the retail local factor and the retail local premium. It is also associated with greater exposure to the attainable world factor and a higher world premium; the total risk premium is significantly higher in developed markets and not significantly different in emerging markets. The evidence thus points to a shift in the composition of risk premia rather than a reduction in their total.

This paper builds on several strands of the literature, most directly on work linking investor-level frictions to international asset pricing. For example, Lau et al. (2010) study the asset-pricing implications of home bias using a partial-equilibrium framework, but treat investors as homogeneous within each country. Akbari et al. (2022) show that cross-border funding frictions at the investor level (leverage constraints on foreign equity positions) generate temporary integration reversals, linking underdiversification to funding liquidity, but also do not distinguish across investor types. By allowing global institutional investors and home-biased retail investors to underdiversify differently, as observed in holdings data, this paper generates a decomposition of local risk premia by investor type. Both components are economically meaningful across a wide range of markets and exhibit distinct cyclical dynamics.

This paper relates to the literature on the asset-pricing implications of investor underdiversification arising from information frictions, preferences, or exclusionary screening (Merton, 1987; Fama and French, 2007; Luo and Balvers, 2017; Zerbib, 2022). Unlike Merton (1987), who studies underdiversification in a one-factor setting in which idiosyncratic risk is priced, this paper allows for a non-diagonal covariance structure across the local components of asset returns, so that systematic local risk is priced, which is essential in an international context. This approach is closer in spirit to Luo and Balvers (2017) and Zerbib (2022), but differs by allowing multiple groups of investors to underdiversify simultaneously, with no common set of securities held across all investor groups. As a result, systematic local risk is priced separately across market segments, depending on the international diversification by global investors.

This paper also contributes to the growing literature on how institutional investment shapes international stock returns. Bartram et al. (2015) and Faias and Ferreira (2017) show that foreign institutional ownership drives return comovement and shifts factor loadings from country to global and industry factors. Institutional portfolio rebalancing also transmits funding shocks across borders through the fire-sale mechanism documented in Jotikasthira et al. (2012). Other work shows that foreign institutional ownership improves price informativeness by replacing less-informed local retail investors (Kacperczyk et al., 2021) and that heterogeneous portfolio reallocations by foreign institutions during periods of elevated global volatility affect the cross-sectional distribution of stock volatility and liquidity (Kacperczyk et al., 2025). While these studies focus on return variation or stock characteristics, this paper emphasizes the levels of global and local risk premia. In the model, global institutional capacity determines the price of institutional local risk at the market level, while coverage shapes individual securities' exposure to global and local risks through cross-border risk-sharing across segmented markets.

Section 1 develops the asset-pricing model; Section 2 introduces the data and econometric framework; Section 3 presents the empirical results; and Section 4 concludes.

1. Model

1.1. Setup and Assumptions

The economy consists of two markets, domestic (D) and foreign (F). Bold letters denote vectors, ' denotes the transpose, and * indicates foreign.

Assumption 1 (Currency denomination). All returns are denominated in domestic currency. To isolate the implications of underdiversification, I assume that purchasing power parity holds, as in Stulz (1981), so that exchange rate risk does not command additional risk premia.⁵

⁵Relaxing this assumption would augment each equilibrium risk premium with an additional component reflecting assets' covariance with real exchange rate fluctuations (Chaieb and Errunza, 2007). Institutional investors face significant foreign exchange risk and routinely hedge their exposures (Du and Huber, 2024), and the cost of hedging may itself contribute to observed market segmentation (Camanho et al., 2022). I also abstract from currency carry arising from interest-rate differentials, which would require modeling currency forwards and country-specific

Assumption 2 (Investors). There are three representative investors: a global institutional investor (i), a domestic retail investor (d), and a foreign retail investor (f). Each investor k chooses its portfolio over an exogenous choice set C^k . The institutional investor's choice set contains a subset of securities from each market. In practice, many investable stocks may remain outside the institutional choice set. Domestic and foreign retail investors are fully home-biased: their choice sets contain all securities in their respective home markets but none abroad. The institutional investor represents globally diversified institutional investors, while local institutions that invest only at home are grouped with retail investors.⁶

Assumption 3 (One-period portfolio choice). Investors have CARA preferences and choose portfolios at time t to maximize expected utility of wealth at time $t + 1$. Each investor $k \in \{d, f, i\}$ has absolute risk aversion γ^k and solves:

$$\begin{aligned} \max_{\mathbf{x}_t^k} E_t[U(W_{t+1}^k)] &= \max_{\mathbf{x}_t^k} E_t[-\exp(-\gamma^k W_{t+1}^k)] \\ W_{t+1}^k &= W_t^k(1 + r_{f,t}) + \sum_{j \in C^k} x_{j,t}^k r_{j,t+1}, \end{aligned}$$

where $\mathbf{x}_t^k$ is the vector of dollar holdings of securities in investor k 's choice set C^k , and $r_{j,t+1}$ is the *excess return* on security j over the risk-free rate.

Assumption 4 (Market structure). A market segment consists of securities priced by the same marginal investors. The home bias of retail investors and the limited choice set of institutional investors imply four market segments: domestic and foreign *institutional securities* (I and I^*) and

short rates; Opie and Riddiough (2024) document that mutual funds' separate long-short currency positions yield approximately zero Sharpe ratios. These considerations lie outside the paper's focus on how *de facto* segmentation observed from holdings data shapes the cross-section of stock returns.

⁶A limited institutional choice set may arise from formal investment mandates, limited information (Merton, 1987), regulatory constraints, environmental, social, and governance (ESG) considerations (Matos, 2020), information costs (De Marco et al., 2022), or agency frictions in institutional money management (Pavlova and Rigobon, 2008). The home bias of domestic institutional investors is studied in Chan et al. (2005), Hau and Rey (2008), and Lau et al. (2010).

domestic and foreign *retail securities* (R and R^*). In total, there are $N = N_R + N_I + N_{I^*} + N_{R^*}$ risky assets, where N_R , N_I , N_{I^*} , and N_{R^*} denote the number of securities in each segment. While strict market segmentation might seem extreme, unlike Zerbib (2022), who interprets segmentation literally as exclusionary screening, here it is a limiting case that captures the heterogeneous preferences of global and local investors.

Assumption 5 (Security returns). $\mathbf{r}_R$, $\mathbf{r}_I$, $\mathbf{r}_{I^*}$, and $\mathbf{r}_{R^*}$ denote the vectors of excess returns for each segment; time subscripts are omitted for simplicity. For example, $\mathbf{r}_I$ is an $N_I \times 1$ vector of domestic institutional returns. Excess returns are jointly normally distributed with an exogenous covariance structure. Let V_{IR} be the $N_I \times N_R$ covariance matrix between domestic institutional and domestic retail securities, and V_{jI} the $1 \times N_I$ vector of covariances between security j and domestic institutional securities; analogous notation applies to other covariances. The supply of securities, measured by market capitalization, is exogenous,⁷ and equilibrium risk premia $\boldsymbol{\mu}$ are determined endogenously by investor optimization and market clearing.

Assumption 6 (Borrowing and short-selling). Investors can borrow and lend at an exogenous risk-free rate r_f . There are no short-sale constraints.

A distinct feature of this setup, relative to existing international asset-pricing models with partial segmentation (e.g., Alexander et al., 1987; De Jong and De Roon, 2005; Chaieb and Errunza, 2007), is the absence of a commonly investable set shared by all investors. When such a set exists, commonly investable securities are priced globally by their covariance with the world market portfolio of all assets.⁸ In the current model, retail investors' choice sets do not overlap across countries due to home bias, and international risk-sharing therefore relies on the institutional investor's global diversification. Although Merton (1987) also considers heterogeneous investment sets, his model imposes a single-factor structure on asset returns and allows all investors to trade

⁷This follows typical CAPM-type models. See Zerbib (2022) for a recent example.

⁸Section 1.2 discusses this in more detail as a special case of the full model.

a forward contract on the common factor.

This market structure resembles models with arbitrageurs operating across segmented markets (e.g., Gromb and Vayanos, 2002; Greenwood et al., 2018), but differs in two respects. First, institutions face a limited choice set rather than being able to trade all securities. Second, the model allows multiple securities within each segment, without imposing strict restrictions on their covariance structure.

1.2. Equilibrium Risk Premia

By symmetry, I present equilibrium results for the domestic side; the foreign results follow analogously. Let D denote the *domestic market portfolio*, the value-weighted portfolio of all domestic securities, with excess return r_D and market capitalization M_D . The domestic *institutional* and *retail portfolios*, I and R , are the value-weighted portfolios of domestic institutional and retail securities, with excess returns r_I and r_R and market capitalizations M_I and M_R . Foreign market, institutional, and retail portfolios are defined analogously, with excess returns r_F , r_{I^*} , and r_{R^*} , and market capitalizations M_F , M_{I^*} , and M_{R^*} .

Any domestic return r_j can be decomposed into a component spanned by domestic institutional securities and a residual. The linear projection of r_j on r_I yields the *attainable return* $\hat{r}_j$:

$$\hat{r}_j = B_{jI} r_I, \quad (1)$$

where $B_{jI} = V_{jI} V_{II}^{-1}$. The attainable return $\hat{r}_j$ is the return on the corresponding mimicking portfolio, termed the *attainable portfolio*, and represents the component of r_j accessible to institutional investors. The residual $\varepsilon_j = r_j - \hat{r}_j$ is held only by domestic retail investors. Institutional securities are attainable by definition. A retail security is also attainable if it can be perfectly replicated by domestic institutional securities.

The *attainable domestic market portfolio* $\hat{D}$ is the portfolio of domestic institutional securities

that optimally mimics the domestic market portfolio, with excess return:

$$r_{\hat{D}} = \frac{M_I}{M_D} r_I + \frac{M_R}{M_D} \hat{r}_R. \quad (2)$$

Because of its limited choice set, the institutional investor holds $\hat{D}$ rather than the full market portfolio D . The attainable foreign market portfolio $\hat{F}$ is defined analogously.

Because the institutional investor trades only foreign institutional securities with foreign retail investors, international risk-sharing depends on how well foreign institutional securities can substitute for the attainable domestic market portfolio $\hat{D}$. To measure this substitutability, $r_{\hat{D}}$ can be decomposed into a component spanned by foreign institutional securities and a residual. The resulting mimicking portfolio, the *substitute portfolio* D^s , has return:

$$r_{D^s} = B_{\hat{D}I^*} \mathbf{r}_{I^*}, \quad (3)$$

where $B_{\hat{D}I^*} = V_{\hat{D}I^*} V_{I^*I^*}^{-1}$. In equilibrium, the institutional investor sells the substitute portfolio to foreign retail investors, who use it as the closest substitute for domestic investment.

The following proposition characterizes equilibrium risk premia for domestic securities.⁹

PROPOSITION 1 (Equilibrium Risk Premia): *Suppose Assumptions 1-6 hold. The equilibrium risk premium of any domestic security j is:*

$$\mu_j = \underbrace{\gamma M_W \text{Cov}(\hat{r}_j, f^{\hat{W}})}_{\text{attainable world premium}} + \underbrace{\frac{\gamma^i}{\gamma^f} \gamma M_D \text{Cov}(\hat{r}_j, r_{\hat{D}} - r_{D^s})}_{\text{institutional local premium}} + \underbrace{\gamma^d M_R \text{Cov}(r_j, r_R - \hat{r}_R)}_{\text{retail local premium}}. \quad (4)$$

In particular,

⁹Internet Appendix A derives the equilibrium risk premia and portfolio holdings.

(i) the risk premium of any domestic institutional security I_j is:

$$\mu_{I_j} = \underbrace{\gamma M_W \text{Cov}(r_{I_j}, f^{\hat{W}})}_{\text{attainable world premium}} + \underbrace{\frac{\gamma^i}{\gamma^f} \gamma M_D \text{Cov}(r_{I_j}, r_{\hat{D}} - r_{D^s})}_{\text{institutional local premium}} \quad (5)$$

(ii) the risk premium of any domestic retail security R_j is:

$$\mu_{R_j} = \underbrace{\gamma M_W \text{Cov}(\hat{r}_{R_j}, f^{\hat{W}})}_{\text{attainable world premium}} + \underbrace{\frac{\gamma^i}{\gamma^f} \gamma M_D \text{Cov}(\hat{r}_{R_j}, r_{\hat{D}} - r_{D^s})}_{\text{institutional local premium}} + \underbrace{\gamma^d M_R \text{Cov}(r_{R_j}, r_R - \hat{r}_R)}_{\text{retail local premium}}, \quad (6)$$

where γ is the aggregate absolute risk aversion defined in Internet Appendix Eq. (A.27). The attainable world market factor $f^{\hat{W}}$ is

$$f^{\hat{W}} = \frac{M_D}{M_W} r_{\hat{D}} + \frac{M_F}{M_W} r_{\hat{F}}, \quad M_W = M_D + M_F. \quad (7)$$

Domestic securities earn three types of risk premia. First, the *attainable world premium* compensates investors for the covariance between a security's attainable return $\hat{r}_j$ and the attainable world market factor, the component of world market risk shared through institutional investors. Its magnitude is governed by γ , the aggregate absolute risk aversion.¹⁰

Domestic securities also earn two local risk premia. The first, the *institutional local premium*, compensates investors for the covariance between $\hat{r}_j$ and the *institutional local factor*, defined as the return difference between the attainable domestic market portfolio and its substitute portfolio,

$$f^{inst} = r_{\hat{D}} - r_{D^s}. \quad (8)$$

This premium arises because the substitute portfolio does not perfectly span the attainable domestic

¹⁰This aggregate risk aversion increases with the absolute correlation between domestic and foreign institutional securities because the equilibrium is derived from the institutional investor's first-order condition. As institutional securities become more correlated across countries, diversification opportunities shrink, raising effective risk aversion. Internet Appendix A provides a detailed discussion.

market portfolio, so institutions cannot fully hedge domestic exposures using foreign institutional securities. The premium rises with the ratio γ^i/γ^f , the global investor's risk aversion relative to foreign retail. When this ratio is low, the global investor holds a larger share of foreign institutional securities, enabling greater cross-border sharing of domestic attainable risk through institutional intermediation, and the unhedgeable residual borne by domestic investors commands a smaller premium. The attainable world premium and the institutional local premium depend on attainable exposures because cross-border risk-sharing operates through return components accessible to institutional investors; thus, any covariance across countries must pass through their attainable components.

The second local risk premium, the *retail local premium*, compensates investors for the covariance between a security's return and the *retail local factor*, defined as the unattainable component of the domestic retail portfolio,

$$f^{retail} = r_R - \hat{r}_R. \quad (9)$$

This risk premium reflects the component of a security's covariance with the domestic retail portfolio that is unrelated to domestic institutional securities and is borne solely by domestic retail investors. Its magnitude, therefore, depends on their risk aversion, γ^d . It is zero for institutional securities, which are orthogonal to f^{retail} .

The model encompasses existing partial segmentation models as special cases.¹¹ First, if retail investors diversify internationally but only into institutional securities, the model reduces to the case of Alexander et al. (1987), De Jong and De Roon (2005), and Chaieb and Errunza (2007), augmented by an explicit institutional investor. In this case, because all investors are marginal in the institutional segment, no institutional security earns a premium beyond its exposure to world market risk, and the institutional local premium is identically zero.¹² Institutional securities

¹¹Internet Appendix B.1 provides detailed analyses of these special cases.

¹²Specifically, if foreign (domestic) retail investors are not home-biased, the domestic (foreign) institutional local risk premium becomes zero. This is distinct from Lau et al. (2010), who provide a partial-equilibrium asset-pricing argument. In their setting, home bias affects prices by altering the relative importance of local versus global risk in domestic investors' first-order conditions. In contrast, this paper develops an equilibrium model with heterogeneous

are therefore priced the same way as they would be in the absence of segmentation.¹³ Retail local risk premia remain as long as the institutional choice set is constrained. Second, when institutional investors fully diversify, the market structure resembles that of Greenwood et al. (2018). In this case, the retail local premium is zero because risk-sharing between institutional and retail investors is perfect.

1.3. Connecting Model to Data

The general pricing result in Eq. (4) can be expressed in the following univariate beta representation, which serves as the basis of the empirical analysis:

$$\mu_j = \tilde{\beta}_j^{\hat{W}} \mu^{\hat{W}} + \tilde{\beta}_j^{inst} \mu^{inst} + \beta_j^{retail} \mu^{retail}, \quad (10)$$

where the raw beta β_j^{retail} is defined based on securities' raw returns in Internet Appendix Eq. (A.41), and the attainable betas $\tilde{\beta}_j^{\hat{W}}$ and $\tilde{\beta}_j^{inst}$ are defined based on attainable returns in Internet Appendix Eq. (A.39) and Eq. (A.40). The attainable world premium $\mu^{\hat{W}}$, the institutional local premium μ^{inst} and the retail local premium μ^{retail} are defined as follows:

$$\mu^{\hat{W}} = \gamma M_W \text{Var}(f^{\hat{W}}) \quad (11)$$

$$\mu^{inst} = \frac{\gamma^i}{\gamma^f} \gamma M_D \text{Var}(f^{inst}) \quad (12)$$

$$\mu^{retail} = \gamma^d M_R \text{Var}(f^{retail}). \quad (13)$$

The model provides a framework for understanding three channels through which institutional investment shapes risk premia in international equity markets. The first channel is the *coverage* of the institutional investment scope, which determines how exposures to global and local factors

investors, in which the pricing of domestic securities depends on which investors are marginal.

¹³Equation (2) of De Jong and De Roon (2005) decomposes the world market risk premium for an investable security into covariance terms associated with investable and non-investable market portfolios. Their sum remains the world market premium, so investable assets are priced the same way as under full market integration. The non-investable component of the world market premium is not a distinct local risk premium arising from imperfect risk-sharing.

affect individual securities' risk premia. There are two distinctions in the pricing of retail and institutional securities. Most importantly, retail securities earn an additional retail local risk premium for their direct exposure to the retail local factor. In the model, institutional securities are orthogonal to the retail local factor and hence do not earn this extra local premium.

A second, more technical distinction concerns which exposure to the world and institutional local factors is priced, and it is worth explaining because the empirical analysis uses raw betas. For institutional securities the raw and attainable betas coincide, as their returns are attainable. For retail securities only the attainable exposure is priced since both factors contain a foreign institutional component, and cross-border risk-sharing operates through attainable exposures. The difference between attainable beta and raw beta is likely to be small for retail stocks. The unattainable return of a retail security is orthogonal to the domestic institutional component of both factors by construction, so the attainable and raw betas coincide whenever the unattainable component is also orthogonal to foreign institutional returns. This is plausible if it captures country-specific exposures not spanned by foreign institutional securities. Therefore, I use raw betas as a proxy for attainable betas in the empirical analysis.

The two other channels operate through the market-level institutional local premium, which depends on the risk-bearing *capacity* of global institutional investors and the *correlation* (or substitutability) between local and foreign institutional securities. Substituting in the aggregate risk aversion, $\gamma = \frac{\mu^{\hat{W}}}{M_W \text{Var}(f^{\hat{W}})}$, from the attainable world premium in Eq. (11), the institutional local premium in Eq. (12) can be expressed relative to the world premium as:

$$\mu^{inst} = \frac{\gamma^i}{\gamma^f} \frac{M_D}{M_W} \times \text{Var}(f^{inst}) \times \frac{\mu^{\hat{W}}}{\text{Var}(f^{\hat{W}})}. \quad (14)$$

The first term in Eq. (14), $\frac{\gamma^i}{\gamma^f} \frac{M_D}{M_W}$, can be calculated from observable quantities under simplifying assumptions. Suppose that all investors have the same coefficient of relative risk aversion, α . Let W^k denote the wealth of investor k ; then absolute risk aversion is inversely proportional to

wealth: $\gamma^i = \frac{\alpha}{W^i}$ and $\gamma^f = \frac{\alpha}{W^f}$. In markets that are small relative to the world market, this ratio is approximately:

$$\frac{\gamma^i}{\gamma^f} \frac{M_D}{M_W} = \frac{W^f}{W^i} \frac{M_D}{M_W} \approx \frac{M_W - W^i}{W^i} \frac{M_D}{M_W} = \left(\frac{M_D}{W^i} - \frac{M_D}{M_W} \right) \approx \frac{M_D}{W^i}. \quad (15)$$

The first approximation holds when the total wealth of domestic retail investors is very small relative to the world market capitalization ($W^d \ll M_W$), so that foreign investor wealth is approximately the difference between the world market capitalization and the total wealth of the institutional investor, $W^f \approx M_W - W^i$. The second approximation holds when the total wealth of global institutional investors who participate in the domestic market is very small relative to the world market capitalization ($W^i \ll M_W$). Both conditions hold for every market in the sample except the United States, whose share of the world market capitalization is large.¹⁴

The second term in Eq. (14) is the variance of the institutional local factor, which is also the unspanned variance of the attainable domestic market return, $Var(f^{inst}) = (1 - \rho^2)Var(r_{\hat{D}})$, where ρ denotes the correlation between the attainable domestic market portfolio and its substitute portfolio. I refer to $1 - \rho^2$ as the *unspanned variance share*.

Substituting these two terms in Eq. (14) yields the following decomposition of the institutional local premium:

$$\mu^{inst} = \frac{1}{Capacity} \times (1 - \rho^2)Var(r_{\hat{D}}) \times \frac{\mu^{\hat{W}}}{Var(f^{\hat{W}})}, \quad (16)$$

where $Capacity = \frac{W^i}{M_D}$ is the ratio of the wealth of global institutional investors to domestic market capitalization. The attainable world premium is the same across all markets, so Eq. (16) speaks to the level of the institutional local premium relative to the world premium. In the estimation I also impose the restriction of a common world premium (Section 2.4).

Two channels determine this relative level: the capacity of global institutional investors in the market and the correlation between local and foreign institutional securities. Higher institutional

¹⁴The United States accounts for 53% of the estimation sample's world market capitalization in December 2024.

capacity lowers the institutional local premium by reducing the price of risk, and higher correlation lowers it by shrinking the unspanned variance share, $1 - \rho^2$. Capacity can be measured directly from holdings data, and ρ can be estimated from returns.

In summary, global institutional investment shapes global versus local pricing through three channels. The *coverage* of the institutional investment scope determines which securities earn which risk premium. The market-level institutional local risk premium depends on the risk-bearing *capacity* of global institutional investors through the price of risk, and on the *correlation* (or substitutability) between domestic and foreign institutional securities through the unspanned variance share.

These three channels guide the subsequent empirical analysis. First, unconditionally, the institutional local premium should be high in markets where the capacity of participating global institutions is low and the unspanned variance share is high; the retail local premium should be high where institutional investment has limited coverage. Second, although the model is static, its decomposition of the total local premium into institutional and retail components has time-series implications. In particular, the institutional local premium should rise when institutional investors' risk-bearing capacity weakens, such as during periods of heightened financial constraints.¹⁵ Third, at the security level, stocks with higher institutional ownership should be less exposed to the retail local factor. Institutional investment should therefore affect the composition of the total local risk premium, with implications for the cost of capital.¹⁶

2. Empirical Framework

This section describes the data, explains how the coverage of global institutional investment is measured from active global open-end mutual fund holdings, and presents the construction of the factors and the econometric framework. Throughout the empirical analysis, bold font is omitted when the distinction between vectors and scalars is unambiguous from the context.

¹⁵Akbari et al. (2022) document instances when global institutions have less capacity to hold global securities due to tightening financial constraints.

¹⁶Internet Appendix B.2 discusses how a stock's risk premium changes in the model after it is first covered by institutional investment.

2.1. Data

Stock Universe The sample includes primary stock listings on major exchanges across 21 developed economies and 20 emerging economies covered by CRSP/Compustat Global between January 2003 and December 2024.¹⁷ All stocks are used in factor construction. To isolate the investor underdiversification channel from the security investability channel, I restrict test assets to 66,425 local stocks, defined as securities listed in their country of domicile whose underlying firm has no cross-listings on other open exchanges, classified following Karolyi and Wu (2018). Appendix A.2 describes the filters and corrections applied. Monthly U.S. dollar stock returns and lagged stock-level characteristics are constructed following Jensen et al. (2023).

Holdings and Stock-level Ownership Quarterly holdings data are sourced from FactSet Ownership. FactSet collects stock-level holdings at both the institution and fund levels from regulatory filings, including 13F filings, as well as direct reports from funds and other ownership disclosures. FactSet’s holdings data are less comprehensive outside the United States and the United Kingdom (Kojien et al., 2024). This limitation is unlikely to bind here: classifying and ranking stocks by their global institutional coverage requires a representative set of global institutions rather than the entire population, and the large global institutions that file regulatory reports collected by FactSet provide such a set. Following Bartram et al. (2015), in each quarter, an institution or fund is classified as global if its maximum country allocation does not exceed 90% and its maximum region allocation is below 80%. Security-level global institutional ownership aggregated across investors is calculated following Ferreira and Matos (2008).

Merging FactSet is merged with the CRSP/Compustat Global stock universe through standard identifiers (CUSIP, ISIN, SEDOL). Stocks whose identifiers cannot be linked to FactSet are excluded; for linkable stocks with no reported holdings, ownership and coverage (Section 2.2) are set to zero.

¹⁷Starting from all developed and emerging markets included in the MSCI All Country World Index (ACWI) as of 2024, a market enters the sample once it has more than twenty stocks and at least 20% of its total market capitalization covered by FactSet.

FactSet-derived variables are lagged to reflect the most recently available holdings.¹⁸

2.2. Measuring Coverage

Estimating the model requires classifying stocks as institutional or retail based on the coverage of global institutional investment. Because the representative global institutional investor aggregates many global institutions, I start by measuring the choice set of each constituent investor and then aggregate these choice sets to obtain a stock-level coverage measure.

I measure the choice sets using holdings at the mutual-fund level. This is because both an investor's global diversification and its choice set need to be measured at a level at which portfolio allocation decisions are made. Mutual funds are one such level: each fund's portfolio is allocated by its own manager. In comparison, an institution's holdings aggregate those of its affiliated managers, which may be global or local, active or passive, and so mask the differences in choice sets across managers. Fund-level data thus identify global investors and their choice sets more cleanly at the level of portfolio decision-making.

Specifically, because institutions in the model are global mean-variance investors, I focus on active global open-end mutual funds. These funds are identified from the open-end mutual funds in FactSet that can be mapped to Morningstar.¹⁹

Each quarter t , I follow Koijen and Yogo (2019) and define the choice set of fund f , C_t^f , as the set of stocks that are currently held or ever held in the previous 11 quarters. Because the choice set requires 12 quarters of holdings history and FactSet holdings begin in 2000, the choice set is first available in 2002Q4. Since stock returns lead classification by one month, the estimation sample starts in January 2003. Stock-level coverage is then the total assets under management

¹⁸For quarter-end months, values from the current quarter-end are used; for other months, values from the previous quarter-end are used. Like the stock-level characteristics, they are lagged by one month relative to returns.

¹⁹FactSet and Morningstar are mapped at the share class level by first matching CUSIP/ISIN and then applying fuzzy name matching. Averaged over the sample period, 89% of the total assets under management (AUM) of open-end funds covered by FactSet is matched to Morningstar, 78% to Morningstar equity funds, and 18% to active global funds according to the Morningstar classification (Internet Appendix Table IA.2). Global mutual funds are those whose declared investment scope is global according to their Morningstar Global Category and whose holdings are not concentrated in one country or region, following Bartram et al. (2015). Active funds are those not classified as index funds by Morningstar.

(AUM) at the previous quarter-end of active global funds whose choice set includes the stock, as a share of the AUM of all active global funds participating in the stock's market, that is, funds whose choice set includes at least one stock from the set of stocks listed in market c , $\mathcal{S}_c$:

$$Coverage_{i,c,t} = \frac{\sum_f AUM_{f,t-1} \mathbf{1}\{i \in C_t^f\}}{\sum_f AUM_{f,t-1} \mathbf{1}\{C_t^f \cap \mathcal{S}_c \neq \emptyset\}}. \quad (17)$$

A stock is institutional in a given quarter if its coverage is positive, that is, if it lies in the choice set of at least one active global fund, and retail otherwise. To ensure that zero coverage is not mechanical for newly listed stocks, a stock is classified only after it has been observable and linkable to FactSet for at least twelve quarters. The average proportion of covered stocks (over time and then across countries) is 58% in developed markets and 44% in emerging markets.²⁰ Around half of the stocks have zero coverage, which is consistent with the motivating evidence presented in Figure 1 that the median global institutional ownership is close to zero across many markets.

For classification and factor construction, the boundary between institutional and retail stocks needs to be exact. Coverage provides an operational boundary based on the revealed choice sets of the observed active global-fund universe, whereas ownership sums positions held by all institutions classified as global, which can be positive without the stock being in the choice set of any global fund. Since coverage is defined over the past 12 quarters of holdings, it is highly persistent (Internet Appendix Table IA.4), and does not move with the size of positions or recent returns. Global institutional ownership, on the other hand, is available for more stocks and ranks them within each market as much as coverage does.²¹ It is therefore used as the stock-specific instrument for factor loadings in the conditional estimation, which requires only the relative ranking of stocks by global institutional coverage, not the boundary.

I provide two validation tests for the coverage measure in Internet Appendix C. First, I show

²⁰Internet Appendix Table IA.3 reports these statistics for individual markets.

²¹Within each market and quarter, the Spearman rank correlation between coverage and global institutional ownership averages 0.76 across the 41 sample markets (0.73 in developed and 0.79 in emerging markets).

that coverage changes around rule-based index revisions: it increases significantly after inclusion in the MSCI All Country World Index (ACWI) and decreases significantly after exclusion. Second, I compare coverage with two alternative indicators of scope: ACWI membership and investability defined following Karolyi and Wu (2018). Index membership is narrow because active funds also invest outside the benchmark, while investability is coarse because it is determined by the availability of cross-listings and so does not vary among stocks within an open market. Coverage predicts entry by global funds and global institutional ownership better than both indicators. These tests support coverage as a measure of global institutional investment scope.

2.3. Factor Construction

The three factors are constructed based on their model definition, which requires projection on the span of domestic and foreign institutional securities, each approximated by a small set of basis portfolios. Following Karolyi and Wu (2018), I use three factors (market, size, value) as basis portfolios for the span of domestic institutional securities. For the span of foreign institutional securities, I use six regional portfolios: North America; Developed Europe; Developed Asia Pacific; Latin America; Emerging Asia; and Emerging Europe, the Middle East, and Africa (EMEA).

The attainable domestic market return $r_{\hat{D}}$ is the fitted value from a time-series regression of the domestic market return onto the three domestic institutional factors. The attainable world market factor is the value-weighted portfolio of these attainable market returns. It is almost perfectly correlated with the return on the world market portfolio, with a sample correlation of 99.9%.

The institutional local factor f^{inst} is the residual from a time-series regression of $r_{\hat{D}}$ onto the regional foreign institutional portfolios. The retail local factor f^{retail} is the residual from a time-series regression of the retail market portfolio onto the three domestic institutional factors. All regressions include a constant to preserve the covariance form of the slope coefficients in the model. Each factor is the sum of the intercept and the residual, so it retains its mean.

Table 1 reports, for each market, the annualized average returns and volatilities of the institu-

tional and retail local factors. Table 2 reports the full-sample R^2 for the projections and correlation among factors. The R^2 of the retail projection ranges between 50% and 90% across most markets. The regional foreign institutional portfolios span $r_{\hat{D}}$ well in developed markets, with an average R^2 exceeding 70%, and less well in emerging markets, where the average R^2 is approximately 50%. In most markets, the correlations among the factors are close to zero.

TABLE 1 AND TABLE 2 ABOUT HERE

2.4. Econometric Framework

I estimate the model using individual stock returns to preserve the within-market variation in global institutional coverage that identifies the local premia. I use the two-pass regression method developed by Gagliardini et al. (2016) and adapted for international markets by Chaieb et al. (2021). This approach is well-suited to individual stocks, as it addresses the errors-in-variables problem inherent in the first-pass estimation and accommodates unbalanced panels. It also allows for time-varying risk premia using conditional information. For each country c , n_c denotes the number of stocks, T_c the length of the time series, and K the number of factors.

The estimation requires two conditions. The first condition disciplines the factor structure of returns. Raw excess returns must admit a linear factor approximation:

$$r_{i,t} = \alpha_{i,t} + \beta'_{i,t} f_{c,t} + \varepsilon_{i,t}, \quad (18)$$

where $E_{t-1}[\varepsilon_{i,t}] = 0$ and $\text{Cov}_{t-1}(\varepsilon_{i,t}, f_{c,t}) = 0$. Consistent estimation of the risk premia and valid inference require that the residuals contain no remaining factor structure. The factor loadings $\beta_{i,t}$ need not coincide with structural betas from the model. This condition can be assessed using the diagnostic statistic of Gagliardini et al. (2019). A negative diagnostic indicates the absence of remaining factor structure. When presenting the estimation results, markets with negative diagnostics are highlighted in bold.

The second condition is the theoretical prediction for the equilibrium risk premia, which must be linear in the same factor exposures as in the linear factor representation:

$$E_{t-1}[r_{i,t}] = \beta'_{i,t} \mu_{c,t}, \quad (19)$$

where $\mu_{c,t} = [\mu_{c,t}^{\hat{W}}, \mu_{c,t}^{inst}, \mu_{c,t}^{retail}]'$ is the vector of conditional factor risk premia. Eq. (19) is based on raw beta exposures from multivariate time-series regressions in Eq. (18), which deviates from the theoretical pricing relation in Eq. (10) in two respects. First, as discussed in Section 1.3, in the model, only the attainable exposures to the world and the institutional local factors are priced for retail securities, and raw betas are exact if their unattainable return components are orthogonal to foreign institutional returns. This approximation is supported empirically by the weak correlations between the retail local factor, a value-weighted portfolio of the unattainable components of individual stocks, and these two factors.

Second, in the model, the risk premia in Eq. (10) are defined with respect to univariate betas, while the multivariate risk premia in Eq. (19) absorb factor covariances. Table 2 shows that these correlations remain close to zero in most markets, suggesting that the multivariate risk premia should closely approximate their univariate structural counterparts in Eq. (11)-Eq. (13).

Combining Eq. (19) and Eq. (18) yields the following asset-pricing restriction:

$$\alpha_{i,t} = \beta'_{i,t} \nu_{c,t}, \quad \nu_{c,t} = \mu_{c,t} - E_{t-1}[f_{c,t}], \quad (20)$$

where $\nu_{c,t}$ captures the wedge between factor risk premia and conditional factor expectations. The wedge is zero when factors are tradable; in practice, $\nu_{c,t}$ need not be zero due to transaction costs involved in rebalancing and short-selling (Gagliardini et al., 2016). Among the three factors, transaction costs are likely minimal for the attainable world market factor, while non-negligible for the two local factors constructed from long-short positions. Therefore in the estimation, I impose the restriction $\nu_W = 0$, while allowing the wedge to be non-zero for the two local factors.

Conditioning information enters the estimation through the dependence of factor risk premia,

conditional factor expectations, and factor loadings on lagged instruments. At the country level, both factor risk premia and conditional expectations in country c are linear in p lagged common instruments $Z_{c,t-1}$:

$$\mu_{c,t} = \Lambda_c Z_{c,t-1} \quad (21)$$

$$E_{t-1}[f_{c,t}] = F_c Z_{c,t-1} \quad (22)$$

$$\nu_{c,t} = (\Lambda_c - F_c) Z_{c,t-1}, \quad (23)$$

where Λ_c and F_c are $K \times p$ matrices, with $K = 3$ in the baseline model. This implies that the wedge between the two is also linear in the instruments, with the world row of $\Lambda_c - F_c$ equal to zero under the restriction $\nu_W = 0$.

Stock-level factor exposures are additionally linear in q stock-specific instruments $Z_{i,t-1}$:

$$\beta_{i,t} = B_i Z_{c,t-1} + C_i Z_{i,t-1}, \quad (24)$$

where B_i is a $K \times p$ matrix and C_i is a $K \times q$ matrix.

Under the asset-pricing restriction in Eq. (20), the linear specifications in Eq. (22)-Eq. (24) imply that the intercept and the factor component of the time-series regression in Eq. (18) take the form:

$$\alpha_{i,t} = Z'_{c,t-1} B'_i (\Lambda_c - F_c) Z_{c,t-1} + Z'_{i,t-1} C'_i (\Lambda_c - F_c) Z_{c,t-1} \quad (25)$$

$$\beta'_{i,t} f_{c,t} = Z'_{c,t-1} B'_i f_{c,t} + Z'_{i,t-1} C'_i f_{c,t}. \quad (26)$$

The resulting intercept in Eq. (25) is a quadratic function of the conditioning instruments, reflecting the interaction between time variation in factor exposures and time variation in risk premia. The factor component in Eq. (26) consists of factors scaled by the same instruments. Collecting these terms, the time-series regression in Eq. (18) becomes:

$$r_{i,t} = b'_{1,i}x_{1,i,t} + b'_{2,i}x_{2,i,t} + \varepsilon_{i,t}. \quad (27)$$

Here, $x_{1,i,t}$ collects all interaction terms from Eq. (25), and $x_{2,i,t}$ contains the scaled factor terms from Eq. (26). Detailed definitions of the regressors are provided in Appendix B.1. Stock-level time-varying factor exposures are then recovered from the first-pass estimates.

I adopt the trimming condition of Gagliardini et al. (2016), retaining stocks having at least 60 months of observations and a first-pass condition number no greater than 25, and extend the instrument selection procedure in Chaieb et al. (2021) to eliminate numerically unstable common and stock-specific instruments. Details of the instrument-selection procedure are provided in Appendix B.2.

I consider two specifications. Unconditional estimation uses only a constant as the instrument. Conditional estimation uses two dividend-yield spreads as instruments: the difference in the dividend yields of the domestic and world market portfolios DY^{D-W} and the difference in the dividend yields between the retail and the institutional portfolios DY^{R-I} , which reflect the valuation differences in the long and short legs of the institutional local and the retail local factors, respectively. The domestic and world market portfolios closely track the two legs of the institutional local factor, the attainable domestic market and the substitute portfolio of foreign institutional securities; the retail and institutional portfolios likewise track the legs of the retail local factor. Both instruments are standardized to have a zero mean and unit variance. For parsimony, I restrict the pricing wedge of each local factor to load only on a constant and its matched dividend-yield spread, while its conditional expectation depends on both spreads. The stock-specific instrument is the within-market percentile rank of global institutional ownership, which ranks stocks by global institutional coverage (Section 2.2).

The second pass runs a cross-sectional weighted least squares (WLS) regression on first-pass coefficients to estimate the pricing wedge $v_c = \text{vec}(\Lambda'_c - F'_c)$. F_c is estimated separately from a

seemingly unrelated regression (SUR) where local factors are predicted by the two dividend-yield spreads and the world factor is predicted by a third instrument, the world dividend yield.²² The time-varying risk premia at the country-level are computed as $\hat{\mu}_{c,t} = \hat{\Lambda}_c Z_{c,t-1}$, and time-varying stock-level risk premia are calculated based on estimated factor exposures from the first pass. Appendix B provides further details on estimation and inference.

3. Empirical Results

Section 3.1 reports unconditional estimates of the institutional and retail local premia across developed and emerging markets and relates the cross-market variation in the institutional local premium to the model’s capacity and correlation channels. Section 3.2 presents conditional estimates and time-varying risk premia. Section 3.3 analyzes how institutional ownership shapes risk premia at the individual-stock level.

3.1. Unconditional Estimation

Table 3 reports the annualized unconditional institutional local and retail local risk premia for each market, along with their standard errors. It also reports the number of stocks used in the second-pass cross-sectional regression, which is substantially smaller than the total number of stocks reported in Table 1, reflecting the trimming condition. Markets with a negative Gagliardini et al. (2019) diagnostic, indicating no remaining factor structure in the residuals, are highlighted in bold. Under the restriction that the pricing wedge of the attainable world factor is zero (Section 2.4), the unconditional world premium is the sample mean of the attainable world factor over each market’s estimation window. While the discussion below focuses on the two local premia, the last row in Table 3 reports, for reference, the unconditional world premium over the full sample and its standard error.

Both institutional and retail local premia are economically meaningful across developed

²²Under the restriction $\nu_W = 0$, the world premium equals the conditional expectation of the world factor, whose specification is common across markets; estimated world premia differ across markets only through differences in the available estimation window.

and emerging markets. In developed markets, the institutional and retail local premia average 2.2% and 2.5% respectively, while in emerging markets the corresponding averages are 3.5% and 4.5%. Relative to the unconditional world premium of 9.0%, both local components contribute considerably to total risk premia.²³ The standard error of each cross-market average is the Newey-West standard error of the cross-market average of the corresponding factor series, which accounts for the correlation of factor returns across markets (Appendix B). All four averages are significantly positive at the 5% level.

At the individual-market level, the institutional local premium is positive and significant at the 5% level (one-sided) in 8 of the 21 developed and 7 of the 20 emerging markets, and the retail local premium in 8 developed and 10 emerging markets.

TABLE 3 ABOUT HERE

The validity of the preceding inference requires that the proposed factors adequately capture the cross-sectional dependence in returns, as indicated by a negative Gagliardini et al. (2019) diagnostic. In the baseline specification, the diagnostic is negative in 20 out of 41 markets. This proportion is lower than that reported in Chaieb et al. (2021) for their specification with a world and an excess country market factor, likely reflecting their longer samples for most markets. The diagnostic is a model selection rule rather than a formal statistical test: if the factor model is correctly specified, the probability of a negative value approaches one as the cross-sectional and time-series dimensions grow (Gagliardini et al., 2019). Therefore, a positive value may reflect finite-sample limitations, particularly in markets with a shorter sample.

The central prediction of the model is not that the local premia are universally important, but that their levels are governed by the capacity, correlation, and coverage of global institutional

²³The closest benchmark to compare these levels is Chaieb et al. (2021), who estimate global and local risk premia from international individual stocks using the same Gagliardini et al. (2016) two-pass methodology. The sum of the institutional and retail local premia estimated here is broadly consistent in magnitude with their estimates of the excess country market premium. Direct comparison cannot be made because they use a much earlier sample period, do not impose the theoretical restriction that the world premium is the same across markets, and estimate a single local premium on a reduced-form excess country market factor. Other studies (e.g., Fama and French, 2012; Karolyi and Wu, 2018) estimate local premia using regional or country-level portfolio test assets rather than individual stocks, making direct comparison difficult.

investment, as discussed in Section 1.3. I now examine whether the cross-market variation in the unconditional institutional local premium aligns with the capacity and correlation channels, deferring the coverage channel to the conditional estimation, where individual stocks' factor exposures are instrumented by the percentile rank of their global institutional ownership. Eq. (16) decomposes the institutional local premium into its capacity (W^i/M_D) and unspanned variance share ($1 - \rho^2$) components, repeated here for convenience:

$$\mu^{inst} = \frac{1}{Capacity} \times (1 - \rho^2) Var(r_{\hat{D}}) \times \frac{\mu^{\hat{W}}}{Var(f^{\hat{W}})}.$$

Since the world premium is restricted to be the same across markets in the estimation, the level of the estimated institutional local premium can be compared directly across markets.

Figure 2 plots the annualized unconditional institutional local risk premium for each market against capacity and unspanned variance share from the decomposition. Since the decomposition is multiplicative, I take the logarithm of each component. Panel A relates the institutional local risk premium to capacity, measured as the time-series average of the log ratio of the total AUM of active global open-end mutual funds participating in the market to its total market capitalization. Panel B plots the institutional local premium against the logarithm of the unspanned variance share, measured as one minus the squared full-sample return correlation between the attainable domestic market portfolio $r_{\hat{D}}$ and its substitute portfolio r_{D^s} . Consistent with the model, the institutional local premium is negatively associated with capacity and positively associated with the unspanned variance share. I also run a cross-market regression of the institutional local risk premium on these two components, controlling for the logarithm of the variance of the attainable domestic market portfolio; the coefficients are reported below the figure. The coefficient on capacity is negative and the coefficient on the unspanned variance share is positive, both significant at the 5% level (one-sided).

FIGURE 2 ABOUT HERE

A natural concern is that the positive diagnostics reflect omitted local factor structure. Institutional investors are known to prefer large stocks (Ferreira and Matos, 2008), so the pricing of the institutional and retail local factors may partially capture the size effect. To assess this concern and to control for additional sources of local risk, I re-estimate the unconditional asset-pricing model in Eq. (10) including three standard local factors: small-minus-big (SMB), high-minus-low (HML), and momentum (MOM), constructed following Jensen et al. (2023).

TABLE 4 ABOUT HERE

The risk premium estimates controlling for these additional local factors are reported in Table 4. The Gagliardini et al. (2019) diagnostic is now positive in only 4 markets, indicating that these factors capture local factor structure not fully spanned by the baseline. Importantly, the institutional and retail local premia remain close to the baseline estimates across most markets: the correlation between the two specifications is 0.87 for both premia, and the cross-market averages are similar. The institutional local premium remains positive and significant in 10 developed and 9 emerging markets, and the retail local premium in 6 developed and 12 emerging markets. The omission of these additional local factors in the baseline estimation therefore does not materially affect the estimated local premia. Once both model-implied local factors are present, the SMB and HML premia are positive and significant in only 4 and 2 markets, whereas the MOM premium remains positive and significant in a majority of markets. These results indicate that the baseline results are not driven by the size effect; if anything, they are consistent with heterogeneous exposure to international risk-sharing through global institutional ownership contributing to the size premium, although the strong correlation between institutional ownership and size makes the two channels difficult to separate.

3.2. Conditional Estimation and Time-variation

The conditional estimation serves two goals: to examine the coverage channel through the stock-specific instrument, and to trace out the dynamics of the risk premia descriptively.²⁴ The stock-specific instrument, the within-market percentile rank of global institutional ownership, captures within-stock variation in factor exposures over time. The common instruments, the dividend-yield spreads, are reduced-form statistical predictors of time-varying local risk premia; the variation they capture may reflect underlying economic primitives that the estimation is not designed to identify.

Table 5 reports the estimated coefficients governing the conditional risk premia. Since the dividend-yield spreads are standardized to have zero mean and unit variance, the coefficient on the constant in Eq. (21) equals the time-series average of the conditional risk premium, and the remaining coefficients capture the response of the premium to a one-standard-deviation change in the corresponding spread. The time-series averages are broadly consistent with the unconditional estimates in Table 3, computed excluding Brazil for comparability. The retail local premium averages 3.1% in developed and 3.7% in emerging markets, compared with 2.5% and 3.9% unconditionally, and the institutional local premium averages 3.4% in emerging markets, compared with 3.7% unconditionally. The cross-market correlations between the conditional time-series averages and the unconditional estimates are 0.73 for the institutional and 0.41 for the retail local premium. The exception is the institutional local premium in developed markets, whose average falls from 2.2% to -0.3% .

TABLE 5 ABOUT HERE

The lower cross-market correlation of the retail local premium partly reflects sample selection. The conditional specification imposes stricter requirements on time-series length because additional instruments and their interaction terms increase the number of regressors, reducing the

²⁴Brazil is excluded from the conditional estimation because erroneous dividend data in Compustat Global prevent construction of its retail-institutional dividend-yield spread; the conditional sample therefore covers 40 markets.

set of stocks used for estimation. Because retail stocks tend to be younger, they are more likely to be excluded, so the sample underlying the retail local premium changes more between the two specifications, moving individual-market estimates while leaving the cross-market averages largely unchanged.

The divergence in the developed-market institutional local premium may instead reflect that unconditional estimates abstract from time-variation in both risk premia and factor exposures. When betas vary over the business cycle, average conditional risk premia need not coincide with unconditional estimates, a manifestation of the market-timing bias documented in Jagannathan and Wang (1996), Lewellen and Nagel (2006), and Boguth et al. (2011).

The coverage channel predicts that as a stock's global institutional coverage rises, its exposure to the retail local factor falls (Section 1.3). Internet Appendix Table IA.6 reports, for each market, the number of stocks for which stock-level global institutional ownership is selected as an instrument, together with the median coefficient on this variable in the time-varying beta specification in Eq. (24). This coefficient captures how changes in a stock's institutional-ownership percentile rank shift its factor loadings over time. Institutional ownership is selected for a large number of stocks in each market, indicating that it helps explain time-variation in factor exposures. Consistent with the coverage channel, the median coefficient for the retail local factor is negative in 29 of the 40 markets, while the median coefficients for the other factors display no clear cross-market pattern. This evidence reflects within-stock time-series associations between ownership and exposure; Section 3.3 examines the cross-sectional counterpart.

Figure 3 plots the time-series evolution of the equal-weighted cross-country averages of the conditional risk premia in developed and emerging markets. The dynamics differ substantially across market types and across risk premium components. Local risk premia account for a larger share of the total risk premium in emerging markets. In emerging markets, the institutional local premium has trended down, although it increased substantially in the early 2020s. In developed markets, the institutional local premium is close to zero over the sample period, while the retail local premium is positive through most of the sample with a slight downward trend.

FIGURE 3 ABOUT HERE

The three risk premium components also exhibit different business-cycle patterns. The attainable world and the emerging-market institutional local premia are countercyclical, spiking during major distress episodes such as the Global Financial Crisis and the Covid-19 crisis. This pattern is consistent with existing evidence that institutional investors, particularly foreign institutions, scale back their equity positions, especially in emerging markets, during periods of global stress (Kacperczyk et al., 2025; Glossner et al., 2025), reflecting tighter financial constraints and reduced risk-bearing capacity (Akbari et al., 2022).

By contrast, the retail local premium in developed markets is procyclical, declining during major distress episodes. This pattern does not indicate an improvement in retail investors' risk-bearing capacity during downturns. Instead, rising within-market correlations during crises imply that a larger share of domestic retail return variance is spanned by the domestic institutional basis, compressing the unattainable component priced by the retail local factor. The retail local premium exhibits no obvious cyclicalities in emerging markets. Since the figure reports equal-weighted averages across countries, the lack of clear cyclicalities in the emerging-market average may reflect that the unspanned retail risk and retail risk-bearing capacity are local and not synchronized across these markets, consistent with the model implication that this risk premium is country-specific.

3.3. Institutional Investment and Risk Premia: Stock-level Evidence

This subsection examines how institutional ownership is associated with risk premia across individual stocks within each market, using the stock-level conditional exposures and risk premia from the Gagliardini et al. (2016) estimation in Section 3.2. Within-market cross-sectional variation in stock-level risk premia is linked to institutional ownership through differences in factor exposures. Although the model's classification of stocks into institutional and retail is binary, factor exposures are continuous in the data, as ownership varies continuously across stocks. Stocks with higher institutional ownership need not be empirically orthogonal to the retail local factor, which

is constructed to be orthogonal to the domestic institutional basis portfolios, but they should be less exposed to it to the extent that they resemble those portfolios. Such stocks may also carry greater exposure to the world and institutional local factors, to the extent that they have larger attainable return components. Given market-level factor risk premia, cross-sectional differences in exposures translate directly into cross-sectional differences in stock-level risk premia.

To investigate this cross-sectional association, I estimate panel regressions of the model-implied factor exposures and risk premia on the lagged within-market percentile rank of global institutional ownership (IO^{rank}), the stock-specific instrument used in the conditional estimation, which captures cross-sectional variation in global institutional coverage (Section 2.2). The regressions control for lagged stock characteristics, including the natural logarithm of market capitalization in millions of U.S. dollars ($LOGMV$), the book-to-market ratio (BM), and the transaction cost measure of Fong et al. (2017) (FHT), and include country-time fixed effects ($\alpha_{c,t}$).

$$y_{i,t} = \alpha + \delta_1 IO_{i,t-1}^{rank} + \delta_2 LOGMV_{i,t-1} + \delta_3 BM_{i,t-1} + \delta_4 FHT_{i,t-1} + \alpha_{c,t} + \epsilon_{i,t}, \quad (28)$$

where $y_{i,t}$ denotes the world, institutional local, or retail local factor exposure ($\beta_{i,t}$), the corresponding risk premium ($\mu_{i,t}$), or the total risk premium.

The results are reported in Table 6. Across both developed and emerging markets, institutional ownership is negatively and significantly related to exposure to the retail local factor, and this association translates into a significant reduction in the retail local premium. In contrast, institutional ownership is positively and significantly associated with exposure to the attainable world factor and its risk premium. A move from the 25th to the 75th percentile of the within-market ownership distribution is associated with a lower retail local premium of 0.3 percentage points in developed and 0.6 in emerging markets, and a higher world premium of 1.4 percentage points in developed and 0.5 in emerging markets. Institutional ownership is also positively associated with exposure to the institutional local factor, though the association with its risk premium is significant

only in developed markets. Taken together, higher institutional ownership is associated with a significantly higher total risk premium in developed markets and an insignificantly lower total premium in emerging markets. These associations are identified from cross-sectional variation across stocks within each market, and thus complement the conditional estimation, where the coefficient on the stock-specific instrument captures how a stock's exposures move with its own ownership rank over time (Section 3.2).

TABLE 6 ABOUT HERE

These estimates should be interpreted with caution. The trimming condition of the conditional estimation disproportionately excludes retail-dominated stocks, which tend to have shorter return histories, so the sample underrepresents the stocks with the largest retail local exposures. This selection attenuates the estimated association between institutional ownership and the retail local premium toward zero. The reduction reported in Table 6 is therefore best read as a lower bound in magnitude, and the positive association with the total premium as an upper bound.

4. Concluding Remarks

This paper proposes an investor underdiversification channel for local risk premia in international equity markets. Holdings data reveal two facts. First, retail investors display near-perfect home bias while institutions are globally diversified. Second, institutions substantially underdiversify within each market, consistent with a limited investment scope that may be shaped by investment mandates or benchmarking incentives. Because retail investors are home-biased, international risk-sharing depends on institutional investors' cross-market investment. Institutions' imperfect cross-market hedging generates an institutional local premium that depends on the risk-bearing capacity of institutional investors, while their limited investment scope leaves residual risk, generating a retail local premium that depends on the risk-bearing capacity of retail investors. Institutional investment thus shapes global versus local pricing through three channels:

capacity, correlation, and coverage. Unlike security-level investability, the underdiversification channel operates even when markets are fully open to foreign investors, so local risk premia persist without formal investment barriers.

The model is estimated in 41 economies from January 2003 to December 2024. Constructing the model-implied local factors requires classifying stocks into institutional and retail segments. To identify the boundary of global institutions' investment scope, I construct a stock-level coverage measure based on the revealed choice sets of active global open-end mutual funds, and show that coverage predicts future entry by global funds and global institutional ownership better than alternative indicators of scope, such as index membership and investability.

I estimate the model with the Gagliardini et al. (2016) two-pass regression applied to individual stock returns in each market. The institutional local and the retail local premia are economically meaningful across a wide range of markets, averaging 2.2% and 2.5% across 21 developed markets and 3.5% and 4.5% across 20 emerging markets. Consistent with the capacity and correlation channels, the institutional local premium is lower in markets with greater institutional risk-bearing capacity and lower unspanned variance share. Over time, the attainable world and the emerging-market institutional local premia are countercyclical, while the retail local premium is procyclical in developed markets and shows no clear cyclicity in emerging markets. Consistent with the coverage channel, in the cross-section of stocks within each market, higher institutional ownership is associated with lower exposure to the retail local factor and a lower retail local premium. It is also associated with higher exposure to the attainable world factor and a higher world premium; the total risk premium is significantly higher in developed markets, with no significant association in emerging markets.

The model takes the heterogeneous diversification patterns of institutional and retail investors as given. Identifying the primitives that drive these patterns and their interaction with the pricing channel documented here is a natural direction for future work.

Appendix A. Additional Information on Data

Appendix A.1. Motivating Evidence

The raw home bias (Bekaert and Wang, 2009) of investor group g in country c is defined as the difference between the group's actual domestic portfolio allocation, $W_{g,c,t}^{act}$, and country c 's weight in the world market portfolio, $W_{c,t}^{BM}$,

$$HB_{g,c,t} = W_{g,c,t}^{act} - W_{c,t}^{BM}. \quad (A1)$$

The benchmark weight $W_{c,t}^{BM}$ is the U.S. dollar market capitalization weight of country c , based on CRSP/Compustat Global. Actual domestic allocations of different investor groups are calculated using data from the IMF Coordinated Portfolio Investment Survey (CPIS),²⁵ the OECD National Accounts, and FactSet, as detailed below.

All Residents CPIS records aggregate cross-border portfolio investment by residents of each country. Domestic investment by residents is calculated as the difference between total domestic market capitalization from CRSP/Compustat Global and foreign residents' investment in the domestic market as recorded in CPIS:

$$I_{c,c,t} = M_{c,t} - \sum_{k \neq c} I_{k,c,t}.$$

Here, $M_{c,t}$ is the domestic market capitalization and $I_{k,c,t}$ denotes investment by residents from country k in target market c , as reported by CPIS.

The total investment by all residents of country c is the sum of their domestic and foreign investments. The actual domestic allocation is therefore given by

²⁵The IMF has since rebranded this dataset as the Portfolio Investment Positions by Counterpart Economy dataset.

$$W_{c,t}^{act} = \frac{I_{c,c,t}}{\sum_k I_{c,k,t}}.$$

Global Institutions Global institutions are defined as institutional investors whose maximum country allocation does not exceed 90% and maximum region allocation is below 80%, following Bartram et al. (2015). For global institutions domiciled in country c , the domestic allocation is computed as the ratio of their total domestic equity investment to total assets under management, using holdings data from FactSet.

Households Total equity investment by households is sourced from OECD National Accounts, Annual Financial Balance Sheets (stocks) for the Household sector. Foreign investment by households is measured using CPIS data for the same sector. Domestic allocation is defined as one minus the share of foreign equity investment.

Appendix A.2. Data Construction

Filters and Corrections The stock universe is constructed from Center for Research in Security Prices (CRSP)/Compustat Global data following Jensen et al. (2023). I retain common shares and American Depositary Receipts (ADRs) that are primary listings on major stock exchanges, defined as stock exchanges with the largest number of listings in each country.²⁶ Each security is assigned to its country of listing. ADRs and foreign direct listings therefore enter the factor portfolios of their listing country rather than their domicile country. I further require that the country of domicile information be available in Compustat. Test assets are restricted to local stocks, defined as securities whose country of listing coincides with their country of domicile and whose underlying firm has no cross-listings on other open exchanges, classified following Karolyi and Wu (2018).

²⁶A few economies have multiple major stock exchanges: United Arab Emirates (Dubai and Abu Dhabi), Brazil (Bovespa and Rio de Janeiro), Canada (Toronto and TSX Venture), Switzerland (Swiss Exchange and Zurich), China (Shanghai and Shenzhen), Germany (XETRA and Deutsche Boerse), France (NYSE Euronext Paris and Paris), India (BSE and National Stock Exchange), Japan (Tokyo and Osaka), Korea (Korea Exchange and KOSDAQ), Taiwan (Taiwan Stock Exchange and Taipei Exchange), and United States (NYSE, NYSE Arca, Amex, and Nasdaq).

This restriction removes cross-listed securities from the test asset set while retaining them in factor construction. Two securities are removed due to data outliers (gvkey=222638, iid=90 and gvkey=025903, iid=92).

FactSet Ownership Data FactSet covers global equity ownership by institutional investors such as mutual funds, pension funds, bank trusts, insurance companies, and sovereign wealth funds. For equities listed in the United States, FactSet gathers holdings data from 13F filings and N-30D filings. For non-U.S. listed equities, FactSet collects data from sources like regulatory agencies, stock exchange announcements, mutual fund managers, and company annual reports. The dataset is structured as two main holdings tables: 13F holdings and fund holdings. I follow Ferreira and Matos (2008) and merge holdings from these two sources at the FactSet institution level (buy- or sell-side institutions as defined by FactSet), yielding a panel of institution-security-quarter holdings by institutional investors.

The following data errors are corrected in FactSet. For 13F holdings, I follow Kojen et al. (2024) and exclude two FactSet entity identifiers (0FSVG4-E and 000V4B-E), which contain known errors in comparison with the EDGAR 13F filings. For fund-level holdings, I remove holdings by FactSet fund 04B9J7-E in security C7R70B-S, which clearly contains an outlier.

Appendix B. Estimation and Inference

This Appendix provides details about the two-pass regression method used to estimate risk premia using individual stock returns. The estimation is performed country by country. For country c , n_c denotes the number of stocks and T_c the length of the available time series.

Appendix B.1. Regression framework

The first-pass time-series regression is:

$$r_{i,t} = b'_{1,i}x_{1,i,t} + b'_{2,i}x_{2,i,t} + \epsilon_{i,t} \quad (\text{B1})$$

$$\begin{aligned} x_{1,i,t} &= (\text{vech}(X_t)', Z'_{c,t-1} \otimes Z'_{i,t-1})' \in \mathbb{R}^{d_1=p(p+1)/2+pq} \\ x_{2,i,t} &= [(f'_{c,t} \otimes Z'_{c,t-1}), (f'_{c,t} \otimes Z'_{i,t-1})] \in \mathbb{R}^{d_2=K(p+q)}, \end{aligned} \quad (\text{B2})$$

where K is the number of factors, p is the number of common instruments, and q is the number of stock-specific instruments. $X_t \in \mathbb{R}^{p \times p}$ is a symmetric matrix such that $X_{t,k,l} = Z_{c,t-1,k}^2$ if $k = l$, and $X_{t,k,l} = 2Z_{c,t-1,k}Z_{c,t-1,l}$ otherwise. $x_{1,i,t}$ contains $d_1 = p(p+1)/2 + pq$ terms, among which $\text{vech}(X_t)$ contains $p(p+1)/2$ interaction terms among common instruments and $Z'_{c,t-1} \otimes Z'_{i,t-1}$ contains pq interaction terms between common instruments and stock-specific instruments. $x_{2,i,t}$ contains $K(p+q)$ factors scaled by common and stock-specific instruments.

The second pass regresses $\hat{b}_1$ on $\hat{b}_3$, which is a transformation of $\hat{b}_2$, to estimate the wedge $\nu_c \in \mathbb{R}^{Kp}$:

$$\hat{b}_{1,i} = \hat{b}_{3,i}\nu_c = \hat{b}_{3,i}H_c\theta_c \quad (\text{B3})$$

$$\hat{b}_{3,i} = \left((N_p[B'_i \otimes I_p])', [W_{p,q}(C'_i \otimes I_p)]' \right)' \quad (\text{B4})$$

$$\nu_c = \text{vec}[\Lambda'_c - F'_c] = H_c\theta_c, \quad (\text{B5})$$

where $N_p = \frac{1}{2}D_p^+(W_{p,p} + I_{p^2})$, $W_{p,q}$ is the commutation matrix such that $\text{vec}[A'] = W_{p,q}\text{vec}[A]$ and D_p^+ is the $p(p+1)/2$ -by- p^2 matrix such that $\text{vech}[A] = D_p^+\text{vec}[A]$. ν_c is ordered factor by factor, so its position $(k-1)p + \ell$ is the loading of factor k 's wedge on common instrument ℓ . In the estimation, the world row of $\Lambda_c - F_c$ is set to zero. $\theta_c \in \mathbb{R}^r$ collects the free wedge parameters of the two local factors only, and the restriction is imposed by expanding θ_c to the full-dimensional wedge ν_c through H_c , a $Kp \times r$ matrix formed by keeping the columns of I_{Kp} that correspond to

the r free positions.

Appendix B.2. The instrument selection procedure

Because interaction terms between factor loadings and factor risk premia generate a large number of regressors, some restrictions need to be imposed on B_i and C_i , the coefficients that govern how factor loadings depend on the instruments. I use an instrument-selection algorithm similar to Chaieb et al. (2021).

Let $\mathbb{I}_{B_i}$ and $\mathbb{I}_{C_i}$ be $K \times p$ and $K \times q$ indicator matrices whose elements are equal to one if the corresponding elements in B_i and C_i are non-zero. $\tilde{\mathbb{I}}_{B_i}$ is the p -vector whose j^{th} element is equal to one if at least one element in the j^{th} column of B_i is not zero, $\tilde{\mathbb{I}}_{C_i}$ is the q -vector whose j^{th} element is equal to one if at least one element in the j^{th} column of C_i is not zero. $\tilde{p}_i = \tilde{\mathbb{I}}_{B_i}' \iota_p$ is the number of columns in B_i with at least one nonzero element, $\tilde{q}_i = \tilde{\mathbb{I}}_{C_i}' \iota_q$ is the number of columns in C_i with at least one nonzero element. Define $d_{11,i} = \frac{p(p-1)}{2} + \tilde{p}_i$ as the number of regressors in the first part of $x_{1,i}$ related to quadratic interactions among common instruments after selection, and $d_{1,i} = d_{11,i} + p\tilde{q}_i$ as the total number of regressors in $x_{1,i}$. $\tilde{B}_i$, $\tilde{C}_i$ are $K\tilde{p} \times Kp$ and $K\tilde{q} \times Kq$ matrices obtained by removing rows in $diag(vec[\mathbb{I}_{B_i}])$, $diag(vec[\mathbb{I}_{C_i}])$ for which all columns are zero. $\tilde{D}_i$ is the $\frac{p(p+1)}{2} \times d_{11,i}$ matrix $diag\left(vech\left[diag(\tilde{\mathbb{I}}_{B_i}) + \iota_p \iota_p' - I_p\right]\right)$ in which columns with all zeros have been removed. $\tilde{E}_i$ is the $pq \times p\tilde{q}$ matrix $diag(\iota_p \otimes \tilde{\mathbb{I}}_{C_i})$ in which columns with all zeros have been removed.

With the restricted $\tilde{B}_i$ and $\tilde{C}_i$, the first-pass regression becomes:

$$r_{i,t} = b'_{1,i} x_{1,i,t} + b'_{2,i} x_{2,i,t} + \epsilon_{i,t} \quad (\text{B6})$$

$$\begin{aligned} x_{1,i,t} &= (vech(X_t)' \tilde{D}_i, (Z'_{c,t-1} \otimes Z'_{i,t-1}) \tilde{E}_i)' \\ x_{2,i,t} &= [(f'_{c,t} \otimes Z'_{c,t-1}) \tilde{B}_i', (f'_{c,t} \otimes Z'_{i,t-1}) \tilde{C}_i']'. \end{aligned} \quad (\text{B7})$$

The coefficients $b_{1,i}$ and $b_{2,i}$ are transformations of the coefficients B_i , C_i , Λ_c and F_c of the

linear specifications Eq. (24), Eq. (21), Eq. (22):

$$b_{1,i} = \left((\tilde{D}'_i N_p [(\Lambda_c - F_c)' \otimes I_p] \tilde{B}'_i \tilde{B}_i \text{vec}[B'_i])', (\tilde{E}'_i [(\Lambda_c - F_c)' \otimes I_q] \text{vec}[C'_i])' \right)' \quad (\text{B8})$$

$$b_{2,i} = \left((\tilde{B}_i \text{vec}[B'_i])', (\tilde{C}_i \text{vec}[C'_i])' \right)' \quad (\text{B9})$$

where $W_{p,q}$ is the $p \times q$ commutation matrix and $N_p = \frac{1}{2} D_p^+ (W_{p,p} + I_{p^2})$ where D_p^+ is the $p(p+1)/2 \times p^2$ matrix such that $\text{vech}(A) = D_p^+ \text{vec}(A)$.

The restrictions $\mathbb{I}_{B_i}$ and $\mathbb{I}_{C_i}$ are determined using an iterative procedure:

1. Run the regression setting all elements in $\mathbb{I}_{B_i}$ and $\mathbb{I}_{C_i}$ to one.
2. Calculate the condition number of the matrix $\hat{Q}_{x,i} = \frac{1}{T_i} \sum_t I_{i,t} x_{i,t} x'_{i,t}$:

$$CN(\hat{Q}_{x,i}) = \sqrt{\text{eig}_{\max}(\hat{Q}_{x,i}) / \text{eig}_{\min}(\hat{Q}_{x,i})} \quad (\text{B10})$$

3. If the condition number is above 25 or the determinant of Q_x is less than machine precision, then find the pair of regressors in $x_{i,2}$ that has the largest cross-correlation in absolute value. Of these two regressors, remove the one with the lowest absolute correlation with $r_{i,t}$ and set the corresponding column in $\mathbb{I}_{B_i}$ or $\mathbb{I}_{C_i}$ to 0.
4. Check the following condition on $\mathbb{I}_{B_i}$: the first column of B_i , i.e. constants are all selected. Otherwise, keep the regressor and look for the next regressor pair with the highest correlation.
5. Construct new regressions using updated indicator matrices $\mathbb{I}_{B_i}$ and $\mathbb{I}_{C_i}$ and repeat the regression.

Appendix B.3. WLS regression and inference

The second-pass cross-sectional regression under instrument selection is:

$$\hat{b}_{1,i} = \hat{b}_{3,i}^R \theta_c, \quad \hat{b}_{3,i}^R \equiv \hat{b}_{3,i} H_c \quad (\text{B11})$$

$$b_{3,i} = \left(\left(\tilde{D}'_i N_p [B'_i \otimes I_p] \right)' , \left[\tilde{E}'_i W_{p,q} (C'_i \otimes I_p) \right]' \right)' \quad (\text{B12})$$

$$v_c = \text{vec}[\Lambda'_c - F'_c] = H_c \theta_c. \quad (\text{B13})$$

A weighted least squares (WLS) estimator for v_c is:

$$\hat{\theta}_c^{WLS} = (\hat{Q}_{b_3}^R)^{-1} \frac{1}{n_c} \sum_i (\hat{b}_{3,i}^R)' \hat{w}_i^R \hat{b}_{1,i}, \quad \hat{v}_c^R = H_c \hat{\theta}_c^{WLS}, \quad (\text{B14})$$

where $\hat{Q}_{b_3}^R = \frac{1}{n_c} \sum_i (\hat{b}_{3,i}^R)' \hat{w}_i^R \hat{b}_{3,i}^R$ and $\hat{w}_i^R = \mathbf{1}_i (\text{diag}[\hat{v}_i^R])^{-1}$ are the weights. The weights are set to be the inverse of the asymptotic variances of the standardized errors $\sqrt{T_c}(\hat{b}_{1,i} - \hat{b}_{3,i}^R \theta_c)$ in the cross-sectional regression for large T from a first-step OLS regression that imposes the same restriction, $\hat{\theta}_c = (\frac{1}{n_c} \sum_i (\hat{b}_{3,i}^R)' \hat{b}_{3,i}^R)^{-1} \frac{1}{n_c} \sum_i (\hat{b}_{3,i}^R)' \hat{b}_{1,i}$. With the OLS estimator $\hat{v}_c^R = H_c \hat{\theta}_c$, I calculate $\hat{v}_i^R$ using the following relations: $\hat{v}_i^R = \tau_{i,c} (C_{\hat{v}_c,i}^R)' \hat{Q}_{x,i}^{-1} S_{ii} \hat{Q}_{x,i}^{-1} C_{\hat{v}_c,i}^R$, where $\tau_i = \frac{T_c}{T_i}$, $\hat{Q}_{x,i} = \frac{1}{T_i} \sum_t I_{i,t} x_{i,t} x'_{i,t}$, $\hat{S}_{ii} = \frac{1}{T_i} \sum_t I_{i,t} \hat{\epsilon}_{i,t}^2 x_{i,t} x'_{i,t}$, $\hat{\epsilon}_{i,t} = r_{i,t} - \hat{b}_i' x_{i,t}$ and $C_{\hat{v}_c,i}^R = (E'_{1,i} - (I_{d_{1,i}} \otimes (\hat{v}_c^R)') J_{a,i} E'_{2,i})'$, $E_{1,i} = (I_{d_{1,i}}, \mathbf{0}_{d_{1,i}, d_{2,i}})'$, $E_{2,i} = (\mathbf{0}_{d_{2,i}, d_{1,i}}, I_{d_{2,i}})'$.

$b_{3,i}$ is obtained by invoking the following identity:

$$\text{vec}(b'_{3,i}) = J_{a,i} b_{2,i} \quad (\text{B15})$$

$$J_{a,i} = \begin{bmatrix} J_{1,i} & O \\ O & J_{2,i} \end{bmatrix} \quad (\text{B16})$$

$$J_{1,i} = W_{p(p-1)/2 + \tilde{p}_i, Kp} (I_{pK} \otimes (\tilde{D}'_i N_p) \times \{I_K \otimes [(W_p \otimes I_p)(I_p \otimes \text{vec}[I_p])]\}) \tilde{B}'_{i,c} \quad (\text{B17})$$

$$J_{2,i} = W_{pq, pK} (I_K \otimes [(I_p \otimes \tilde{E}'_i W_{p,q})(W_{p,q} \otimes I_p)(I_q \otimes \text{vec}(I_p))]) \tilde{C}'_{i,c}, \quad (\text{B18})$$

where $W_{p,q}$ is the $p \times q$ commutation matrix and $N_p = \frac{1}{2} D_p^+ (W_{p,p} + I_{p^2})$, and D_p^+ is the $p(p+1)/2 \times p^2$ matrix such that $\text{vech}(A) = D_p^+ \text{vec}(A)$.

The distribution of the estimator $\hat{\theta}_c^{WLS}$ is:

$$\sqrt{n_c T_c}(\hat{\theta}_c^{WLS} - \frac{1}{T_c} \hat{B}_{\theta_c} - \theta_c) \Rightarrow N(0, \Sigma_{\theta_c}), \quad (B19)$$

The convergence rate is unchanged by the restriction; identification requires $H'_c Q_{b_3} H_c$ to be positive definite. $\hat{B}_{\theta_c}$ is the bias correction:

$$\hat{B}_{\theta_c} = (\hat{Q}_{b_3}^R)^{-1} \frac{1}{n_c} \sum_i \tau_i H'_c J_{b,i} \text{vec}[E'_{2,i} \hat{Q}_{x,i}^{-1} \hat{S}_{ii} \hat{Q}_{x,i}^{-1} C_{\hat{y}_c,i}^R \hat{w}_i^R]. \quad (B20)$$

And the covariance matrix of the free wedge parameters θ_c is:

$$\begin{aligned} \hat{\Sigma}_{\theta_c} &= (\hat{Q}_{b_3}^R)^{-1} H'_c \hat{S}^R H_c (\hat{Q}_{b_3}^R)^{-1} \\ \hat{S}^R &= \frac{1}{n_c} \sum_{i,j} \frac{\tau_i \tau_j}{\tau_{i,j}} b'_{3,i} w_i^R (C_{\hat{y}_c,i}^R)' \hat{Q}_{x,i}^{-1} \tilde{S}_{ij} \hat{Q}_{x,j}^{-1} C_{\hat{y}_c,j}^R (w_j^R)' b_{3,j}, \tau_{i,j} = \frac{T_c}{T_{ij}} \\ J_{b,i} &= (\text{vec}[I_{d_{1,i}}]' \otimes I_{Kp})(I_{d_{1,i}} \otimes J_{a,i}) \\ C_{\hat{y}_c,i}^R &= (E'_{1,i} - (I_{d_{1,i}} \otimes (\hat{y}_c^R)') J_{a,i} E'_{2,i})'. \end{aligned} \quad (B21)$$

The covariance matrix of the full restricted vector, $\hat{\Sigma}_{v_c}^R = H_c \hat{\Sigma}_{\theta_c} H'_c$, has rank r : the restricted positions have zero sampling variance, so inference is conducted on θ_c .

The bias correction $\hat{B}_{\theta_c}$ accounts for the error-in-variable problem in the first-pass regression. A hard thresholded estimator $\tilde{S}_{ij} = \hat{S}_{ij} \mathbf{1}_{\|\hat{S}_{ij}\| \geq \kappa_{n_c, T_c}}$ is used, where $\hat{S}_{ij} = \frac{1}{T_{ij}} \sum_t I_{i,t} I_{j,t} \hat{\epsilon}_{i,t} \hat{\epsilon}_{j,t} x_{i,t} x'_{j,t}$, $\|\hat{S}_{ij}\|$ is the Frobenius norm, $\kappa_{n_c, T_c} = M \sqrt{\frac{\log(n_c)}{T_c}}$ is a data-dependent threshold and M is a positive number set by cross-validation.

To obtain estimates of time-varying risk premia $\hat{\Lambda}_c$, a SUR of $f_{c,t}$ is run on lagged common instruments $Z_{c,t-1}$:

$$f_{c,t} = F_c Z_{c,t-1} + u_t \quad (B22)$$

The estimator for F_c is:

$$\hat{F}_c = (\sum f_{c,t} Z'_{c,t-1}) (\sum_t Z_{c,t-1} Z'_{c,t-1})^{-1}. \quad (\text{B23})$$

The world row of F_c is estimated on a constant and the lagged world dividend yield only, so that it is common across markets, and its loadings on the local instruments are zero; the local rows are estimated on $Z_{c,t-1}$.

$\hat{\Lambda}_c$ follows from $\text{vec}(\hat{\Lambda}'_c) = \text{vec}(\hat{F}'_c) + H_c \hat{\theta}_c$, with $\hat{\theta}_c$ the bias-corrected WLS estimate. At the restricted positions $\hat{\Lambda}_c$ equals $\hat{F}_c$ exactly, so the world risk premium and its standard error come from the factor regression alone. The asymptotic distribution of Λ_c is: $\sqrt{T_c} \text{vec}[\hat{\Lambda}'_c - \Lambda'_c] \Rightarrow N(0, \Sigma_{\Lambda_c})$, where

$$\Sigma_{\Lambda_c} = (\mathbb{I}_K \otimes Q_z^{-1}) \Sigma_u (\mathbb{I}_K \otimes Q_z^{-1}) \quad (\text{B24})$$

$$\Sigma_u = E[u_t u'_t \otimes Z_{c,t-1} Z'_{c,t-1}]$$

$$Q_z = E[Z_{c,t-1} Z'_{c,t-1}].$$

Eq. (B24) is the expression for Σ_{Λ_c} when all rows share the same instruments. With the restriction on the world row, Σ_{Λ_c} is computed equation by equation with each row's own instruments, including the cross-equation blocks. The asymptotic distribution of Λ_c is dominated by the asymptotic distribution of F_c because $\hat{v}_c$ has a faster convergence rate than F_c . Finally, the time-varying risk premia $\mu_{c,t}$ are calculated as $\hat{\mu}_{c,t} = \hat{\Lambda}_c Z_{c,t-1}$, with asymptotic distribution:

$$\sqrt{T}(\hat{\mu}_t - \mu_t) \Rightarrow N(0, (Z'_{c,t-1} Q_z^{-1} \otimes I_K) E[Z_{t-1} Z'_{t-1} \otimes u_t u'_t] (Q_z^{-1} Z_{t-1} \otimes I_K)). \quad (\text{B25})$$

In the unconditional estimation the constant is the only common instrument, $p = 1$, so $v_c \in \mathbb{R}^K$ and H_c selects the two local wedge parameters; the world premium is then the sample mean of the attainable world factor, with the standard error of a sample mean.

Tables 3 and 4 also report the cross-market averages of the local premia with their standard

errors. Because $\hat{\nu}_c$ converges faster than F_c , the estimation error of the average $\frac{1}{N} \sum_c \hat{\lambda}_c$ is dominated by that of $\frac{1}{N} \sum_c \bar{f}_c$, where $\bar{f}_c = \frac{1}{T_c} \sum_{t=1}^{T_c} f_{c,t}$ is the factor sample mean over market c 's estimation window. Substituting for $\bar{f}_c$ and exchanging the order of summation, the dominating component becomes $\sum_{t=1}^T \frac{1}{N} \sum_c f_{c,t}/T_c$, where the outer sum runs over the full sample window and the inner sum over the markets whose estimation window includes month t . Its sampling error can therefore be written as $\sum_t h_t$, where $h_t = \frac{1}{N} \sum_c (f_{c,t} - E[f_{c,t}])/T_c$. The reported standard error is the Newey-West standard error of $\sum_t h_t$, with $E[f_{c,t}]$ replaced by the sample mean $\bar{f}_c$ in estimation. Because markets are aggregated within each month before the long-run variance is computed, this standard error accounts for both serial correlation and cross-market correlation of the factors.

References

- Akbari, Amir, Francesca Carrieri, and Aytek Malkhozov (2022). “Can Cross-border Funding Frictions Explain Financial Integration Reversals?” *Review of Financial Studies* 35 (1), pp. 394–437.
- Alexander, Gordon, Cheol S Eun, and Sundaram Janakiramanan (1987). “Asset Pricing and Dual Listing on Foreign Capital Markets: A Note”. *Journal of Finance* 42 (1), pp. 151–158.
- Bartram, Söhnke M, John M Griffin, Tae-Hoon Lim, and David T Ng (2015). “How Important Are Foreign Ownership Linkages for International Stock Returns?” *Review of Financial Studies* 28 (11), pp. 3036–3072.
- Basak, Suleyman and Anna Pavlova (2013). “Asset Prices and Institutional Investors”. *American Economic Review* 103 (5), pp. 1728–58.
- Bekaert, Geert and Campbell R Harvey (1995). “Time-varying World Market Integration”. *Journal of Finance* 50 (2), pp. 403–444.
- Bekaert, Geert, Campbell R Harvey, Andrea Kiguel, and Xiaozheng Wang (2016). “Globalization and Asset Returns”. *Annual Review of Financial Economics* 8 (1), pp. 221–288.
- Bekaert, Geert, Campbell R Harvey, and Christian Lundblad (2007). “Liquidity and Expected Returns: Lessons from Emerging Markets”. *Review of Financial Studies* 20 (6), pp. 1783–1831.
- Bekaert, Geert, Campbell R Harvey, Christian Lundblad, and Stephan Siegel (2011). “What Segments Equity Markets?” *Review of Financial Studies* 24 (12), pp. 3841–3890.
- Bekaert, Geert and Xiaozheng Sandra Wang (2009). “Home Bias Revisited”. *Columbia Business School Working Paper*.
- Boguth, Oliver, Murray Carlson, Adlai Fisher, and Mikhail Simutin (2011). “Conditional Risk and Performance Evaluation: Volatility Timing, Overconditioning, and New Estimates of Momentum Alphas”. *Journal of Financial Economics* 102 (2), pp. 363–389.
- Camanho, Nelson, Harald Hau, and Hélène Rey (2022). “Global Portfolio Rebalancing and Exchange Rates”. *Review of Financial Studies* 35 (11), pp. 5228–5274.

- Chaieb, Ines and Vihang Errunza (2007). “International Asset Pricing under Segmentation and PPP Deviations”. *Journal of Financial Economics* 86 (2), pp. 543–578.
- Chaieb, Ines, Hugues Langlois, and Olivier Scaillet (2021). “Factors and Risk Premia in Individual International Stock Returns”. *Journal of Financial Economics* 141 (2), pp. 669–692.
- Chan, Kalok, Vicentiu Covrig, and Lilian Ng (2005). “What Determines the Domestic Bias and Foreign Bias? Evidence from Mutual Fund Equity Allocations Worldwide”. *Journal of Finance* 60 (3), pp. 1495–1534.
- De Jong, Frank and Frans A De Roon (2005). “Time-varying Market Integration and Expected Returns in Emerging Markets”. *Journal of Financial Economics* 78 (3), pp. 583–613.
- De Marco, Filippo, Marco Macchiavelli, and Rosen Valchev (2022). “Beyond Home Bias: International Portfolio Holdings and Information Heterogeneity”. *Review of Financial Studies* 35 (9), pp. 4387–4422.
- Du, Wenxin and Amy W Huber (2024). *Dollar Asset Holdings and Hedging around the Globe*. NBER Working Paper.
- Errunza, Vihang and Etienne Losq (1985). “International Asset Pricing under Mild Segmentation: Theory and Test”. *Journal of Finance* 40 (1), pp. 1–20.
- Faias, José A and Miguel A Ferreira (2017). “Does Institutional Ownership Matter for International Stock Return Comovement?” *Journal of International Money and Finance* 78, pp. 64–83.
- Fama, Eugene F and Kenneth R French (2012). “Size, Value, and Momentum in International Stock Returns”. *Journal of Financial Economics* 105 (3), pp. 457–472.
- (2007). “Disagreement, Tastes, and Asset Prices”. *Journal of Financial Economics* 83 (3), pp. 667–689.
- Ferreira, Miguel A and Pedro Matos (2008). “The Colors of Investors’ Money: The Role of Institutional Investors around the World”. *Journal of Financial Economics* 88 (3), pp. 499–533.
- Fong, Kingsley Y L, Craig W Holden, and Charles A Trzcinka (2017). “What Are the Best Liquidity Proxies for Global Research?” *Review of Finance* 21 (4), pp. 1355–1401.

- Gagliardini, Patrick, Elisa Ossola, and Olivier Scaillet (2016). “Time-varying Risk Premium in Large Cross-sectional Equity Data Sets”. *Econometrica* 84 (3), pp. 985–1046.
- (2019). “A Diagnostic Criterion for Approximate Factor Structure”. *Journal of Econometrics* 212 (2), pp. 503–521.
- Glossner, Simon, Pedro Matos, Stefano Ramelli, and Alexander F Wagner (2025). “Do Institutional Investors Stabilize Equity Markets in Crisis Periods? Evidence from COVID-19”. *Management Science* 71 (10), pp. 8623–8641.
- Greenwood, Robin, Samuel G Hanson, and Gordon Y Liao (2018). “Asset Price Dynamics in Partially Segmented Markets”. *Review of Financial Studies* 31 (9), pp. 3307–3343.
- Gromb, Denis and Dimitri Vayanos (2002). “Equilibrium and Welfare in Markets with Financially Constrained Arbitrageurs”. *Journal of Financial Economics* 66 (2-3), pp. 361–407.
- Hau, Harald and Helene Rey (2008). “Home Bias at the Fund Level”. *American Economic Review* 98 (2), pp. 333–338.
- Hollstein, Fabian (2022). “Local, Regional, or Global Asset Pricing?” *Journal of Financial and Quantitative Analysis* 57 (1), pp. 291–320.
- Hou, Kewei, Andrew Karolyi, and Bong-Chan Kho (2011). “What Factors Drive Global Stock Returns?” *Review of Financial Studies* 24 (8), pp. 2527–2574.
- Jagannathan, Ravi and Zhenyu Wang (1996). “The Conditional CAPM and the Cross-section of Expected Returns”. *Journal of Finance* 51 (1), pp. 3–53.
- Jensen, Theis Ingerslev, Bryan Kelly, and Lasse Heje Pedersen (2023). “Is There a Replication Crisis in Finance?” *Journal of Finance* 78 (5), pp. 2465–2518.
- Jotikasthira, Chotibhak, Christian T. Lundblad, and Tarun Ramadorai (2012). “Asset Fire Sales and Purchases and the International Transmission of Funding Shocks”. *Journal of Finance* 67 (6), pp. 2015–2050.
- Kacperczyk, Marcin, Jaromir Nosal, and Tianyu Wang (2025). “Global Volatility and Firm-level Capital Flows”. *Journal of Financial Economics* 169, p. 104078.

- Kacperczyk, Marcin, Savitar Sundaresan, and Tianyu Wang (2021). “Do Foreign Institutional Investors Improve Price Efficiency?” *Review of Financial Studies* 34 (3), pp. 1317–1367.
- Karolyi, Andrew and Rene Stulz (2003). “Are Financial Assets Priced Locally or Globally?” *Handbook of the Economics of Finance*. Vol. 1. Elsevier, pp. 975–1020.
- Karolyi, G Andrew and Ying Wu (2018). “A New Partial-segmentation Approach to Modeling International Stock Returns”. *Journal of Financial and Quantitative Analysis* 53 (2), pp. 507–546.
- Koijen, Ralph SJ, Robert J Richmond, and Motohiro Yogo (2024). “Which Investors Matter for Equity Valuations and Expected Returns?” *Review of Economic Studies* 91 (4), pp. 2387–2424.
- Koijen, Ralph SJ and Motohiro Yogo (2019). “A Demand System Approach to Asset Pricing”. *Journal of Political Economy* 127 (4), pp. 1475–1515.
- Lau, Sie Ting, Lilian Ng, and Bohui Zhang (2010). “The World Price of Home Bias”. *Journal of Financial Economics* 97 (2), pp. 191–217.
- Lewellen, Jonathan and Stefan Nagel (2006). “The Conditional CAPM Does Not Explain Asset-pricing Anomalies”. *Journal of Financial Economics* 82 (2), pp. 289–314.
- Lewis, Karen K. (2011). “Global Asset Pricing”. *Annual Review of Financial Economics* 3 (1), pp. 435–466.
- Luo, H. Arthur and Ronald J. Balvers (2017). “Social Screens and Systematic Investor Boycott Risk”. *Journal of Financial and Quantitative Analysis* 52 (1), pp. 365–399.
- Matos, Pedro (2020). “ESG and Responsible Institutional Investing around the World: A Critical Review”. *CFA Institute Research Foundation*.
- Merton, Robert C (1987). “A Simple Model of Capital Market Equilibrium with Incomplete Information”. *Journal of Finance* 42 (3), pp. 483–510.
- Opie, Wei and Steven Riddiough (2024). “On the Use of Currency Forwards: Evidence from International Equity Mutual Funds”. *University of Toronto Working Paper*.
- Pavlova, Anna and Roberto Rigobon (2008). “The Role of Portfolio Constraints in the International Propagation of Shocks”. *Review of Economic Studies* 75 (4), pp. 1215–1256.

- Pavlova, Anna and Taisiya Sikorskaya (2023). “Benchmarking Intensity”. *Review of Financial Studies* 36 (3), pp. 859–903.
- Petzev, Ivan, Andreas Schrimpf, and Alexander F Wagner (2016). “Has the Pricing of Stocks Become More Global?” *BIS Working Paper*.
- Stulz, Rene M (1981). “On the Effects of Barriers to International Investment”. *Journal of Finance* 36 (4), pp. 923–934.
- Zerbib, Olivier David (2022). “A Sustainable Capital Asset Pricing Model (S-CAPM): Evidence from Environmental Integration and Sin Stock Exclusion”. *Review of Finance* 26 (6), pp. 1345–1388.

Table 1. Summary Statistics.

This table reports summary statistics of the sample of individual stocks and local pricing factors used in the estimation. For each market, it reports the number of local stocks (# Stocks), the starting date, and the average excess return and volatility of the institutional local factor (Inst. Local) and the retail local factor (Retail Local). The sample spans January 2003 to December 2024. Individual stock U.S. dollar returns are constructed following Jensen et al. (2023). Average returns and volatilities are annualized and expressed in percentage points.

			Average return (%)		Volatility (%)	
Market	# Stocks	Start Date	Inst. Local	Retail Local	Inst. Local	Retail Local
Panel A: Developed Markets						
Australia	3,446	31-Jan-2003	-1.289	0.643	9.586	8.628
Austria	160	31-Jan-2003	-0.327	5.795	9.859	13.806
Belgium	279	31-Jan-2003	-0.394	4.599	8.891	13.706
Canada	2,447	31-Jan-2003	-0.941	2.050	8.363	14.145
Denmark	364	31-Jan-2003	5.262	-0.794	10.227	11.689
Finland	266	31-Jan-2003	-1.818	1.516	10.721	9.719
France	1,523	31-Jan-2003	0.229	2.142	5.036	9.987
Germany	1,247	31-Jan-2003	-0.238	3.699	6.565	8.430
Hong Kong	1,731	31-Jan-2003	0.986	-6.504	11.479	10.490
Israel	829	31-Jan-2003	1.100	1.263	13.944	7.870
Italy	719	31-Jan-2003	-0.810	-5.694	8.178	9.203
Japan	5,213	31-Jan-2003	0.032	-3.419	10.200	5.708
Netherlands	270	31-Jan-2003	-0.301	0.156	7.219	14.650
New Zealand	248	31-Jan-2003	1.872	-1.371	11.143	10.665
Norway	546	31-Jan-2003	0.422	5.152	13.147	10.513
Singapore	933	31-Jan-2003	1.539	-0.183	7.991	8.578
Spain	427	31-Jan-2003	2.058	1.269	10.268	13.838
Sweden	1,046	31-Jan-2003	1.749	-5.140	8.735	13.140
Switzerland	372	31-Jan-2003	3.012	-1.416	6.407	10.731
UK	3,809	31-Jan-2003	-1.744	-3.705	5.951	7.653
US	13,620	31-Jan-2003	3.915	-3.919	6.512	7.177
Panel B: Emerging Markets						
Brazil	373	31-Oct-2003	2.227	4.413	17.443	18.012
Chile	260	31-Jan-2003	1.231	4.923	14.992	10.138
China	5,300	31-Oct-2005	10.817	-3.544	25.363	9.280
Egypt	289	30-Nov-2003	7.239	-2.118	23.300	15.562
Greece	426	31-Jan-2003	-5.574	-1.492	20.504	11.705
India	6,208	31-Jan-2003	6.135	1.248	17.882	18.906
Indonesia	1,018	31-Jan-2003	4.208	11.215	15.600	13.245
Korea	3,607	31-Jan-2003	-2.156	-5.051	13.637	9.904
Kuwait	227	31-May-2008	-2.500	-0.680	13.683	7.995
Malaysia	1,511	31-Jan-2003	1.064	4.695	10.186	5.816
Mexico	206	31-Jan-2003	0.777	3.514	12.939	12.515
Philippines	341	31-Jan-2003	4.471	2.434	14.295	18.183
Poland	1,221	31-Jan-2003	-1.768	4.127	15.869	11.193
Qatar	59	31-May-2008	1.875	3.804	14.725	24.700
Saudi Arabia	358	31-Jan-2010	3.577	-1.035	15.584	9.718
South Africa	801	31-Jan-2003	0.458	-1.459	13.492	12.994
Taiwan	2,749	31-Jan-2003	2.014	-1.450	13.305	6.551
Thailand	1,103	31-Jan-2003	3.506	1.070	14.369	8.726
Turkiye	706	31-Jan-2003	5.953	2.465	27.180	30.904
UAE	167	31-Oct-2008	2.744	1.652	14.685	12.653

Table 2. Additional Factor Summary Statistics.

This table reports, for each market, the R^2 of the full-sample static OLS projections that define the attainable retail and substitute portfolios in Section 2.3, and the full-sample correlations among the institutional local (f^{inst}), retail local (f^{retail}), and attainable world (f^W) factors. The sample spans January 2003 to December 2024.

Market	R-squared		Correlations		
	$\hat{R}$	D_s	$\text{Corr}(f^{inst}, f^{retail})$	$\text{Corr}(f^{retail}, f^W)$	$\text{Corr}(f^{inst}, f^W)$
Panel A: Developed Markets					
Australia	0.903	0.811	-0.055	0.022	0.013
Austria	0.412	0.827	-0.041	-0.010	-0.003
Belgium	0.381	0.804	-0.215	0.050	-0.001
Canada	0.684	0.804	0.066	-0.038	0.012
Denmark	0.645	0.714	-0.124	0.070	0.002
Finland	0.747	0.769	-0.142	0.063	0.001
France	0.689	0.936	-0.084	0.020	0.004
Germany	0.649	0.894	-0.020	-0.021	0.013
Hong Kong	0.830	0.703	-0.095	0.075	0.034
Israel	0.880	0.536	-0.038	0.040	0.009
Italy	0.818	0.867	-0.149	0.063	0.012
Japan	0.864	0.505	0.190	-0.170	0.054
Netherlands	0.555	0.886	0.022	-0.008	-0.003
New Zealand	0.695	0.661	-0.049	0.030	0.000
Norway	0.761	0.728	-0.087	0.031	0.007
Singapore	0.846	0.821	-0.060	0.018	0.001
Spain	0.424	0.791	-0.079	0.023	0.005
Sweden	0.795	0.854	0.006	0.000	0.005
Switzerland	0.563	0.835	0.110	-0.041	0.008
UK	0.858	0.880	-0.090	0.051	0.024
US	0.872	0.816	-0.232	0.082	0.193
<i>Average</i>	0.708	0.783	-0.056	0.017	0.019
Panel B: Emerging Markets					
Brazil	0.589	0.671	-0.017	-0.006	0.006
Chile	0.696	0.524	-0.047	0.030	0.008
China	0.917	0.237	0.020	-0.006	0.075
Egypt	0.739	0.332	0.020	-0.030	-0.003
Greece	0.750	0.588	-0.107	0.075	0.001
India	0.783	0.563	0.066	-0.091	0.022
Indonesia	0.694	0.603	0.023	0.004	0.001
Korea	0.887	0.714	-0.084	0.073	0.012
Kuwait	0.754	0.418	-0.053	0.050	0.003
Malaysia	0.906	0.597	-0.045	0.049	0.000
Mexico	0.577	0.668	0.058	-0.071	-0.003
Philippines	0.377	0.470	-0.034	0.024	-0.001
Poland	0.798	0.702	-0.098	0.055	0.000
Qatar	0.275	0.477	0.051	-0.035	-0.002
Saudi Arabia	0.789	0.263	-0.041	0.048	0.012
South Africa	0.753	0.721	-0.111	0.107	0.002
Taiwan	0.930	0.608	-0.057	0.034	0.009
Thailand	0.839	0.596	0.096	-0.102	-0.004
Turkiye	0.582	0.416	-0.078	0.046	0.000
UAE	0.442	0.454	0.152	-0.188	0.005
<i>Average</i>	0.704	0.531	-0.014	0.003	0.007

Table 3. Unconditional Estimation.

This table reports the unconditional estimates of the institutional local premium (μ^{inst}) and the retail local premium (μ^{retail}) for 21 developed and 20 emerging markets, obtained using the Gagliardini et al. (2016) two-pass regression applied to individual stock returns. For each market, the table reports the number of stocks used in the estimation (# Stocks) and the estimated premia with standard errors in parentheses beneath. The standard error beneath *Average* is the Newey-West standard error of the cross-market average; # Pos. Sig. counts markets significantly positive at the 5% level (one-sided). The world premium (μ^W) is restricted to carry no pricing wedge and equals the sample mean of the attainable world factor over each market's estimation window; the *World* line reports its full-sample value and Newey-West standard error (window means for markets with a later start date or missing months differ slightly and are not reported). Markets with a negative Gagliardini et al. (2019) diagnostic are in bold. The sample spans January 2003 to December 2024. Individual stock U.S. dollar returns are constructed following Jensen et al. (2023). Local factors are constructed as in Section 2.3, with institutional stocks those with positive coverage (Section 2.2). Premia are annualized, in percentage points.

Developed Markets				Emerging Markets			
Market	# Stocks	Inst. Local	Retail Local	Market	# Stocks	Inst. Local	Retail Local
Australia	1,462	3.476 (1.642)	0.032 (1.851)	Brazil	156	-0.540 (3.623)	15.242 (3.748)
Austria	59	-0.874 (2.161)	6.401 (3.415)	Chile	147	2.024 (2.893)	4.906 (1.847)
Belgium	128	4.775 (1.831)	6.647 (2.741)	China	2,938	10.916 (6.549)	-0.513 (1.879)
Canada	906	12.097 (1.572)	-0.355 (3.047)	Egypt	157	9.449 (5.237)	0.444 (4.018)
Denmark	148	1.118 (2.133)	3.213 (2.635)	Greece	232	-1.683 (4.323)	5.257 (2.419)
Finland	125	-0.300 (2.040)	-4.881 (1.979)	India	2,119	10.170 (3.452)	-2.328 (3.801)
France	648	-0.558 (1.062)	9.878 (1.853)	Indonesia	445	6.334 (3.187)	8.585 (2.852)
Germany	637	1.228 (1.362)	3.831 (1.965)	Korea	1,942	1.270 (2.489)	-0.808 (2.179)
Hong Kong	1,177	1.818 (2.040)	0.929 (2.626)	Kuwait	157	-2.434 (3.484)	-0.280 (1.907)
Israel	378	5.318 (2.992)	7.598 (1.633)	Malaysia	944	1.649 (2.209)	2.226 (1.349)
Italy	246	-3.646 (1.466)	-7.257 (1.982)	Mexico	87	0.043 (2.536)	9.074 (2.405)
Japan	3,708	4.304 (1.984)	-0.705 (1.222)	Philippines	229	6.173 (2.896)	11.915 (3.616)
Netherlands	84	-1.232 (1.384)	21.406 (3.204)	Poland	509	-1.077 (2.779)	5.636 (2.820)
New Zealand	118	2.639 (2.215)	-0.051 (2.176)	Qatar	44	2.187 (3.361)	6.904 (4.860)
Norway	157	-2.796 (2.803)	9.630 (2.187)	Saudi Arabia	167	2.640 (3.786)	2.307 (2.298)
Singapore	543	1.707 (1.734)	3.845 (1.993)	South Africa	282	6.227 (2.574)	-6.271 (2.627)
Spain	129	-0.760 (2.139)	-2.285 (2.700)	Taiwan	1,648	4.976 (2.238)	-4.678 (1.228)
Sweden	346	4.208 (1.747)	-4.881 (2.920)	Thailand	604	2.446 (2.910)	3.206 (1.952)
Switzerland	199	6.228 (1.142)	0.732 (1.993)	Turkiye	380	9.002 (5.487)	16.599 (6.352)

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Developed Markets				Emerging Markets			
Market	# Stocks	Inst. Local	Retail Local	Market	# Stocks	Inst. Local	Retail Local
UK	1,310	0.698 (1.156)	-0.203 (1.812)	UAE	83	0.655 (3.930)	12.693 (2.890)
US	4,990	6.429 (1.360)	-1.901 (1.553)				
<i>Average</i>		2.185 (0.412)	2.458 (0.760)	<i>Average</i>		3.521 (0.710)	4.506 (0.708)
<i># Pos. Sig.</i>		8	8	<i># Pos. Sig.</i>		7	10
<i>World</i>	8.982 (3.581)						

Table 4. Unconditional Estimation Controlling for Additional Local Factors.

This table reports the unconditional estimates of the institutional local premium (μ^{inst}) and the retail local premium (μ^{retail}), controlling for additional local market factors including small-minus-big (SMB), high-minus-low (HML), and momentum (MOM), obtained using the Gagliardini et al. (2016) two-pass regression applied to individual stock returns. For each market, the table reports the number of stocks used in the estimation (# Stocks) and the estimated premia with standard errors in parentheses beneath. The standard error beneath *Average* is the Newey-West standard error of the cross-market average; # Pos. Sig. counts markets significantly positive at the 5% level (one-sided). The world premium (μ^W) is restricted to carry no pricing wedge and equals the sample mean of the attainable world factor over each market's estimation window; the *World* line reports its full-sample value and Newey-West standard error (window means for markets with a later start date or missing months differ slightly and are not reported). Markets with a negative Gagliardini et al. (2019) diagnostic are in bold. The sample spans January 2003 to December 2024. Individual stock U.S. dollar returns and local SMB, HML, MOM factors are constructed following Jensen et al. (2023). Premia are annualized, in percentage points.

Developed Markets							Emerging Markets						
Market	# Stocks	Inst. Local	Retail Local	SMB	HML	MOM	Market	# Stocks	Inst. Local	Retail Local	SMB	HML	MOM
Australia	1,462	5.242 (1.642)	0.427 (1.851)	2.333 (2.923)	-7.714 (2.776)	10.090 (3.069)	Brazil	155	3.708 (3.620)	14.060 (3.735)	-4.776 (3.779)	2.856 (3.752)	6.318 (4.694)
Austria	59	-1.912 (2.161)	3.570 (3.415)	-2.345 (2.716)	14.572 (4.755)	27.639 (5.053)	Chile	147	2.914 (2.893)	4.431 (1.847)	0.654 (2.133)	-4.399 (2.066)	12.908 (2.646)
Belgium	128	2.274 (1.831)	8.114 (2.741)	-4.050 (2.135)	-5.940 (4.479)	22.950 (4.795)	China	2,938	12.524 (6.549)	-1.091 (1.879)	1.191 (5.728)	-13.044 (4.337)	3.395 (4.264)
Canada	906	11.565 (1.572)	-0.452 (3.047)	2.030 (2.565)	-10.690 (3.027)	17.893 (3.529)	Egypt	151	5.599 (5.592)	-1.855 (3.794)	-3.021 (3.607)	-7.772 (3.818)	18.056 (4.415)
Denmark	148	1.333 (2.133)	0.854 (2.635)	3.761 (2.815)	0.794 (4.356)	4.649 (4.456)	Greece	232	-3.056 (4.280)	7.830 (2.367)	0.767 (3.550)	-11.218 (6.368)	20.897 (6.975)
Finland	125	1.442 (2.040)	-3.149 (1.979)	-0.986 (2.578)	-8.833 (4.209)	28.677 (3.389)	India	2,119	14.434 (3.452)	-1.921 (3.801)	1.811 (3.282)	-6.453 (3.895)	9.253 (4.815)
France	648	-0.269 (1.062)	9.183 (1.853)	0.473 (1.933)	-6.970 (3.022)	10.144 (3.052)	Indonesia	445	7.324 (3.187)	6.024 (2.852)	2.277 (3.027)	0.433 (4.532)	3.545 (4.644)
Germany	637	-1.230 (1.362)	2.542 (1.965)	3.930 (2.051)	0.384 (2.672)	-3.708 (3.258)	Korea	1,942	2.409 (2.489)	1.399 (2.179)	0.916 (3.167)	-9.103 (3.602)	18.282 (3.711)
Hong Kong	1,177	3.926 (2.040)	2.329 (2.626)	-0.322 (3.576)	-1.464 (3.344)	5.895 (4.410)	Kuwait	157	-1.310 (3.484)	-0.515 (1.907)	-1.912 (3.384)	-2.360 (3.189)	1.901 (3.263)
Israel	378	6.772 (2.992)	6.113 (1.633)	3.795 (3.847)	-3.735 (4.289)	5.165 (4.435)	Malaysia	944	5.261 (2.209)	2.240 (1.349)	-2.056 (2.287)	-5.309 (2.681)	7.727 (2.955)
Italy	246	-0.914 (1.466)	-8.004 (1.982)	-1.799 (2.377)	-12.116 (3.405)	17.609 (3.405)	Mexico	87	3.615 (2.536)	11.319 (2.405)	-3.877 (3.518)	-2.447 (4.478)	10.785 (4.245)
Japan	3,708	3.160 (1.984)	-1.651 (1.222)	1.633 (1.888)	-7.765 (2.770)	7.201 (2.746)	Philippines	229	8.404 (2.905)	12.871 (3.821)	-2.742 (3.253)	-4.762 (3.384)	13.064 (4.110)
Netherlands	84	-0.075	17.554	5.214	-19.889	4.870	Poland	509	1.095	5.464	3.865	-8.906	0.191

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[illegible]

Table 5. Conditional Estimation.

This table reports the estimated loadings of the institutional local premium and the retail local premium on the common instruments for 21 developed and 19 emerging markets, obtained using the Gagliardini et al. (2016) conditional two-pass regression applied to individual stock returns. The common instruments are the world dividend yield (DY^W), the domestic-minus-world dividend yield spread (DY^{D-W}), and the retail-minus-institutional dividend yield spread (DY^{R-I}). Each local premium is presented with its loadings on the constant and its matched spread. Local premia do not load on DY^W and their loadings on the other spread through conditional factor expectations are estimated but not reported. For each market, the table reports the number of stocks used in the estimation (# Stocks) and the estimated loadings with standard errors in parentheses beneath. The world premium loads on a constant and DY^W , with the spread loadings restricted to zero; the *World* line reports its full-sample loadings and standard errors (values for markets with a later start date or missing months differ slightly and are not reported). Markets with a negative Gagliardini et al. (2019) diagnostic are in bold. The sample spans January 2003 to December 2024. Individual stock U.S. dollar returns are constructed following Jensen et al. (2023). All premia are annualized and expressed in percentage points.

Developed Markets						Emerging Markets					
Market	# Stocks	Inst. Local		Retail Local		Market	# Stocks	Inst. Local		Retail Local	
		<i>const</i>	DY^{D-W}	<i>const</i>	DY^{R-I}			<i>const</i>	DY^{D-W}	<i>const</i>	DY^{R-I}
Australia	1,233	2.160	2.192	1.636	0.486	Chile	120	1.645	2.065	4.344	1.692
		(2.010)	(2.142)	(1.819)	(1.812)			(3.182)	(3.917)	(2.146)	(2.446)
Austria	49	-2.733	7.634	14.292	5.440	China	2,829	7.728	-0.299	3.616	4.511
		(2.097)	(4.049)	(2.922)	(6.392)			(5.768)	(7.351)	(2.111)	(2.565)
Belgium	118	-0.135	-1.586	6.993	-3.358	Egypt	138	5.506	-0.157	-3.575	-5.026
		(1.882)	(3.376)	(2.908)	(3.506)			(5.081)	(5.524)	(3.394)	(3.385)
Canada	779	2.465	-0.657	1.436	1.813	Greece	222	-2.374	-0.504	2.986	-2.476
		(1.778)	(2.070)	(3.003)	(3.267)			(4.382)	(4.625)	(2.476)	(2.914)
Denmark	138	-3.901	2.870	8.406	-2.556	India	1,750	10.262	22.261	3.245	-2.545
		(2.170)	(2.231)	(2.463)	(2.952)			(3.769)	(6.069)	(4.014)	(6.665)
Finland	117	-1.309	-1.409	1.457	-4.113	Indonesia	410	6.884	5.049	12.309	-5.589
		(2.278)	(2.698)	(2.063)	(2.479)			(3.313)	(3.501)	(2.807)	(3.421)
France	533	-1.172	-0.466	3.239	4.250	Korea	1,507	-5.024	2.024	-5.428	-10.463
		(1.070)	(1.205)	(2.121)	(2.791)			(2.898)	(3.525)	(2.096)	(2.759)
Germany	564	-0.520	1.204	-0.211	0.485	Kuwait	150	-4.427	2.296	4.016	-5.194
		(1.395)	(2.044)	(1.730)	(2.587)			(3.334)	(6.010)	(1.953)	(2.532)
Hong Kong	970	-2.769	-3.485	3.376	-8.295	Malaysia	845	3.778	9.063	3.860	1.624
		(2.440)	(3.367)	(2.222)	(2.749)			(2.138)	(2.456)	(1.235)	(1.533)
Israel	354	2.714	-1.866	3.746	1.131	Mexico	73	5.199	7.160	8.726	5.094
		(2.963)	(3.305)	(1.670)	(1.601)			(2.752)	(2.887)	(2.622)	(3.533)
Italy	224	-5.021	1.796	-2.190	8.570	Philippines	179	9.800	11.370	1.013	-10.089
		(1.732)	(2.470)	(1.925)	(2.654)			(3.018)	(3.123)	(3.850)	(3.081)
Japan	2,830	3.716	-1.486	-0.058	0.744	Poland	407	-2.142	-0.108	4.470	0.790
		(2.170)	(3.256)	(1.214)	(1.663)			(3.374)	(4.525)	(2.332)	(2.774)

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Developed Markets						Emerging Markets					
Market	# Stocks	Inst. Local		Retail Local		Market	# Stocks	Inst. Local		Retail Local	
		$const$	DY^{D-W}	$const$	DY^{R-I}			$const$	DY^{D-W}	$const$	DY^{R-I}
Netherlands	74	-3.694 (1.535)	-1.854 (1.932)	4.291 (3.107)	0.267 (4.139)	Qatar	44	0.911 (3.559)	4.643 (3.910)	12.210 (6.031)	-8.394 (6.780)
New Zealand	105	-1.817 (2.351)	2.377 (3.466)	7.330 (2.266)	0.092 (2.962)	Saudi Arabia	166	-4.667 (3.981)	9.343 (4.778)	6.157 (2.487)	-0.718 (3.489)
Norway	146	-0.105 (2.789)	-3.303 (5.344)	6.739 (2.201)	-1.328 (4.000)	South Africa	253	9.339 (2.854)	4.932 (2.939)	-0.657 (2.756)	4.877 (2.217)
Singapore	440	5.077 (1.698)	8.222 (1.840)	-0.294 (1.809)	-0.423 (1.917)	Taiwan	1,453	3.386 (2.825)	6.632 (4.965)	-1.121 (1.392)	0.290 (1.910)
Spain	117	-3.913 (2.184)	1.110 (3.841)	-0.670 (2.914)	3.897 (4.245)	Thailand	489	2.619 (3.047)	9.774 (3.107)	0.924 (1.824)	4.727 (1.686)
Sweden	300	3.138 (1.847)	-0.156 (2.117)	-2.016 (2.775)	0.967 (2.747)	Turkiye	334	15.638 (5.730)	29.280 (7.906)	5.769 (6.565)	15.862 (5.517)
Switzerland	170	-0.028 (1.362)	-0.175 (1.565)	11.488 (2.265)	2.541 (4.745)	UAE	76	0.515 (3.559)	6.285 (4.154)	6.700 (3.110)	-0.071 (3.192)
UK	1,000	-1.022 (1.266)	2.146 (2.018)	-1.860 (1.616)	0.717 (2.608)						
US	4,303	3.448 (1.380)	-2.699 (1.679)	-1.960 (1.524)	1.743 (1.986)						
Average		-0.258	0.496	3.103	0.622	Average		3.399	6.901	3.661	-0.584
World	$const$: 8.982 (3.240)		DY^W : 1.373 (5.320)								

Table 6. Institutional Ownership and Risk Premia: Security-Level Evidence.

This table reports the estimates from panel regressions as specified in Eq. (28) of stock-level attainable world, institutional local, and retail local factor exposures (β) and risk premia (μ) on the within-market percentile rank of global institutional ownership (IO^{rank} , in $[0, 1]$), the stock-specific instrument used in the conditional estimation. The coefficient on IO^{rank} is the differential associated with moving from the bottom to the top of the within-market ownership distribution. Stock-level controls include the logarithm of market capitalization ($LOGMV$), book-to-market ratio (BM), and Fong et al. (2017) illiquidity measure (FHT). All specifications include country-time fixed effects. Standard errors are clustered at the stock level and reported in parentheses. Conditional betas, BM , and FHT are winsorized at the 1st and 99th percentile within market. The sample spans January 2003 to December 2024. The risk-premia coefficients are annualized and expressed in percentage points. ***, **, and * denote significance at the 1%, 5%, and 10% levels, respectively.

	World		Inst. Local		Retail Local		Total
	β	μ	β	μ	β	μ	μ
Panel A: Developed Markets							
IO^{rank}	0.317*** (0.014)	2.889*** (0.123)	0.227*** (0.021)	0.655*** (0.073)	-0.338*** (0.021)	-0.660*** (0.071)	2.884*** (0.161)
$LOGMV$	-0.026*** (0.003)	-0.233*** (0.023)	0.005 (0.004)	-0.048*** (0.016)	-0.005 (0.004)	0.016 (0.011)	-0.264*** (0.030)
BM	-0.005* (0.003)	0.018 (0.027)	-0.003 (0.004)	-0.108*** (0.019)	-0.052*** (0.005)	-0.060** (0.024)	-0.150*** (0.040)
FHT	1.379*** (0.157)	12.371*** (1.438)	1.120*** (0.262)	5.315*** (1.363)	2.652*** (0.256)	3.225*** (0.631)	20.912*** (2.068)
Obs.	2,194,309	2,194,309	2,194,309	2,194,309	2,194,309	2,194,309	2,194,309
R2 Adj.	0.194	0.237	0.067	0.303	0.123	0.125	0.230
Panel B: Emerging Markets							
IO^{rank}	0.106*** (0.013)	1.041*** (0.117)	0.049*** (0.014)	-0.135 (0.104)	-0.666*** (0.025)	-1.220*** (0.119)	-0.314 (0.211)
$LOGMV$	0.004 (0.003)	0.023 (0.026)	-0.001 (0.003)	-0.043 (0.026)	-0.072*** (0.005)	-0.241*** (0.027)	-0.260*** (0.052)
BM	0.033*** (0.003)	0.304*** (0.028)	0.031*** (0.003)	0.325*** (0.036)	-0.057*** (0.005)	0.013 (0.028)	0.642*** (0.058)
FHT	0.178 (0.266)	1.739 (2.404)	-0.505* (0.287)	-3.754** (1.747)	1.921*** (0.599)	8.478*** (2.275)	6.464 (4.072)
Obs.	1,785,784	1,785,784	1,785,784	1,785,784	1,785,784	1,785,784	1,785,784
R2 Adj.	0.190	0.355	0.075	0.800	0.180	0.194	0.581

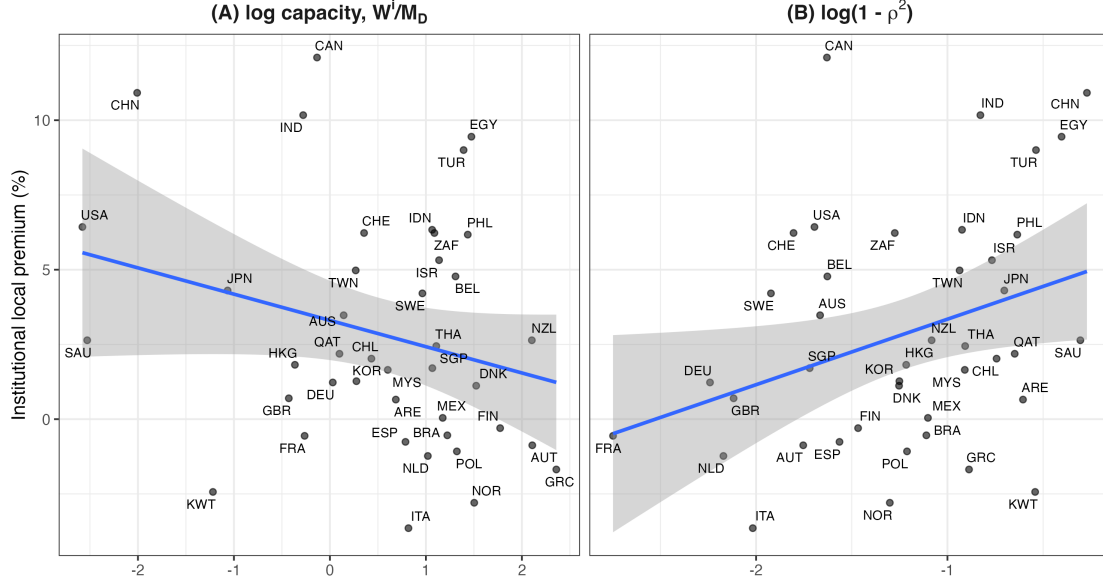


Figure 2. Unconditional Institutional Local Premium: Cross-market Evidence

This figure plots the estimated annualized unconditional institutional local risk premium reported in Table 3 against the capacity and correlation channels of Eq. (16). Panel A plots the premium against log institutional risk-bearing capacity, $\log(W^i/M_D)$, the time-series mean of active global assets under management scaled by local market capitalization. Panel B plots it against the log unspanned variance share $\log(1 - \rho^2)$, where ρ is the correlation between the attainable domestic market and its substitute portfolio. Markets are labeled by their three-letter ISO codes. The solid line is the fitted linear association and the shaded area the 95% confidence interval. The fitted pooled regression is

$$\hat{\mu}_c^{inst} = 9.16 - 0.96 \log(W^i/M_D)_c + 1.81 \log(1 - \rho^2)_c + 1.18 \log \text{Var}(r_{\hat{D}})_c$$

(5.24) (0.52) (0.92) (1.63)

$N = 41$ markets, $R^2 = 0.17$. Heteroskedasticity-robust standard errors in parentheses.

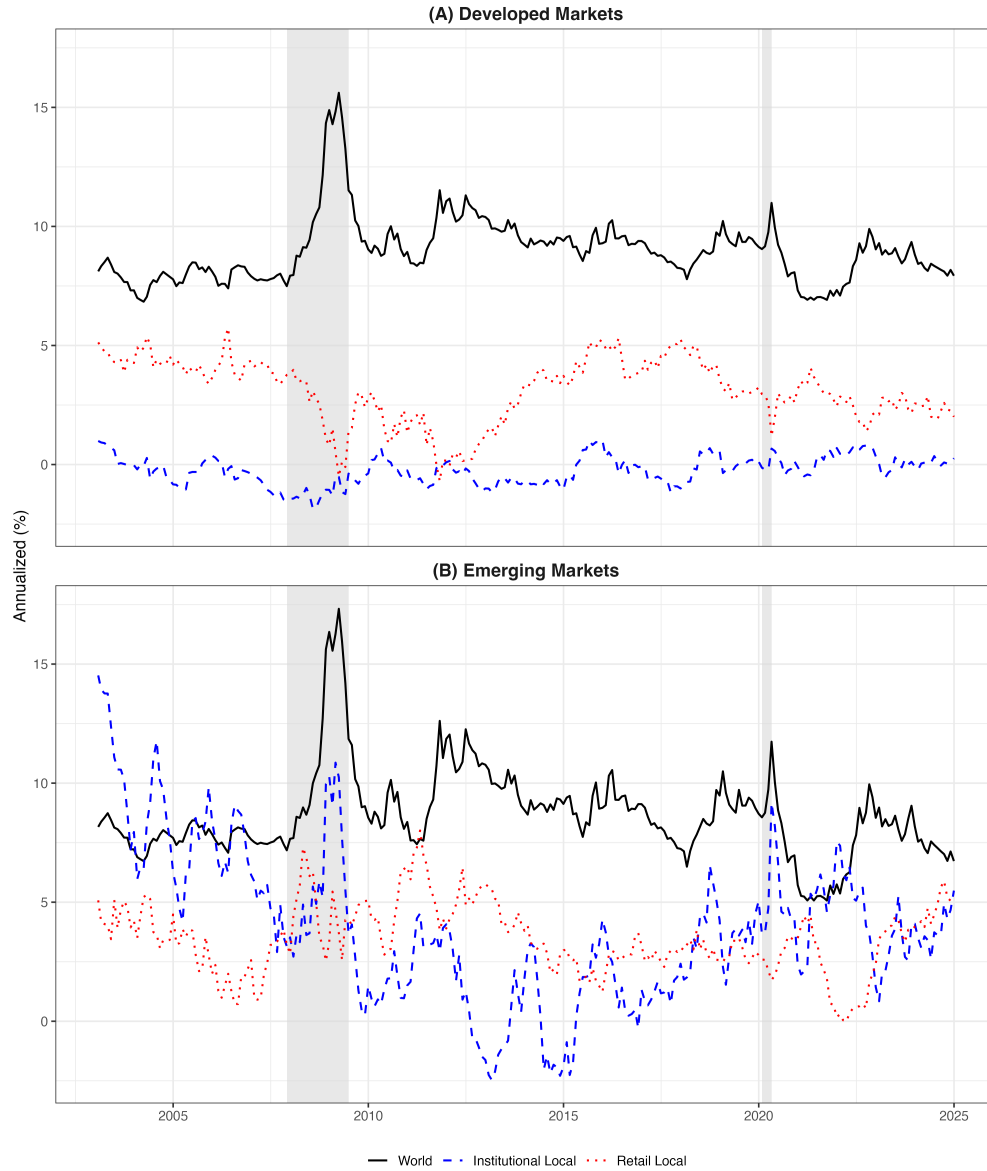


Figure 3. Conditional Risk Premia in Developed and Emerging Markets

This figure plots the equally-weighted average time-varying attainable world market premium, institutional local premium, and retail local premium in developed markets (Panel A) and emerging markets (Panel B). For each country, the time-varying risk premium is calculated using the estimated coefficients reported in Table 5 and the time-varying common instruments, as in Eq. (21), and then annualized. The sample spans January 2003 to December 2024. Shaded areas represent recession months defined by the National Bureau of Economic Research (NBER).

Internet Appendix for Institutional Investment and International Risk-sharing

(Not for publication)

This Online Appendix presents supplementary materials and results not included in the main body of the paper.

Appendix A. Proof of Model Equilibrium

This appendix provides proof of the model equilibrium. By symmetry, the focus is on interpreting the results based on domestic securities and investors, with results about foreign securities and investors introduced where necessary.

Appendix A.1. Problem setup

To simplify notation, I omit time subscripts unless necessary. Domestic and foreign securities are partitioned into four segments and I stack securities' excess returns into an $N \times 1$ vector $\mathbf{r} = [\mathbf{r}'_R, \mathbf{r}'_I, \mathbf{r}'_{I^*}, \mathbf{r}'_{R^*}]'$. The market capitalization of risky securities is exogenous and I use $\mathbf{s}_R$, $\mathbf{s}_I$, $\mathbf{s}_{I^*}$ and $\mathbf{s}_{R^*}$ to denote vectors of the total supply of risky securities in terms of their market capitalization. For example, $\mathbf{s}_I = [s_{I_1}, \dots, s_{I_{N_I}}]'$ is an $N_I \times 1$ vector of the market capitalization of domestic institutional securities.

Investor $k \in \{d, f, i\}$ has CARA utility and solves a one-period mean-variance portfolio choice problem by choosing the optimal dollar investment $\mathbf{x}^k$ of securities in their choice set C^k . Specifically, $C^d = \{R, I\}$, $C^f = \{R^*, I^*\}$ and $C^i = \{I, I^*\}$. The optimization problem investor k

solves is:

$$\begin{aligned} \max_{\mathbf{x}_X^k} E[U(W_{t+1}^k)] &= \max_{\mathbf{x}_X^k} E[-\exp(-\gamma^k W_{t+1}^k)] \\ W_{t+1}^k &= W_t^k (1 + r_f) + \sum_{X \in C^k} \mathbf{x}_X^{k'} \mathbf{r}_X, \end{aligned}$$

where γ^k is investor k 's absolute risk aversion. X represents a market segment in investor k 's choice set, for example, for the domestic retail investor, $X \in \{R, I\}$. $\mathbf{x}_X^k = [x_1, \dots, x_{N_X}]'$ is an $N_X \times 1$ vector containing investor k 's dollar investment in securities in segment X , $\mathbf{r}_X = [r_{X_1}, \dots, r_{X_{N_X}}]$ is a vector of one-period excess returns of assets in segment X , and $\mathbf{1}$ is a vector of ones with the appropriate length.

With normally distributed returns, the expected utility of investor k is:

$$E[U(W_{t+1}^k)] = -\exp \left[-\gamma^k W_t^k (1 + r_f) \right] \exp \left[-\gamma^k \sum_{X \in C^k} \mathbf{x}_X^{k'} \boldsymbol{\mu}_X + \frac{(\gamma^k)^2}{2} \sum_{X, Y \in C^k} \mathbf{x}_X^{k'} V_{XY} \mathbf{x}_Y^k \right], \quad (\text{A.1})$$

where $\boldsymbol{\mu}_X$ is a $n_X \times 1$ vector of risk premia of securities in segment X . The optimization problem in Eq. (A.1) is equivalent to:

$$\max_{\mathbf{x}_X^k} \sum_{X \in C^k} \mathbf{x}_X^{k'} \boldsymbol{\mu}_X - \frac{\gamma^k}{2} \sum_{X, Y \in C^k} \mathbf{x}_X^{k'} V_{XY} \mathbf{x}_Y^k. \quad (\text{A.2})$$

The first order condition (FOC) of each investor's portfolio choice depends on their choice sets, C_s . The FOC of the institutional investor i is:

$$V_{II} \mathbf{x}_I^i + V_{II^*} \mathbf{x}_{I^*}^i = \frac{1}{\gamma^i} \boldsymbol{\mu}_I \quad (\text{A.3})$$

$$V_{I^*I} \mathbf{x}_I^i + V_{I^*I^*} \mathbf{x}_{I^*}^i = \frac{1}{\gamma^i} \boldsymbol{\mu}_{I^*}. \quad (\text{A.4})$$

The FOC of the domestic retail investor d is:

$$V_{RR}\mathbf{s}_R + V_{RI}\mathbf{x}_I^d = \frac{1}{\gamma^d}\boldsymbol{\mu}_R \quad (\text{A.5})$$

$$V_{IR}\mathbf{s}_R + V_{II}\mathbf{x}_I^d = \frac{1}{\gamma^d}\boldsymbol{\mu}_I. \quad (\text{A.6})$$

The FOC of the foreign retail investor f :

$$V_{I^*I^*}\mathbf{x}_{I^*}^f + V_{I^*R^*}\mathbf{s}_{R^*} = \frac{1}{\gamma^f}\boldsymbol{\mu}_{I^*} \quad (\text{A.7})$$

$$V_{R^*I^*}\mathbf{x}_{I^*}^f + V_{R^*R^*}\mathbf{s}_{R^*} = \frac{1}{\gamma^f}\boldsymbol{\mu}_{R^*}, \quad (\text{A.8})$$

where I used in the FOCs of domestic and foreign retail investors the following market-clearing conditions for domestic and foreign retail securities:

$$\mathbf{x}_R^d = \mathbf{s}_R$$

$$\mathbf{x}_{R^*}^f = \mathbf{s}_{R^*}.$$

From Eq. (A.6) and Eq. (A.7), the dollar demand for institutional securities by retail investors can be expressed in terms of the risk premia of institutional securities.

$$\mathbf{x}_I^d = V_{II}^{-1} \left[\frac{1}{\gamma^d}\boldsymbol{\mu}_I - V_{IR}\mathbf{s}_R \right] \quad (\text{A.9})$$

$$\mathbf{x}_{I^*}^f = V_{I^*I^*}^{-1} \left[\frac{1}{\gamma^f}\boldsymbol{\mu}_{I^*} - V_{I^*R^*}\mathbf{s}_{R^*} \right]. \quad (\text{A.10})$$

Substituting Eq. (A.9) into Eq. (A.5) yields an expression of the risk premia of domestic retail securities in terms of the risk premia of domestic institutional securities:

$$\boldsymbol{\mu}_R = V_{RI}V_{II}^{-1}\boldsymbol{\mu}_I + \gamma^d(V_{RR} - V_{RI}V_{II}^{-1}V_{IR})\mathbf{s}_R. \quad (\text{A.11})$$

The risk premia of domestic retail securities have two components. The first term $V_{RI}V_{II}^{-1}\boldsymbol{\mu}_I$

reflects the risk premia earned by retail securities for their returns spanned by their institutional counterpart. This expression is similar to the pricing equation for noninvestable assets of De Jong and De Roon (2005) (Eq.9 of their paper). $V_{II}^{-1}V_{IR}$ is the regression coefficient of a multiple regression of domestic retail securities on domestic institutional securities. $V_{RR} - V_{RI}V_{II}^{-1}V_{IR}$ is the residual, conditional covariance among domestic retail securities that cannot be explained by this linear regression. Intuitively, the pricing of retail securities is "benchmarked" against their institutional counterpart, and the level of this risk premium depends on how well domestic institutional securities can span domestic retail securities. The second term is compensation for this residual covariance and depends on the absolute risk aversion of the retail investor γ^d .

To solve for equilibrium risk premia and portfolio holdings, I invoke the market-clearing conditions for domestic and foreign institutional securities:

$$\mathbf{x}_I^d + \mathbf{x}_I^i = \mathbf{s}_I \quad (\text{A.12})$$

$$\mathbf{x}_{I^*}^f + \mathbf{x}_{I^*}^i = \mathbf{s}_{I^*}. \quad (\text{A.13})$$

Appendix A.2. Equilibrium risk premia

Proof. Using the market-clearing conditions for institutional securities in Eq. (A.12) and Eq. (A.13), the dollar investment in institutional securities by the institutional investor is the residual market supply after subtracting the demand from retail investors:

$$\mathbf{x}_I^i = \mathbf{s}_I - V_{II}^{-1} \left[\frac{1}{\gamma^d} \boldsymbol{\mu}_I - V_{IR} \mathbf{s}_R \right] \quad (\text{A.14})$$

$$\mathbf{x}_{I^*}^i = \mathbf{s}_{I^*} - V_{I^*I^*}^{-1} \left[\frac{1}{\gamma^f} \boldsymbol{\mu}_{I^*} - V_{I^*R^*} \mathbf{s}_{R^*} \right]. \quad (\text{A.15})$$

Substituting the institutional investor's investment in institutional securities Eq. (A.14) and Eq. (A.15) into its FOC Eq. (A.3) and Eq. (A.4) yields a linear system from which I solve for the

institutional risk premia μ_I and μ_{I^*} :

$$\frac{1}{\gamma^i} \mu_I = V_{II} \left[\mathbf{s}_I - V_{II}^{-1} \left(\frac{1}{\gamma^d} \mu_I - V_{IR} \mathbf{s}_R \right) \right] + V_{II^*} \left[\mathbf{s}_{I^*} - V_{II^*}^{-1} \left(\frac{1}{\gamma^f} \mu_{I^*} - V_{I^*R} \mathbf{s}_R \right) \right] \quad (\text{A.16})$$

$$\frac{1}{\gamma^i} \mu_{I^*} = V_{I^*I} \left[\mathbf{s}_I - V_{II}^{-1} \left(\frac{1}{\gamma^d} \mu_I - V_{IR} \mathbf{s}_R \right) \right] + V_{I^*I^*} \left[\mathbf{s}_{I^*} - V_{II^*}^{-1} \left(\frac{1}{\gamma^f} \mu_{I^*} - V_{I^*R} \mathbf{s}_R \right) \right]. \quad (\text{A.17})$$

Rewrite the linear system Eq. (A.16) and Eq. (A.17) more compactly in matrix form:

$$\begin{bmatrix} \frac{1}{\gamma^D} \mathbf{I} & \frac{1}{\gamma^f} V_{II^*} V_{I^*I}^{-1} \\ \frac{1}{\gamma^d} V_{I^*I} V_{II}^{-1} & \frac{1}{\gamma^F} \mathbf{I} \end{bmatrix} \begin{bmatrix} \mu_I \\ \mu_{I^*} \end{bmatrix} = \begin{bmatrix} V_{II} \mathbf{s}_I + V_{IR} \mathbf{s}_R + V_{II^*} \mathbf{s}_{I^*} + V_{II^*} V_{I^*I}^{-1} V_{I^*R} \mathbf{s}_R \\ V_{I^*I} \mathbf{s}_I + V_{I^*R} \mathbf{s}_R + V_{I^*I^*} \mathbf{s}_{I^*} + V_{I^*I} V_{II}^{-1} V_{IR} \mathbf{s}_R \end{bmatrix}, \quad (\text{A.18})$$

where I define the aggregate effective risk aversion of investors in the domestic and foreign markets γ^D and γ^F as:

$$\frac{1}{\gamma^D} = \frac{1}{\gamma^d} + \frac{1}{\gamma^i} \quad (\text{A.19})$$

$$\frac{1}{\gamma^F} = \frac{1}{\gamma^f} + \frac{1}{\gamma^i}. \quad (\text{A.20})$$

Rearranging the linear system in Eq. (A.18) provides an expression of the risk premia of institutional securities in terms of exogenous inputs:

$$\begin{aligned} \begin{bmatrix} \mu_I \\ \mu_{I^*} \end{bmatrix} &= \begin{bmatrix} V_{II} & \mathbf{O} \\ \mathbf{O} & V_{I^*I^*} \end{bmatrix} \begin{bmatrix} \frac{1}{\gamma^D} V_{II} & \frac{1}{\gamma^f} V_{II^*} \\ \frac{1}{\gamma^d} V_{I^*I} & \frac{1}{\gamma^F} V_{I^*I^*} \end{bmatrix}^{-1} \begin{bmatrix} V_{II} \mathbf{s}_I + V_{IR} \mathbf{s}_R + V_{II^*} \mathbf{s}_{I^*} + V_{II^*} V_{I^*I}^{-1} V_{I^*R} \mathbf{s}_R \\ V_{I^*I} \mathbf{s}_I + V_{I^*R} \mathbf{s}_R + V_{I^*I^*} \mathbf{s}_{I^*} + V_{I^*I} V_{II}^{-1} V_{IR} \mathbf{s}_R \end{bmatrix} \\ &= \begin{bmatrix} \gamma^D V_{II} \Phi_I^{-1} & -\frac{\gamma^D \gamma^F}{\gamma^f} V_{II} \Phi_I^{-1} V_{II^*} V_{I^*I}^{-1} \\ -\frac{\gamma^D \gamma^F}{\gamma^d} V_{I^*I^*} \Phi_{I^*}^{-1} V_{I^*I} V_{II}^{-1} & \gamma^F V_{I^*I^*} \Phi_{I^*}^{-1} \end{bmatrix} \begin{bmatrix} V_{II} \mathbf{s}_I + V_{IR} \mathbf{s}_R + V_{II^*} \mathbf{s}_{I^*} + V_{II^*} V_{I^*I}^{-1} V_{I^*R} \mathbf{s}_R \\ V_{I^*I} \mathbf{s}_I + V_{I^*R} \mathbf{s}_R + V_{I^*I^*} \mathbf{s}_{I^*} + V_{I^*I} V_{II}^{-1} V_{IR} \mathbf{s}_R \end{bmatrix}, \end{aligned} \quad (\text{A.21})$$

where Φ_I and Φ_{I^*} are defined as follows:

$$\Phi_I = V_{II} - \frac{\gamma^D \gamma^F}{\gamma^d \gamma^f} V_{II^*} V_{I^*I}^{-1} V_{I^*I} \quad (\text{A.22})$$

$$\Phi_{I^*} = V_{I^*I^*} - \frac{\gamma^D \gamma^F}{\gamma^d \gamma^f} V_{I^*I} V_{II}^{-1} V_{II^*}. \quad (\text{A.23})$$

Φ_I is the conditional covariance among domestic institutional securities that is not explained by the span of foreign institutional securities. Unlike the conditional covariance matrices in Eq. (A.11), here the amount of conditioning is adjusted by a factor of $\frac{\gamma^D \gamma^F}{\gamma^d \gamma^f}$. The higher this ratio, the lower the risk-bearing capacity of the institutional investor relative to retail investors.^{IA-1}

To simplify the conditional covariance matrix Φ_I and Φ_{I^*} , I assume that the following approximations hold:

$$\Phi_I^{-1} \approx \theta V_{II}^{-1} \quad (\text{A.24})$$

$$\Phi_{I^*}^{-1} \approx \theta V_{I^*I^*}^{-1},$$

where $\theta > 1$ is a scalar that increases if domestic and foreign institutional securities can better substitute for each other.

The approximations hold exactly when there is only one institutional and one retail security in each country. Suppose $\text{corr}(r_I, r_{I^*}) = \rho$, then:

$$V_{II^*} V_{I^*I}^{-1} V_{I^*I} = \rho^2 V_{II}$$

$$V_{I^*I} V_{II}^{-1} V_{II^*} = \rho^2 V_{I^*I^*}.$$

$\rho^2 < 1$ captures the imperfect substitution between domestic and foreign institutional securities. In the limit case when $\rho^2 = 1$, $V_{II^*} V_{I^*I}^{-1} V_{I^*I} = V_{II}$ and $V_{I^*I} V_{II}^{-1} V_{II^*} = V_{I^*I^*}$, indicating that domestic and foreign institutional securities perfectly span each other. θ can be expressed as:

^{IA-1}To see this, note that $\frac{\gamma^D \gamma^F}{\gamma^d \gamma^f} = \frac{(\gamma^i)^2}{(\gamma^d + \gamma^i)(\gamma^f + \gamma^i)}$, which increases with the risk aversion of the institutional investor γ^i , hence decreases with the risk-bearing capacity of the institutional investor.

$$\theta = [1 - \frac{\gamma^D \gamma^F}{\gamma^d \gamma^f} \rho^2]^{-1}.$$

θ is influenced by the institutional investor's decision to trade off risk across domestic and foreign institutional securities. This is determined by both ρ^2 and $\frac{\gamma^D \gamma^F}{\gamma^d \gamma^f}$. ρ^2 measures the diversification the institutional investor enjoys by investing in both domestic and foreign institutional securities. When the absolute correlation is higher, domestic and foreign institutional securities better substitute for each other, resulting in less diversification for the institution and a higher θ . When retail investors' relative wealth with respect to the institutional investor increases, $\frac{\gamma^D \gamma^F}{\gamma^d \gamma^f}$ increases, leading to an increase in θ .

Solving for the risk premia of domestic institutional securities, using the approximation Eq. (A.24) yields:

$$\begin{aligned} \mu_I &= \gamma^D V_{II} \Phi_I^{-1} \left[(V_{II} - \frac{\gamma^F}{\gamma^f} V_{II^*} V_{I^*I}^{-1} V_{I^*I}) \mathbf{s}_I + (V_{IR} - \frac{\gamma^F}{\gamma^f} V_{II^*} V_{I^*I}^{-1} V_{I^*I} V_{II}^{-1} V_{IR}) \mathbf{s}_R + (1 - \frac{\gamma^F}{\gamma^f}) V_{II^*} \mathbf{s}_{I^*} \right. \\ &\quad \left. + (1 - \frac{\gamma^F}{\gamma^f}) V_{II^*} V_{I^*I}^{-1} V_{I^*R} \mathbf{s}_{R^*} \right] \\ &= \gamma [V_{II} \mathbf{s}_I + V_{IR} \mathbf{s}_R + V_{II^*} \mathbf{s}_{I^*} + V_{II^*} V_{I^*I}^{-1} V_{I^*R} \mathbf{s}_{R^*}] \\ &\quad + \frac{\gamma^i}{\gamma^f} \gamma [(V_{II} - V_{II^*} V_{I^*I}^{-1} V_{I^*I}) \mathbf{s}_I + (V_{IR} - V_{II^*} V_{I^*I}^{-1} V_{I^*I} V_{II}^{-1} V_{IR}) \mathbf{s}_R]. \end{aligned} \tag{A.25}$$

Substituting Eq. (A.25) into Eq. (A.11) provides the risk premia of domestic retail securities:

$$\begin{aligned} \mu_R &= \gamma V_{RI} V_{II}^{-1} [V_{II} \mathbf{s}_I + V_{IR} \mathbf{s}_R + V_{II^*} \mathbf{s}_{I^*} + V_{II^*} V_{I^*I}^{-1} V_{I^*R} \mathbf{s}_{R^*}] \\ &\quad + \frac{\gamma^i}{\gamma^f} \gamma V_{RI} V_{II}^{-1} [(V_{II} - V_{II^*} V_{I^*I}^{-1} V_{I^*I}) \mathbf{s}_I + (V_{IR} - V_{II^*} V_{I^*I}^{-1} V_{I^*I} V_{II}^{-1} V_{IR}) \mathbf{s}_R] + \gamma^d (V_{RR} - V_{RI} V_{II}^{-1} V_{IR}) \mathbf{s}_R, \end{aligned} \tag{A.26}$$

where I use the fact that $1 - \frac{\gamma^F}{\gamma^f} = \frac{\gamma^F}{\gamma^i}$. γ is the aggregate risk aversion of the economy:

$$\gamma = \frac{\gamma^D \gamma^F}{\gamma^i} \theta. \quad (\text{A.27})$$

γ is easier to interpret in the special case when there is only one institutional security in each country. Assume that the correlation between domestic and foreign institutional securities is $\text{corr}(r_I, r_{I^*}) = \rho$, the aggregate risk-bearing capacity γ is:

$$\begin{aligned} \frac{1}{\gamma} &= \frac{\gamma^i}{\gamma^D \gamma^F} \theta^{-1} \\ &= \frac{(\gamma^d + \gamma^i)(\gamma^f + \gamma^i)}{\gamma^d \gamma^f \gamma^i} \frac{\gamma^d \gamma^f + \gamma^d \gamma^i + \gamma^f \gamma^i + (1 - \rho^2)(\gamma^i)^2}{(\gamma^d + \gamma^i)(\gamma^f + \gamma^i)} \\ &= \frac{1}{\gamma^d} + \frac{1}{\gamma^f} + \frac{1}{\gamma^i} + (1 - \rho^2) \frac{\gamma^i}{\gamma^d \gamma^f}. \end{aligned} \quad (\text{A.28})$$

When $\rho^2 = 1$, the aggregate risk-bearing capacity is the total risk-bearing capacity of the three representative investors in a frictionless economy. When $\rho^2 \neq 1$, the aggregate risk-bearing capacity decreases with ρ^2 . Intuitively, as ρ^2 decreases, institutional investors enjoy more international diversification, resulting in higher aggregate risk-bearing capacity.

I define the *attainable return* of any domestic security j as the fitted value by regressing its return r_j onto domestic institutional securities:

$$\hat{r}_j = B_{jI} \mathbf{r}_I, \quad (\text{A.29})$$

where $B_{jI} = \mathbf{V}_{jI} \mathbf{V}_{II}^{-1}$ is a $1 \times N_I$ row vector of the multiple regression coefficients of regressing the return of security j onto domestic institutional securities.

I define the *attainable domestic market portfolio* $\hat{D}$ as the portfolio of domestic institutional securities that optimally mimic the domestic market portfolio. The return on this portfolio is the same as that of the value-weighted portfolio of the domestic institutional portfolio and the attainable domestic retail portfolio:

$$r_{\hat{D}} = \frac{\mathbf{s}'_I}{M_D} \mathbf{r}_I + \frac{\mathbf{s}'_R}{M_D} B_{RI} \mathbf{r}_I, \quad (\text{A.30})$$

where $B_{RI} = V_{RI} V_{II}^{-1}$ is an $n_R \times n_I$ matrix of multiple regression coefficients of retail securities on institutional securities. The *attainable foreign market portfolio* is defined in the same way. I define the value-weighted portfolio of the attainable domestic portfolio ($\hat{D}$) and the attainable foreign portfolio ($\hat{F}$) as the *attainable world market portfolio* ($\hat{W}$):

$$r_{\hat{W}} = \frac{M_D}{M_W} r_{\hat{D}} + \frac{M_F}{M_W} r_{\hat{F}}. \quad (\text{A.31})$$

The attainable world market portfolio represents the world market risk that can be shared through institutional investment.

I define the *substitute portfolio* for the attainable domestic market portfolio as the portfolio of foreign institutional securities that optimally mimics the attainable domestic market portfolio. The return on this portfolio is the fitted value from a multiple linear regression of the attainable domestic market portfolio returns onto foreign institutional securities:

$$r_{D^s} = \frac{\mathbf{s}'_I}{M_D} B_{II^*} \mathbf{r}_{I^*} + \frac{\mathbf{s}'_R}{M_D} B_{RI} B_{II^*} \mathbf{r}_{I^*}, \quad (\text{A.32})$$

where $B_{II^*} = V_{II^*} V_{I^* I^*}^{-1}$ is a $n_I \times n_{I^*}$ matrix of multiple regression coefficients of domestic institutional securities on foreign institutional securities.

From Eq. (A.25), the equilibrium risk premia of any domestic institutional security I_j can be

expressed in terms of their covariance with aggregate factor portfolios as follows:

$$\begin{aligned}
\mu_{I_j} &= \gamma[V_{I_j I} \mathbf{s}_I + V_{I_j R} \mathbf{s}_R + V_{I_j I^*} \mathbf{s}_{I^*} + V_{I_j I^*} V_{I^* I^*}^{-1} V_{I^* R^*} \mathbf{s}_{R^*}] \\
&+ \frac{\gamma^i}{\gamma^f} \gamma [(V_{I_j I} - V_{I_j I^*} V_{I^* I^*}^{-1} V_{I^* I}) \mathbf{s}_I + (V_{I_j R} - V_{I_j I^*} V_{I^* I^*}^{-1} V_{I^* I} V_{I^* I}^{-1} V_{I^* R}) \mathbf{s}_R] \\
&= \gamma M_D \text{Cov}(r_{I_j}, \frac{\mathbf{s}'_I}{M_D} \mathbf{r}_I + \frac{\mathbf{s}'_R}{M_D} \mathbf{r}_R) + \gamma M_F \text{Cov}(r_{I_j}, \frac{\mathbf{s}'_{I^*}}{M_F} \mathbf{r}_{I^*} + \frac{\mathbf{s}'_{R^*}}{M_F} B_{R^* I^*} \mathbf{r}_{I^*}) \\
&+ \frac{\gamma^i}{\gamma^f} \gamma M_D \text{Cov}(r_{I_j}, \frac{\mathbf{s}'_I}{M_D} \mathbf{r}_I + \frac{\mathbf{s}'_R}{M_D} \mathbf{r}_R - \frac{\mathbf{s}'_I}{M_D} B_{I I^*} \mathbf{r}_{I^*} - \frac{\mathbf{s}'_R}{M_D} B_{R I} B_{I I^*} \mathbf{r}_{I^*}) \\
&= \gamma M_W \text{Cov}(r_{I_j}, r_{\hat{W}}) + \frac{\gamma^i}{\gamma^f} \gamma M_D \text{Cov}(r_{I_j}, r_{\hat{D}} - r_{D^s}),
\end{aligned} \tag{A.33}$$

where I use the fact that $\text{Cov}(r_I, r_D) = \text{Cov}(r_I, r_{\hat{D}})$. Domestic institutional securities earn two risk premia. The first one is an attainable world risk premium for their covariance with the attainable world market return $r_{\hat{W}}$, which I also refer to as the *attainable world market factor* $f^{\hat{W}}$. The second one is for their covariance with the return difference between the attainable domestic market portfolio and its substitute portfolio of foreign securities, which I define as the *institutional local factor*:

$$f^{inst} = r_{\hat{D}} - r_{D^s}. \tag{A.34}$$

From Eq. (A.26), the equilibrium risk premia of any domestic retail security R_j can be expressed similarly in terms of its covariances with factor portfolios:

$$\mu_{R_j} = \gamma M_W \text{Cov}(\hat{r}_{R_j}, r_{\hat{W}}) + \frac{\gamma^i}{\gamma^f} \gamma M_D \text{Cov}(\hat{r}_{R_j}, r_{\hat{D}} - r_{D^s}) + \gamma^d M_R \text{Cov}(r_{R_j}, r_R - \hat{r}_R), \tag{A.35}$$

where $\hat{r}_{R_j}$ is the attainable return of security R_j defined in Eq. (A.29).

Domestic retail securities earn a third risk premium for their covariance with the residual retail risk that is orthogonal to the span of institutional securities, which I define as the *retail local*

factor:

$$f^{retail} = r_R - \hat{r}_R. \quad (\text{A.36})$$

The risk premium of any domestic security j can be expressed as the following generic expression:

$$\mu_j = \gamma M_W \text{Cov}(\hat{r}_j, r_{\hat{W}}) + \frac{\gamma^i}{\gamma^f} \gamma M_D \text{Cov}(\hat{r}_j, r_{\hat{D}} - r_{D^s}) + \gamma^d M_R \text{Cov}(r_j, r_R - \hat{r}_R), \quad (\text{A.37})$$

where I used the fact that for institutional securities, $\text{Cov}(r_{I_j}, r_R - \hat{r}_R) = 0$, because $r_R - \hat{r}_R$ is the residual from a multiple regression of r_R on the set of all domestic institutional securities.

□

The price of the attainable world market risk γM_W is the same across all four segments. The price of the institutional local risk is $\frac{\gamma^i}{\gamma^f} \gamma M_D$, which increases when the institutional investor becomes more risk averse, resulting in less international risk-sharing, hence a higher institutional local premium relative to the attainable world premium. γ^i also has a direct level effect on the overall level of risk premium through the aggregate risk aversion γ , as can be seen from Eq. (A.28).

Consider the special case in which there is only one institutional security and one retail security in each country. Suppose the correlation between domestic and foreign institutional securities is $\text{corr}(r_I, r_{I^*}) = \rho$, the correlation between domestic retail and institutional securities is $\text{corr}(r_R, r_I) = \rho_R$, and the correlation between foreign retail and institutional securities is $\text{corr}(r_{R^*}, r_{I^*}) = \rho_R^*$. For any domestic security j , suppose its volatility is σ_j and its correlation with the domestic institutional security is $\text{corr}(r_j, r_I) = \rho_j$. Then its attainable return is $\hat{r}_j = \rho_j \frac{\sigma_j}{\sigma_I} r_I$. Simplifying Eq. (A.37), the risk premium of a domestic security $j \in \{R, I\}$ can be expressed as:

$$\mu_j = \gamma M_W \rho_j \frac{\sigma_j}{\sigma_I} \text{Cov}(r_I, r_{\hat{W}}) + \gamma \frac{\gamma^i}{\gamma^f} M_D \rho_j \frac{\sigma_j}{\sigma_I} (1 - \rho^2) \text{Cov}(r_I, r_{\hat{D}}) + \gamma^d M_R (1 - \rho_j^2) \sigma_j \sigma_R. \quad (\text{A.38})$$

In this simplified case, domestic security j earns attainable world and institutional local risk premia through its correlation with the domestic institutional security, ρ_j . The novelty in the model is the indirect covariance between domestic and foreign retail securities embedded in the attainable world risk premium of the domestic retail security: $\text{Cov}(\hat{r}_R, \hat{r}_{R^*}) = \rho_R \rho \rho_R^* \sigma_R \sigma_{R^*}$. These two remote markets do not share a single common investor yet share risk through a chain of correlations: the correlation between domestic retail and institutional securities ρ_R , the correlation between domestic and foreign institutional securities ρ , and the correlation between foreign institutional and retail securities ρ_R^* .

The relative importance of global versus local risk premia depends on three aspects of institutional investment:

- *Coverage* of the institution's choice set, represented by ρ_j .
- *Correlation* between domestic and foreign institutional securities ρ , which determines the quantity of risk that can be shared internationally by trading institutional securities.
- *Capacity*, the risk-bearing capacity of the institutional investor, $1/\gamma^i$, which determines the level of the institutional local risk premium.

The correlation between a domestic security and the domestic institutional security reflects to what extent its exposure can be spanned by the institutional security. Since the retail local risk premium is compensation for the unspanned retail risk, the higher the correlation, the lower the retail local risk premium. $\rho_j = 1$ when either j is an institutional security or a retail security that is perfectly correlated with the domestic institutional security. In this case, the retail local risk premium becomes zero. Notably, the correlation between a domestic retail security and foreign securities does not enter the equilibrium risk premium because international risk-sharing in the model is channeled through its correlation with the domestic institutional security.

Beta representation Define the univariate beta exposure of security j with respect to the attainable world factor $f^{\hat{W}}$, the institutional local factor f^{inst} , and the retail local factor f^{retail} as follows:

$$\tilde{\beta}_j^{\hat{W}} = \frac{\text{Cov}(\hat{r}_j, f^{\hat{W}})}{\text{Var}(f^{\hat{W}})} \quad (\text{A.39})$$

$$\tilde{\beta}_j^{inst} = \frac{\text{Cov}(\hat{r}_j, f^{inst})}{\text{Var}(f^{inst})} \quad (\text{A.40})$$

$$\beta_j^{retail} = \frac{\text{Cov}(r_j, f^{retail})}{\text{Var}(f^{retail})}, \quad (\text{A.41})$$

and the risk premium associated with each beta exposure as:

$$\mu^{\hat{W}} = \gamma M_W \text{Var}(f^{\hat{W}}) \quad (\text{A.42})$$

$$\mu^{inst} = \frac{\gamma^i}{\gamma^f} \gamma M_D \text{Var}(f^{inst}) \quad (\text{A.43})$$

$$\mu^{retail} = \gamma^d M_R \text{Var}(f^{retail}). \quad (\text{A.44})$$

The risk premium of domestic security j in Eq. (A.37) has the following equivalent beta representation:

$$\mu_j = \tilde{\beta}_j^{\hat{W}} \mu^{\hat{W}} + \tilde{\beta}_j^{inst} \mu^{inst} + \beta_j^{retail} \mu^{retail}. \quad (\text{A.45})$$

Appendix A.3. Equilibrium investments

Substituting Eq. (A.25) into the FOC of the domestic retail investor Eq. (A.9), I obtain the holdings of domestic institutional securities by the domestic retail investor:

$$\begin{aligned} \mathbf{x}_I^d = & \frac{\gamma^D}{\gamma^d} \Phi_I^{-1} \left[\left(V_{II} - \frac{\gamma^F}{\gamma^f} V_{II^*} V_{I^*I^*}^{-1} V_{I^*I} \right) \mathbf{s}_I + \left(V_{IR} - \frac{\gamma^F}{\gamma^f} V_{II^*} V_{I^*I^*}^{-1} V_{I^*I} V_{II}^{-1} V_{IR} \right) \mathbf{s}_R + \left(1 - \frac{\gamma^F}{\gamma^f} \right) V_{II^*} \mathbf{s}_{I^*} \right. \\ & \left. + \left(1 - \frac{\gamma^F}{\gamma^f} \right) V_{II^*} V_{I^*I^*}^{-1} V_{I^*R^*} \mathbf{s}_{R^*} \right] - V_{II}^{-1} V_{IR} \mathbf{s}_R. \end{aligned} \quad (\text{A.46})$$

Using the approximation in Eq. (A.24), I simplify the equilibrium investment of the domestic retail investor in domestic institutional securities in Eq. (A.46):

$$\begin{aligned} \mathbf{x}_I^d = & \frac{\gamma^D}{\gamma^d} \theta \underbrace{\left[\mathbf{s}_I + \mathbf{V}_{II}^{-1} \mathbf{V}_{IR} \mathbf{s}_R \right]}_{\hat{D}} + \frac{\gamma^D \gamma^F}{\gamma^d \gamma^i} \theta \underbrace{\left[\mathbf{V}_{II}^{-1} \mathbf{V}_{II^*} \mathbf{s}_{I^*} + \mathbf{V}_{II}^{-1} \mathbf{V}_{II^*} \mathbf{V}_{I^* I^*}^{-1} \mathbf{V}_{I^* R^*} \mathbf{s}_{R^*} \right]}_{F^s} - \underbrace{\mathbf{V}_{II}^{-1} \mathbf{V}_{IR} \mathbf{s}_R}_{\hat{R}} \\ & - \frac{\gamma^D \gamma^F}{\gamma^d \gamma^f} \theta \underbrace{\left[\mathbf{V}_{II}^{-1} \mathbf{V}_{II^*} \mathbf{V}_{I^* I^*}^{-1} \mathbf{V}_{I^* I} \mathbf{s}_I + \mathbf{V}_{II}^{-1} \mathbf{V}_{II^*} \mathbf{V}_{I^* I^*}^{-1} \mathbf{V}_{I^* I} \mathbf{V}_{II}^{-1} \mathbf{V}_{IR} \mathbf{s}_R \right]}_H, \end{aligned} \quad (\text{A.47})$$

where I use the fact that $1 - \frac{\gamma^F}{\gamma^f} = \frac{\gamma^F}{\gamma^i}$.

Define constants A^d , A^f , γ as:

$$A^d = \frac{\gamma^D}{\gamma^d} \theta \quad (\text{A.48})$$

$$A^f = \frac{\gamma^F}{\gamma^f} \theta, \quad (\text{A.49})$$

where A^d (A^f) measures the relative risk-bearing capacity of the domestic (foreign) retail investor with respect to the institutional investor.

By market clearing for the domestic retail securities, the domestic retail investor holds the domestic retail portfolio:

$$\mathbf{x}_R^d = \mathbf{s}_R. \quad (\text{A.50})$$

The domestic retail investor's dollar investment in domestic securities can be expressed in terms of its four components:

$$\mathbf{x}_I^d + \mathbf{s}_R^d = \underbrace{A^d \hat{D}}_{\text{attainable domestic}} + \underbrace{D - \hat{D}}_{\text{unattainable local risk}} + \underbrace{\frac{\gamma}{\gamma^d} F^s}_{\text{substitute for foreign}} - \frac{\gamma^i}{\gamma^d \gamma^f} \gamma H, \quad (\text{A.51})$$

where H is a portfolio of domestic institutional securities that optimally mimics the substitute

portfolio for domestic investment D^s . The institutional investor tilts away from D^s in its trade with foreign retail investors.^{IA-2} Portfolio H serves as a hedge that the domestic retail investor provides to the institutional investor for that tilt.

The size of this position depends on the level of correlation between domestic and foreign institutional securities. To see this, consider the special case in which there is only one institutional security and one retail security in each country. Assuming that the correlation between the domestic institutional security and the foreign institutional security is $\text{corr}(r_I, r_{I^*}) = \rho$, H can be simplified as:

$$H = \rho^2 \hat{D}. \quad (\text{A.52})$$

Unless the correlation between domestic and foreign institutional securities is very high, the magnitude of this term is very small compared to the other components of investors' portfolios.

Using the market clearing condition for domestic institutional securities, the dollar investment in domestic institutional securities by the institutional investor is:

$$\mathbf{x}_I^i = (1 - A^d) \hat{D} - \frac{\gamma}{\gamma^d} F^s + \frac{\gamma^i}{\gamma^d \gamma^f} \gamma H. \quad (\text{A.53})$$

Figure IA.1 shows the portfolio frontiers of the domestic retail investor and the institutional investor. Unlike existing partial segmentation models in which the frontier of an unrestricted investor is superior to that of a restricted investor, the portfolio frontiers of the two investors intersect.

^{IA-2}Because the foreign retail investor tilts toward and the institutional investor tilts away from D^s , it is as if the institution "sells" the substitute portfolio to the foreign retail investor.

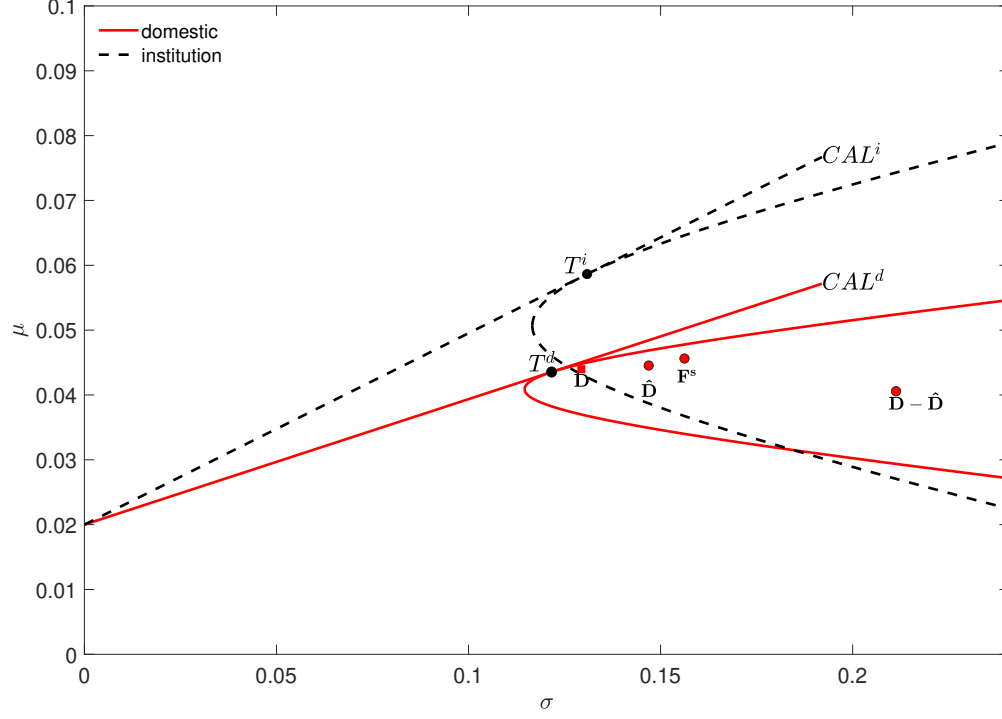


Figure IA.1. Efficient frontiers of the domestic retail investor and the institutional investor

This figure plots the efficient frontiers of the domestic retail investor and the institutional investor. T^i and T^d are the tangency portfolios of the institutional investor and the domestic retail investor, respectively. CAL^i and CAL^d are the capital allocation lines of the two investors.

Appendix B. Special Cases and Additional Analyses

Appendix B.1. Special cases under alternative model assumptions

This appendix discusses the equilibrium risk premia of different assets under alternative assumptions about the choice sets of institutional investors and retail investors.

S1: Institutional investors fully diversify Suppose the institutional investor's choice set is unrestricted. Define the vector of domestic (foreign) excess returns as $\mathbf{r}_D = [\mathbf{r}'_R, \mathbf{r}'_I]'$ ($\mathbf{r}_F = [\mathbf{r}'_{I^*}, \mathbf{r}'_{R^*}]'$). The supply of domestic (foreign) securities in terms of market capitalization is denoted as $\mathbf{s}_D = [\mathbf{s}'_R, \mathbf{s}'_I]'$ ($\mathbf{s}_F = [\mathbf{s}'_{I^*}, \mathbf{s}'_{R^*}]'$). Denote the covariance matrix between domestic and foreign

securities as:

$$\begin{bmatrix} V_{DD} & V_{DF} \\ V_{FD} & V_{FF} \end{bmatrix} = \begin{bmatrix} V_{RR} & V_{RI} & V_{RI^*} & V_{RR^*} \\ V_{IR} & V_{II} & V_{II^*} & V_{IR^*} \\ V_{I^*R} & V_{I^*I} & V_{I^*I^*} & V_{I^*R^*} \\ V_{R^*R} & V_{R^*I} & V_{R^*I^*} & V_{R^*R^*} \end{bmatrix}.$$

The FOCs of domestic and foreign retail investors remain the same and can be rewritten as:

$$V_{DD}\mathbf{x}_D^d = \frac{1}{\gamma^d}\boldsymbol{\mu}_D \quad (\text{B.1})$$

$$V_{FF}\mathbf{x}_F^f = \frac{1}{\gamma^f}\boldsymbol{\mu}_F. \quad (\text{B.2})$$

The FOC of the institutional investor becomes:

$$V_{DD}\mathbf{x}_D^i + V_{DF}\mathbf{x}_F^i = \frac{1}{\gamma^i}\boldsymbol{\mu}_D \quad (\text{B.3})$$

$$V_{FD}\mathbf{x}_D^i + V_{FF}\mathbf{x}_F^i = \frac{1}{\gamma^i}\boldsymbol{\mu}_F. \quad (\text{B.4})$$

From Eq. (B.1) and Eq. (B.2) solve the demand by domestic and foreign retail investors:

$$\mathbf{x}_D^d = \frac{1}{\gamma^d}V_{DD}^{-1}\boldsymbol{\mu}_D \quad (\text{B.5})$$

$$\mathbf{x}_F^f = \frac{1}{\gamma^f}V_{FF}^{-1}\boldsymbol{\mu}_F. \quad (\text{B.6})$$

Utilizing the market-clearing conditions of domestic and foreign assets, and Eq. (B.5) and Eq. (B.6), the holdings by institutional investors are:

$$\mathbf{x}_D^i = \mathbf{s}_D - \frac{1}{\gamma^d}V_{DD}^{-1}\boldsymbol{\mu}_D \quad (\text{B.7})$$

$$\mathbf{x}_F^i = \mathbf{s}_F - \frac{1}{\gamma^f}V_{FF}^{-1}\boldsymbol{\mu}_F. \quad (\text{B.8})$$

Substituting Eq. (B.7) and Eq. (B.8) into the FOCs Eq. (B.3) and Eq. (B.4) yields the following linear system involving the risk premium of domestic and foreign securities:

$$\begin{aligned} V_{DD}\mathbf{s}_D - \frac{1}{\gamma^d}\boldsymbol{\mu}_D + V_{DF}\mathbf{s}_F - \frac{1}{\gamma^f}V_{DF}V_{FF}^{-1}\boldsymbol{\mu}_F &= \frac{1}{\gamma^i}\boldsymbol{\mu}_D \\ V_{FD}\mathbf{s}_D - \frac{1}{\gamma^d}V_{FD}V_{DD}^{-1}\boldsymbol{\mu}_D + V_{FF}\mathbf{s}_F - \frac{1}{\gamma^f}\boldsymbol{\mu}_F &= \frac{1}{\gamma^i}\boldsymbol{\mu}_F. \end{aligned} \quad (\text{B.9})$$

The above linear system can be rewritten as:

$$\begin{aligned} \left(\frac{1}{\gamma^i} + \frac{1}{\gamma^d}\right)\boldsymbol{\mu}_D + \frac{1}{\gamma^f}V_{DF}V_{FF}^{-1}\boldsymbol{\mu}_F &= V_{DD}\mathbf{s}_D + V_{DF}\mathbf{s}_F \\ \left(\frac{1}{\gamma^i} + \frac{1}{\gamma^f}\right)\boldsymbol{\mu}_F + \frac{1}{\gamma^d}V_{FD}V_{DD}^{-1}\boldsymbol{\mu}_D &= V_{FD}\mathbf{s}_D + V_{FF}\mathbf{s}_F. \end{aligned} \quad (\text{B.10})$$

Define the aggregate absolute risk aversion of investors in the domestic and foreign markets:

$$\begin{aligned} \frac{1}{\gamma^D} &= \frac{1}{\gamma^i} + \frac{1}{\gamma^d} \\ \frac{1}{\gamma^F} &= \frac{1}{\gamma^i} + \frac{1}{\gamma^f}. \end{aligned}$$

Rewrite the linear system in matrix form

$$\begin{bmatrix} \frac{1}{\gamma^D}V_{DD} & \frac{1}{\gamma^F}V_{DF} \\ \frac{1}{\gamma^D}V_{FD} & \frac{1}{\gamma^F}V_{FF} \end{bmatrix} \begin{bmatrix} V_{DD}^{-1} & \mathbf{O} \\ \mathbf{O} & V_{FF}^{-1} \end{bmatrix} \begin{bmatrix} \boldsymbol{\mu}_D \\ \boldsymbol{\mu}_F \end{bmatrix} = \begin{bmatrix} V_{DD}\mathbf{s}_D + V_{DF}\mathbf{s}_F \\ V_{FD}\mathbf{s}_D + V_{FF}\mathbf{s}_F \end{bmatrix}. \quad (\text{B.11})$$

Solve the linear system by inverting the block matrices

$$\begin{aligned} \begin{bmatrix} \boldsymbol{\mu}_D \\ \boldsymbol{\mu}_F \end{bmatrix} &= \begin{bmatrix} V_{DD} & \mathbf{O} \\ \mathbf{O} & V_{FF} \end{bmatrix} \begin{bmatrix} \frac{1}{\gamma^D}V_{DD} & \frac{1}{\gamma^F}V_{DF} \\ \frac{1}{\gamma^D}V_{FD} & \frac{1}{\gamma^F}V_{FF} \end{bmatrix}^{-1} \begin{bmatrix} V_{DD}\mathbf{s}_D + V_{DF}\mathbf{s}_F \\ V_{FD}\mathbf{s}_D + V_{FF}\mathbf{s}_F \end{bmatrix} \\ &= \begin{bmatrix} V_{DD} & \mathbf{O} \\ \mathbf{O} & V_{FF} \end{bmatrix} \begin{bmatrix} \gamma^D\Phi_D^{-1} & -\frac{\gamma^D\gamma^F}{\gamma^f}\Phi_D^{-1}V_{DF}V_{FF}^{-1} \\ -\frac{\gamma^D\gamma^F}{\gamma^d}\Phi_F^{-1}V_{FD}V_{DD}^{-1} & \gamma^F\Phi_F^{-1} \end{bmatrix} \begin{bmatrix} V_{DD}\mathbf{s}_D + V_{DF}\mathbf{s}_F \\ V_{FD}\mathbf{s}_D + V_{FF}\mathbf{s}_F \end{bmatrix}, \end{aligned} \quad (\text{B.12})$$

where the matrices Φ_D and Φ_F are defined as

$$\Phi_D = V_{DD} - \frac{\gamma^D \gamma^F}{\gamma^d \gamma^f} V_{DF} V_{FF}^{-1} V_{FD} \quad (\text{B.13})$$

$$\Phi_F = V_{FF} - \frac{\gamma^D \gamma^F}{\gamma^d \gamma^f} V_{FD} V_{DD}^{-1} V_{DF}. \quad (\text{B.14})$$

Φ_D (Φ_F) can be interpreted as the conditional variance of domestic (foreign) assets not spanned by foreign (domestic) assets. The amount of conditioning is modulated by the ratio $\frac{\gamma^D \gamma^F}{\gamma^d \gamma^f}$. Without loss of generality, assume that all investors have the same relative risk aversion $\alpha = W^k \gamma^k$. Define the aggregate wealth of investors investing in the domestic market as $W^D = W^d + W^i$ and the aggregate wealth of investors investing in the foreign market as $W^F = W^f + W^i$, then

$$\begin{aligned} \gamma^D &= \frac{\alpha}{W^D} \\ \gamma^F &= \frac{\alpha}{W^F} \\ \frac{\gamma^D \gamma^F}{\gamma^d \gamma^f} &= \frac{W^d W^f}{W^D W^F}. \end{aligned}$$

Intuitively, the ratio represents the relative importance of retail investors in both markets. The more important retail investors are, the more the spanning between domestic and foreign securities matters because only retail investors need to replicate exposures in the other country using their home assets.

In order to further simplify Eq. (B.12), I assume the following approximation:

$$\Phi_D^{-1} \approx \theta V_{DD}^{-1} \quad (\text{B.15})$$

$$\Phi_F^{-1} \approx \theta V_{FF}^{-1}. \quad (\text{B.16})$$

When there is only one domestic and one foreign security, the approximation holds exactly.

Suppose that the correlation between the domestic and foreign security is $\text{corr}(r_D, r_F) = \rho$, then

$$\begin{aligned}\Phi_D &= (1 - \frac{\gamma^D \gamma^F}{\gamma^d \gamma^f} \rho^2) V_{DD} \\ \Phi_D^{-1} &= [1 - \frac{\gamma^D \gamma^F}{\gamma^d \gamma^f} \rho^2]^{-1} V_{DD}^{-1} \\ \Phi_F &= (1 - \frac{\gamma^D \gamma^F}{\gamma^d \gamma^f} \rho^2) V_{FF} \\ \Phi_F^{-1} &= [1 - \frac{\gamma^D \gamma^F}{\gamma^d \gamma^f} \rho^2]^{-1} V_{FF}^{-1},\end{aligned}$$

in this case, $\theta = [1 - \frac{\gamma^D \gamma^F}{\gamma^d \gamma^f} \rho^2]^{-1}$.

Applying the approximation Eq. (B.15) to further simplify the linear system Eq. (B.12) and solve for the risk premium of domestic securities:

$$\begin{aligned}\mu_D &= \gamma^D \theta (V_{DD} \mathbf{s}_D + V_{DF} \mathbf{s}_F) - \frac{\gamma^D \gamma^F}{\gamma^f} \theta V_{DF} V_{FF}^{-1} (V_{FD} \mathbf{s}_D + V_{FF} \mathbf{s}_F) \\ &= \gamma^D \theta \left[(V_{DD} - \frac{\gamma^F}{\gamma^f} V_{DF} V_{FF}^{-1} V_{FD}) \mathbf{s}_D + (V_{DF} - \frac{\gamma^F}{\gamma^f} V_{DF}) \mathbf{s}_F \right] \\ &= \gamma^D \theta \left[(1 - \frac{\gamma^F}{\gamma^f}) (V_{DD} \mathbf{s}_D + V_{DF} \mathbf{s}_F) + \frac{\gamma^F}{\gamma^f} (V_{DD} - V_{DF} V_{FF}^{-1} V_{FD}) \mathbf{s}_D \right] \\ &= \frac{\gamma^D \gamma^F}{\gamma^i} \theta \left[(V_{DD} \mathbf{s}_D + V_{DF} \mathbf{s}_F) + \frac{\gamma^i}{\gamma^f} (V_{DD} - V_{DF} V_{FF}^{-1} V_{FD}) \mathbf{s}_D \right],\end{aligned}\tag{B.17}$$

where I used $1 - \frac{\gamma^F}{\gamma^f} = \frac{\gamma^i}{\gamma^f}$.

Define aggregate risk aversion γ as:

$$\gamma = \frac{\gamma^D \gamma^F}{\gamma^i} \theta.$$

The above pricing result can be expressed in terms of covariance with the world market portfolio and with the domestic market portfolio:

$$\mu_D = \gamma M_W \text{Cov}(\mathbf{r}_D, r_W) + \frac{\gamma^i}{\gamma^f} \gamma M_D \text{Cov}(\mathbf{r}_D, r_D - r_{D^s}),\tag{B.18}$$

where the excess return of the substitute portfolio for domestic securities is:

$$r_{D^s} = \frac{\mathbf{s}'_D}{M_D} \mathbf{V}_{DF} \mathbf{V}_{FF}^{-1} \mathbf{r}_F.$$

By symmetry, the risk premium of foreign securities is:

$$\mu_F = \frac{\gamma^D \gamma^F}{\gamma^i} \theta \left[(\mathbf{V}_{FD} \mathbf{s}_D + \mathbf{V}_{FF} \mathbf{s}_F) + \frac{\gamma^i}{\gamma^d} (\mathbf{V}_{FF} - \mathbf{V}_{FD} \mathbf{V}_{DD}^{-1} \mathbf{V}_{DF}) \mathbf{s}_F \right]. \quad (\text{B.19})$$

Compared with the full model, when institutions fully diversify, the retail local risk premium disappears, as institutional and retail investors in each country commonly access all local securities in the market.

S2: Retail investors invest in foreign institutional securities In this scenario, retail investors' home bias is relaxed, and their choice set includes institutional securities from the other country. Institutional securities then become globally investable securities in Alexander et al. (1987), De Jong and De Roon (2005), and Chaieb and Errunza (2007). Retail securities, by contrast, correspond to noninvestable securities held only by local investors.

Define the set of domestic and foreign institutional securities as eligible securities. The vector of eligible securities excess returns is $\mathbf{r}_E = [\mathbf{r}'_I, \mathbf{r}'_{I^*}]'$. The supply of eligible securities in terms of market capitalization is denoted as $\mathbf{s}_E = [\mathbf{s}'_I, \mathbf{s}'_{I^*}]'$. The covariance matrices between eligible and domestic and foreign retail securities are

$$\mathbf{V}_{ER} = \begin{bmatrix} \mathbf{V}_{IR} \\ \mathbf{V}_{I^*R} \end{bmatrix} \quad \mathbf{V}_{ER^*} = \begin{bmatrix} \mathbf{V}_{IR^*} \\ \mathbf{V}_{I^*R^*} \end{bmatrix}.$$

The FOC of domestic retail investors is now:

$$V_{RR}\mathbf{s}_R + V_{RE}\mathbf{x}_E^d = \frac{1}{\gamma^d}\boldsymbol{\mu}_R \quad (\text{B.20})$$

$$V_{ER}\mathbf{s}_R + V_{EE}\mathbf{x}_E^d = \frac{1}{\gamma^d}\boldsymbol{\mu}_E. \quad (\text{B.21})$$

The FOC of foreign retail investors is now:

$$V_{EE}\mathbf{x}_E^f + V_{ER^*}\mathbf{s}_{R^*} = \frac{1}{\gamma^f}\boldsymbol{\mu}_E \quad (\text{B.22})$$

$$V_{R^*E}\mathbf{x}_E^f + V_{R^*R^*}\mathbf{s}_{R^*} = \frac{1}{\gamma^f}\boldsymbol{\mu}_{R^*}. \quad (\text{B.23})$$

The FOC of institutional investors remains the same:

$$V_{EE}\mathbf{x}_E^i = \frac{1}{\gamma^i}\boldsymbol{\mu}_E. \quad (\text{B.24})$$

Summing the FOCs for investment in eligible securities, Eq. (B.21), Eq. (B.22) and Eq. (B.24), and applying the market-clearing condition for eligible securities yields:

$$\begin{aligned} \boldsymbol{\mu}_E &= \gamma[V_{ER}\mathbf{s}_R + V_{EE}\mathbf{s}_E + V_{ER^*}\mathbf{s}_{R^*}] \\ &= \gamma M_W \text{Cov}(\mathbf{r}_E, r_W), \end{aligned} \quad (\text{B.25})$$

where $\frac{1}{\gamma} = \frac{1}{\gamma^d} + \frac{1}{\gamma^i} + \frac{1}{\gamma^f}$, and $r_W = \frac{1}{M_W} \left[\mathbf{s}'_R \mathbf{r}_R + \mathbf{s}'_E \mathbf{r}_E + \mathbf{s}'_{R^*} \mathbf{r}_{R^*} \right]$ is the return of the *world market portfolio of all assets*. Eligible securities are priced globally by their covariance with the world market portfolio, the same as under full market integration, consistent with the intuition that risk-sharing is perfect in these securities. All investors are marginal in the eligible segment, so the aggregate risk tolerance in Eq. (B.25) is that of the whole economy. In the full model, foreign retail investors do not participate in domestic institutional securities, and it is this exclusion that generates the institutional local premium.

Eq. (B.25) is the same result for globally accessible securities in Alexander et al. (1987) and

Karolyi and Wu (2018) and that of eligible securities in Chaieb and Errunza (2007) in the absence of PPP deviation, with the addition of institutional investors who only invest in the investable securities.

Eq. (B.25) is also equivalent to the pricing of investable assets in equation (2) of De Jong and De Roon (2005), who assume multiple markets but only one investable asset and one non-investable asset within each market. The current special case (S2) considers an alternative setting with only two markets, each having multiple investable and noninvestable assets. To see the equivalence, rewrite Eq. (B.25) using an alternative benchmark, the *world portfolio of eligible/investable securities*:

$$\begin{aligned}\mu_E &= \gamma M_W \left[V_{EE} \frac{s_E}{M_E} \frac{M_E}{M_W} \right] + \gamma M_W \left[V_{ER} \frac{s_R}{M_R} \frac{M_R}{M_W} + V_{ER^*} \frac{s_{R^*}}{M_{R^*}} \frac{M_{R^*}}{M_W} \right] \\ &= \gamma M_W \text{Cov}(\mathbf{r}_E, r_{WE})(1 - q - q^*) + \gamma M_W \text{Cov}(\mathbf{r}_E, r_R)q + \gamma M_W \text{Cov}(\mathbf{r}_E, r_{R^*})q^*,\end{aligned}\tag{B.26}$$

where $q = \frac{M_R}{M_W}$ and $q^* = \frac{M_{R^*}}{M_W}$ are the ratios of the market capitalization of non-investable securities in domestic and foreign markets to the market value of all assets, M_E is the aggregate market capitalization of eligible securities, so that $M_W = M_R + M_E + M_{R^*}$, $r_{WE} = \frac{s'_E}{M_E} \mathbf{r}_E$ is the return on the world market portfolio of investable securities (denoted r^w in De Jong and De Roon (2005), which is the return on a subportfolio of the all-asset world market portfolio), and γM_W corresponds to their global aggregate relative risk aversion γ^m .

In their equation (2), which corresponds to Eq. (B.26), De Jong and De Roon (2005) take the first term, the *world investable market premium*, as the benchmark, and interpret the remaining terms as evidence that “the pricing of the investable assets is also affected by the investment restrictions” (p. 586). This reading is benchmark-dependent. Collecting the three terms recovers Eq. (B.25), which is the relation that prices every security under full market integration. The last two terms in Eq. (B.26) are the contribution of non-investable assets to the all-asset (rather than the investable) world market portfolio, so their economic source is the composition of the world market portfolio rather than segmentation. They are therefore not local risk premia, and investable securities are

priced by their covariance with the all-asset world market portfolio whether there is segmentation or not.

The pricing of retail securities is also the same as that of inaccessible/non-investable securities in Alexander et al. (1987), De Jong and De Roon (2005), and Chaieb and Errunza (2007). Focusing on the domestic side since the market structure is symmetric, first substitute Eq. (B.25) into the FOC Eq. (B.21) to solve for the investment in eligible securities by the domestic retail investor:

$$\mathbf{x}_E^d = \frac{1}{\gamma^d} \mathbf{V}_{EE}^{-1} \boldsymbol{\mu}_E - \mathbf{V}_{EE}^{-1} \mathbf{V}_{ER} \mathbf{s}_R. \quad (\text{B.27})$$

Substituting Eq. (B.27) into Eq. (B.20) yields the risk premium of retail securities:

$$\boldsymbol{\mu}_R = \mathbf{V}_{RE} \mathbf{V}_{EE}^{-1} \boldsymbol{\mu}_E + \gamma^d \left[\mathbf{V}_{RR} - \mathbf{V}_{RE} \mathbf{V}_{EE}^{-1} \mathbf{V}_{ER} \right] \mathbf{s}_R. \quad (\text{B.28})$$

Eq. (B.28) is exactly the risk premium of noninvestable securities in equation (9) of De Jong and De Roon (2005). They do not solve the risk premium of noninvestable securities explicitly in terms of economic primitives, but this can be done by substituting the risk premium of eligible securities from Eq. (B.25) into Eq. (B.28):

$$\boldsymbol{\mu}_R = \gamma [\mathbf{V}_{RE} \mathbf{V}_{EE}^{-1} \mathbf{V}_{ER} \mathbf{s}_R + \mathbf{V}_{RE} \mathbf{s}_E + \mathbf{V}_{RE} \mathbf{V}_{EE}^{-1} \mathbf{V}_{ER}^* \mathbf{s}_R^*] + \gamma^d \left[\mathbf{V}_{RR} - \mathbf{V}_{RE} \mathbf{V}_{EE}^{-1} \mathbf{V}_{ER} \right] \mathbf{s}_R. \quad (\text{B.29})$$

Eq. (B.29) can be rewritten in terms of covariance forms:

$$\boldsymbol{\mu}_R = \gamma M_W \text{Cov}(\hat{\mathbf{r}}_R, r_W) + \gamma^d M_R \text{Cov}(\mathbf{r}_R, r_R - \hat{r}_R), \quad (\text{B.30})$$

where the attainable return of domestic retail securities is defined as the component of domestic retail securities returns that can be spanned by the entire set of eligible securities, including both

domestic and foreign institutional securities:

$$\hat{\mathbf{r}}_R = \mathbf{V}_{RE} \mathbf{V}_{EE}^{-1} \mathbf{r}_E$$

The attainable domestic retail portfolio is the component of the domestic retail portfolio that can be spanned by the same set of eligible securities:

$$\hat{r}_R = \frac{\mathbf{s}_R'}{M_R} \mathbf{V}_{RE} \mathbf{V}_{EE}^{-1} \mathbf{r}_E.$$

The solution Eq. (B.29) is equivalent to the risk premia of local stocks in Alexander et al. (1987) and Karolyi and Wu (2018) or the ineligible securities in Chaieb and Errunza (2007). To see the equivalence to Alexander et al. (1987) and Karolyi and Wu (2018), rewrite Eq. (B.29) by benchmarking the total risk premium with respect to the domestic market risk premium:

$$\begin{aligned} \mu_R &= \gamma^d \left[\mathbf{V}_{RR} \mathbf{s}_R + \mathbf{V}_{RE} \mathbf{s}_E \right] - (\gamma^d - \gamma) \left[\mathbf{V}_{RE} \mathbf{s}_E + \mathbf{V}_{RE} \mathbf{V}_{EE}^{-1} \mathbf{V}_{ER} \mathbf{s}_R \right] + \gamma \left[\mathbf{V}_{RE} \mathbf{V}_{EE}^{-1} \mathbf{V}_{ER^*} \mathbf{s}_{R^*} \right] \quad (\text{B.31}) \\ &= \gamma^d M_{\tilde{D}} \text{Cov}(\mathbf{r}_R, r_{\tilde{D}}) - (\gamma^d - \gamma) M_{\tilde{D}} \text{Cov}(\hat{\mathbf{r}}_R, r_{\tilde{D}}) + \gamma M_{\tilde{F}} \text{Cov}(\hat{\mathbf{r}}_R, r_{R^*}), \end{aligned}$$

where

$$M_{\tilde{D}} = M_R + M_E$$

$$M_{\tilde{F}} = M_{R^*}$$

$$r_{\tilde{D}} = \frac{M_R}{M_{\tilde{D}}} r_R + \frac{M_E}{M_{\tilde{D}}} r_E.$$

To align with the notation of Karolyi and Wu (2018) and Alexander et al. (1987) for the domestic benchmark, $M_{\tilde{D}}$ is the aggregate market value of all domestic retail securities and eligible securities, and $M_{\tilde{F}}$ is the aggregate market value of all foreign retail securities.^{IA-3}

^{IA-3}In Alexander et al. (1987), eligible securities are cross-listed domestic stocks priced globally, and the domestic benchmark is the domestic market (local + eligible). Here, because institutional securities are globally accessible in this special case, the analog of this benchmark naturally spans all eligible securities (both I and I^*) together with domestic retail.

Eq. (B.31) corresponds to equation (2c) in Karolyi and Wu (2018) or equation (20) in Alexander et al. (1987). The second term is the domestic externality (discount) for the domestic risk that can be shared with institutional and foreign retail investors through their diversification, and the third term is the foreign externality (premium) for the compensation for bearing foreign systematic risk through their correlation with eligible securities.

To see the equivalence to the ineligible securities in the partial segmentation case of Chaieb and Errunza (2007), in the absence of PPP deviation, rewrite the risk premium using the world market as a benchmark:

$$\begin{aligned}\boldsymbol{\mu}_R &= \gamma \left[\mathbf{V}_{RR} \mathbf{s}_R + \mathbf{V}_{RE} \mathbf{s}_E + \mathbf{V}_{RR^*} \mathbf{s}_{R^*} \right] + (\gamma^d - \gamma) \left[\mathbf{V}_{RR} - \mathbf{V}_{RE} \mathbf{V}_{EE}^{-1} \mathbf{V}_{ER} \right] \mathbf{s}_R \\ &\quad - \gamma \left[\mathbf{V}_{RR^*} - \mathbf{V}_{RE} \mathbf{V}_{EE}^{-1} \mathbf{V}_{ER^*} \right] \mathbf{s}_{R^*} \\ &= \gamma M_W \text{Cov}(\mathbf{r}_R, r_W) + (\gamma^d - \gamma) M_R \text{Cov}(\mathbf{r}_R, r_R | \mathbf{r}_E) - \gamma M_{R^*} \text{Cov}(\mathbf{r}_R, r_{R^*} | \mathbf{r}_E),\end{aligned}\tag{B.32}$$

where the first local risk premium is the “conditional market risk premium” and the second local premium is the “cross-market risk premium”.

From Eq. (B.25) and Eq. (B.29), the generic expression for the risk premium of any domestic security j becomes:

$$\mu_j = \gamma M_W \text{Cov}(\hat{r}_j, r_W) + \gamma^d M_R \text{Cov}(r_j, r_R - \hat{r}_R),\tag{B.33}$$

which represents a special case of the pricing result in the full model.

This special case differs from the full model in three respects. First, the institutional local risk premium disappears because retail investors now directly access institutional stocks from the other market, rather than gaining indirect exposures by trading with institutions. Second, retail securities still command a retail local risk premium, but their attainable returns are now defined as what can be spanned by both domestic and foreign institutional securities, rather than just the corresponding local institutional securities. Third, the relevant global risk-sharing benchmark

is the return on the all-asset world market portfolio, not the attainable world market portfolio, since international risk-sharing no longer depends on the span of institutional investment. Taken together, the full model nests the partially segmented market structure in Alexander et al. (1987), De Jong and De Roon (2005), and Chaieb and Errunza (2007), and generates local risk premiums for even globally accessible securities.

S3: Retail investors fully diversify In a less plausible scenario, retail investors fully diversify across all securities, and the model reduces to the market structure of Errunza and Losq (1985), Luo and Balvers (2017), and Zerbib (2022). In this case, institutional investors become the restricted investors who only invest in institutional securities, and retail investors become unrestricted investors who invest in all securities. Define the set of domestic and foreign institutional securities as eligible securities (E) and retail securities as excluded securities (X). The vector of eligible securities excess returns is $\mathbf{r}_E = [\mathbf{r}'_I, \mathbf{r}'_{I^*}]'$. The supply of eligible securities in terms of market capitalization is denoted as $\mathbf{s}_E = [\mathbf{s}'_I, \mathbf{s}'_{I^*}]'$. Denote the covariance matrix between eligible and excluded securities as:

$$\begin{bmatrix} V_{EE} & V_{EX} \\ V_{XE} & V_{XX} \end{bmatrix} = \begin{bmatrix} V_{II} & V_{II^*} & V_{IR} & V_{IR^*} \\ V_{I^*I} & V_{I^*I^*} & V_{I^*R} & V_{I^*R^*} \\ V_{RI} & V_{RI^*} & V_{RR} & V_{RR^*} \\ V_{R^*I} & V_{R^*I^*} & V_{R^*R} & V_{R^*R^*} \end{bmatrix}$$

The FOCs of domestic and foreign retail investors now become the same. Define γ^R , given by $\frac{1}{\gamma^R} = \frac{1}{\gamma^d} + \frac{1}{\gamma^f}$, as the aggregate absolute risk aversion of retail investors, and $\mathbf{x}_E^R$ and $\mathbf{x}_X^R$ as their aggregate holdings in eligible and excluded securities:

$$V_{EE}\mathbf{x}_E^R + V_{EX}\mathbf{x}_X^R = \frac{1}{\gamma^R}\boldsymbol{\mu}_E \quad (\text{B.34})$$

$$V_{XE}\mathbf{x}_E^R + V_{XX}\mathbf{x}_X^R = \frac{1}{\gamma^R}\boldsymbol{\mu}_X. \quad (\text{B.35})$$

The FOC of institutional investors remains the same as Eq. (B.24).

The equilibrium risk premium can be solved as in Errunza and Losq (1985), Luo and Balvers (2017), and Zerbib (2022). Summing Eq. (B.34) and Eq. (B.24) yields

$$\boldsymbol{\mu}_E = \gamma[\mathbf{V}_{EE}\mathbf{s}_E + \mathbf{V}_{EX}\mathbf{s}_X], \quad (\text{B.36})$$

where γ , given by $\frac{1}{\gamma} = \frac{1}{\gamma^R} + \frac{1}{\gamma^I}$, is the aggregate absolute risk aversion of all investors, and I used the market-clearing conditions:

$$\begin{aligned} \mathbf{x}_E^R + \mathbf{x}_E^I &= \mathbf{s}_E \\ \mathbf{x}_X^R &= \mathbf{s}_X. \end{aligned} \quad (\text{B.37})$$

Solving Eq. (B.34) gives the demand for eligible securities by retail investors

$$\mathbf{x}_E^R = \frac{1}{\gamma^R} \mathbf{V}_{EE}^{-1} \boldsymbol{\mu}_E - \mathbf{V}_{EE}^{-1} \mathbf{V}_{EX} \mathbf{s}_X. \quad (\text{B.38})$$

Substituting Eq. (B.36) into Eq. (B.38) yields:

$$\begin{aligned} \mathbf{x}_E^R &= \frac{\gamma}{\gamma^R} \mathbf{s}_E + \left(\frac{\gamma}{\gamma^R} - 1 \right) \mathbf{V}_{EE}^{-1} \mathbf{V}_{EX} \mathbf{s}_X \\ \mathbf{x}_E^I &= \left(1 - \frac{\gamma}{\gamma^R} \right) \left[\mathbf{s}_E + \mathbf{V}_{EE}^{-1} \mathbf{V}_{EX} \mathbf{s}_X \right]. \end{aligned} \quad (\text{B.39})$$

Substituting Eq. (B.38) into Eq. (B.35) and simplifying yields:

$$\boldsymbol{\mu}_X = \gamma[\mathbf{V}_{XE}\mathbf{s}_E + \mathbf{V}_{XX}\mathbf{s}_X] + (\gamma^R - \gamma)[\mathbf{V}_{XX}\mathbf{s}_X - \mathbf{V}_{XE}\mathbf{V}_{EE}^{-1}\mathbf{V}_{EX}\mathbf{s}_X]. \quad (\text{B.40})$$

The pricing equations Eq. (B.36) and Eq. (B.40) can be expressed in terms of their covariance

with two factors:

$$\begin{aligned}\mu_E &= \gamma M_W \text{Cov}(\mathbf{r}_E, r_W) \\ \mu_X &= \gamma M_W \text{Cov}(\mathbf{r}_X, r_W) + (\gamma^R - \gamma) M_X \text{Cov}(\mathbf{r}_X, r_X | \mathbf{r}_E),\end{aligned}$$

where r_W is the return on the world market portfolio and r_X is the return on the market portfolio of excluded (domestic and foreign retail) securities.

The generic expression for the risk premium of any security j becomes

$$\mu_j = \gamma M_W \text{Cov}(r_j, r_W) + (\gamma^R - \gamma) M_X \text{Cov}(r_j, r_X | \mathbf{r}_E). \quad (\text{B.41})$$

The second risk premium earned by excluded securities corresponds to the retail local premium in the full model, capturing the risk aversion specific to retail investors. However, there are two key differences. First, the local systematic risk in the excluded segment now includes both domestic and foreign retail securities and is influenced by the risk aversion of retail investors from both countries. Second, this local risk is conditioned on the entire span of both domestic and foreign institutional securities, rather than just the local institutional securities of each retail security. As a result, this special case is not entirely nested within the solution of the full model.

Eq. (B.41) can be further expressed in terms of beta representation with respect to the market portfolio of eligible securities as in De Jong and De Roon (2005) and Zerbib (2022), or with respect to the world market portfolio as in Luo and Balvers (2017).

Appendix B.2. The effect of global institutional investment on risk premia

How does the risk premium of a domestic retail stock k change after it is included in the choice set of global institutions? On the one hand, it would enjoy direct international risk-sharing and no longer earn the retail local risk premium. On the other hand, the attainable world market risk premium and the institutional local risk premium now depend on the covariance between the raw return of the security (r_k) with the risk factors rather than its attainable component ($\hat{r}_k$). Taking

the difference between the risk premium of an institutional security k in Eq. (A.33) and a retail security k in Eq. (A.35) yields the change in its risk premium:

$$\begin{aligned}
\Delta\mu_k &= \gamma M_W \text{Cov}(r_k - \hat{r}_k, r_{\hat{W}}) + \frac{\gamma^i}{\gamma^f} \gamma M_D \text{Cov}(r_k - \hat{r}_k, r_{\hat{D}} - r_{D^s}) - \gamma^d M_R \text{Cov}(r_k, r_R - \hat{r}_R) \\
&= \gamma M_W \text{Cov}(r_k - \hat{r}_k, \frac{M_D}{M_W} r_{\hat{D}} + \frac{M_F}{M_W} r_{\hat{F}}) + \frac{\gamma^i}{\gamma^f} \gamma M_D \text{Cov}(r_k - \hat{r}_k, r_{\hat{D}} - r_{D^s}) - \gamma^d M_R \text{Cov}(r_k, r_R - \hat{r}_R) \\
&= \gamma M_F \text{Cov}(r_k - \hat{r}_k, r_{\hat{F}}) - \frac{\gamma^i}{\gamma^f} \gamma M_D \text{Cov}(r_k - \hat{r}_k, r_{D^s}) - \gamma^d M_R \text{Cov}(r_k - \hat{r}_k, r_R - \hat{r}_R),
\end{aligned}$$

where the second equality follows from $\text{Cov}(\hat{r}_k, r_R - \hat{r}_R) = \text{Cov}(r_k - \hat{r}_k, r_{\hat{D}}) = 0$. In the plausible scenario that the unattainable component of the retail security, $r_k - \hat{r}_k$, is orthogonal to the foreign institutional securities that constitute r_{D^s} , the change in total risk premium is:

$$\gamma M_F \text{Cov}(r_k - \hat{r}_k, r_{\hat{F}}) - \gamma^d M_R \text{Cov}(r_k - \hat{r}_k, r_R - \hat{r}_R).$$

If a retail security can be perfectly replicated by its institutional counterpart, then inclusion in the institutional choice set has no effect on its risk premium. Otherwise, the effect of inclusion depends on how its unattainable exposure covaries with the retail local factor and with foreign institutional securities. Securities that covary more with local retail securities and less with foreign attainable exposures get more reduction in risk premium post-inclusion.

Appendix C. Measuring Coverage

This appendix provides additional details about the coverage measure. Table IA.3 reports, for each market, the average share of stocks covered by at least one active global fund, the share of market capitalization these stocks represent, and the median coverage among covered stocks. Table IA.4 reports the persistence of choice sets at the fund level and of the coverage-based classification at the stock level. As in Koijen and Yogo (2019), choice sets of active global open-end mutual funds are highly persistent: the median share of a fund's current positions that it also

held in the previous quarter is 92%, and 94% of current positions were held at least once over the previous 11 quarters. The stock-level classification inherits this persistence: 99% of stock-quarters retain their institutional or retail classification after one quarter, and 88% after 11 quarters. This persistence indicates that coverage captures the boundary of investment scope rather than the size of positions or recent returns.

I run two validation tests for the coverage measure. The first test examines whether coverage changes around rule-based index reconstitution events. Figure IA.4 plots coefficients on event-time dummies from regressions of coverage on quarters relative to a stock's first addition to or first deletion from the MSCI ACWI, with stock and country-quarter fixed effects and controls for log market capitalization, the FHT illiquidity measure, book-to-market, and momentum. The reference quarter is $k = -1$, non-switching stocks serve as controls, and a deletion counts only when no addition occurred in the previous 20 quarters. Coverage increases significantly after additions and decreases significantly after deletions. Around additions, coverage begins to rise before the event, indicating that coverage leads index inclusion rather than merely restating it; around deletions, coverage is flat before the event and declines only afterward. These patterns are consistent with coverage tracking changes in global institutions' investment scope associated with index reconstitution.

The second validation test compares the predictive power of coverage with that of two alternative indicators of scope, MSCI ACWI membership and the investability indicator of Karolyi and Wu (2018), for two future stock-level ownership outcomes: the number of global funds entering the stock and total global institutional ownership. I estimate the following regression:

$$y_{i,t} = \delta_1 Coverage_{i,t-1} + \delta_2 ACWI_{i,t-1} + \delta_3 Investable_{i,t-1} + \psi' X_{i,t-1} + \alpha_{c,t} + \epsilon_{i,t},$$

where $y_{i,t}$ is fund entries or global institutional ownership, $ACWI_{i,t-1}$ and $Investable_{i,t-1}$ are the index membership and investability indicators, $X_{i,t-1}$ includes the same controls as the first

validation test, and $\alpha_{c,t}$ denotes country-quarter fixed effects. Table IA.5 shows that coverage has the highest predictive power for both outcomes, with a within R^2 for fund entries of 0.53, against 0.47 for index membership and 0.38 for investability. Controlling for coverage attenuates both alternatives: for fund entries, the index membership coefficient falls from 10.7 to 6.8 and the investability coefficient from 4.8 to 1.6, and for ownership the investability coefficient becomes insignificant. The horse race confirms that coverage contains information about investment scope beyond the existing indicators.

Table IA.1. Home Bias by Investor Group as of December 2024.

This table reports the raw home bias, defined as the difference between investors' actual domestic allocation and the benchmark weight of the domestic country in the world market portfolio in Eqn. (A1), for three investor groups: aggregate residents (All), global institutional investors (Inst.), and households. A blank indicates that household data are not available for that market. Appendix A.1 provides details about the data and calculation.

Developed Markets				Emerging Markets			
Market	All	Inst.	Households	Market	All	Inst.	Households
Australia	0.532	0.270		Brazil	0.856	0.496	
Austria	0.210	0.044	0.783	Chile	0.448	0.241	
Belgium	0.202	0.017	0.730	China	0.828	0.426	
Canada	0.573	0.204		Greece	0.758	0.434	0.921
Switzerland	0.471	0.223		Hungary	0.539	0.043	0.939
Germany	0.250	0.118	0.629	Korea	0.590	0.423	
Denmark	0.307	0.028	0.938	Mexico	0.833	0.607	
Spain	0.423	0.093	0.866	Malaysia	0.745	0.245	
Finland	0.188	0.153	0.915	Poland	0.833	0.501	0.976
France	0.463	0.110	0.909	Saudi Arabia	0.845	-0.022	
UK	0.099	0.078	0.886	Thailand	0.854	0.536	
Hong Kong	0.275	0.020		South Africa	0.522	0.426	
Israel	0.651	0.537		<i># Markets</i>	12	12	3
Italy	0.264	0.106	0.734	<i>Average</i>	0.721	0.363	0.945
Japan	0.617	0.331					
Netherlands	0.199	0.015	0.877				
Norway	0.114	0.010	0.936				
New Zealand	0.382	0.201					
Portugal	0.255	0.054	0.922				
Singapore	0.176	0.024					
Sweden	0.350	0.210	0.919				
US	0.311	0.157					
<i># Markets</i>	22	22	13				
<i>Average</i>	0.332	0.136	0.850				

Table IA.2. FactSet Funds Matched to Morningstar.

This table reports, at the fourth quarter of each year, the number of FactSet open-end funds with reported AUM and their total AUM in trillions of USD, and the proportions of these funds, by count and by AUM, that are matched to a Morningstar fund, to a Morningstar equity fund, and, within equity funds, that are declared-global active open-end funds. A fund is declared global when its Morningstar global category is “Global Equity” or “Global Emerging Markets Equity” and its investment area is unrestricted or global; active funds are open-end funds that are not classified as index funds by Morningstar. Across matched funds, the median correlation between FactSet and Morningstar monthly returns is 0.98 and between quarterly AUM is 0.97; 90% and 70% of matched funds have AUM and return correlations exceeding 0.9. The Average row is the simple average over the yearly rows.

Year	Funds	AUM (USD trn)	Matched		Equity		Global active	
			Count	AUM	Count	AUM	Count	AUM
2002	17,193	3.1	0.81	0.91	0.52	0.82	0.09	0.13
2003	18,975	4.5	0.80	0.90	0.51	0.80	0.09	0.13
2004	19,640	5.6	0.79	0.90	0.52	0.80	0.09	0.14
2005	19,965	6.5	0.80	0.90	0.53	0.80	0.09	0.16
2006	20,867	8.2	0.80	0.90	0.54	0.80	0.09	0.18
2007	22,244	9.2	0.80	0.90	0.53	0.80	0.09	0.19
2008	23,610	5.1	0.79	0.89	0.52	0.77	0.10	0.18
2009	24,155	7.2	0.78	0.88	0.51	0.76	0.09	0.19
2010	25,083	8.3	0.77	0.88	0.50	0.77	0.09	0.19
2011	25,488	7.4	0.77	0.89	0.49	0.77	0.10	0.19
2012	25,206	8.4	0.77	0.89	0.49	0.77	0.10	0.20
2013	24,504	10.6	0.78	0.89	0.50	0.78	0.10	0.19
2014	25,616	11.2	0.78	0.89	0.49	0.77	0.11	0.19
2015	26,982	10.9	0.77	0.89	0.48	0.78	0.11	0.19
2016	27,677	11.2	0.75	0.89	0.46	0.77	0.10	0.19
2017	28,105	13.6	0.74	0.89	0.46	0.78	0.11	0.20
2018	28,924	12.1	0.74	0.89	0.45	0.78	0.11	0.20
2019	29,621	14.8	0.72	0.89	0.45	0.78	0.11	0.20
2020	31,348	17.5	0.71	0.88	0.43	0.78	0.11	0.20
2021	33,553	20.8	0.69	0.87	0.42	0.77	0.10	0.19
2022	32,465	15.9	0.67	0.86	0.43	0.76	0.11	0.19
2023	33,426	18.5	0.62	0.87	0.41	0.76	0.10	0.19
2024	33,612	20.5	0.59	0.86	0.39	0.76	0.10	0.18
<i>Average</i>	26,011	10.9	0.75	0.89	0.48	0.78	0.10	0.18

Table IA.3. Coverage by Market.

This table reports, for each market, the number of distinct linkable stocks over the sample (# Stocks), and the average over time of the share of stocks covered by at least one active global fund (Coverage > 0; Covered), the share of market capitalization they represent (Mkt cap), and the median coverage among covered stocks (Median). Coverage is the AUM of active global funds whose choice set contains the stock, as a share of the total AUM of active global funds participating in the stock's market. The sample spans 2002Q4 to 2024Q4. The Average rows report the simple average across markets.

Developed Markets					Emerging Markets				
Market	# Stocks	Covered	Mkt cap	Median	Market	# Stocks	Covered	Mkt cap	Median
Australia	2,662	0.361	0.954	0.018	Brazil	339	0.707	0.934	0.038
Austria	123	0.662	0.955	0.044	Chile	196	0.461	0.814	0.070
Belgium	249	0.593	0.891	0.022	China	5,200	0.139	0.418	0.007
Canada	4,728	0.294	0.959	0.015	Egypt	197	0.322	0.715	0.020
Denmark	272	0.595	0.926	0.024	Greece	323	0.401	0.882	0.039
Finland	256	0.683	0.980	0.024	India	3,293	0.357	0.889	0.018
France	1,293	0.500	0.954	0.017	Indonesia	821	0.393	0.813	0.022
Germany	1,130	0.549	0.961	0.015	Korea	2,772	0.511	0.917	0.013
Hong Kong	2,861	0.589	0.961	0.013	Kuwait	178	0.279	0.834	0.221
Israel	773	0.372	0.881	0.029	Malaysia	1,220	0.370	0.914	0.031
Italy	677	0.653	0.971	0.022	Mexico	187	0.746	0.964	0.030
Japan	4,897	0.649	0.962	0.018	Philippines	291	0.410	0.855	0.035
Netherlands	231	0.721	0.933	0.034	Poland	777	0.265	0.880	0.035
New Zealand	184	0.572	0.891	0.050	Qatar	55	0.683	0.903	0.087
Norway	574	0.627	0.929	0.021	Saudi Arabia	290	0.367	0.755	0.138
Singapore	911	0.481	0.951	0.016	South Africa	498	0.544	0.909	0.029
Spain	349	0.684	0.949	0.026	Taiwan	2,040	0.524	0.921	0.016
Sweden	996	0.547	0.908	0.021	Thailand	892	0.391	0.872	0.017
Switzerland	320	0.795	0.974	0.024	Turkiye	549	0.459	0.875	0.027
United Kingdom	3,282	0.561	0.978	0.014	United Arab Emirates	145	0.466	0.733	0.031
United States	15,000	0.723	0.980	0.007					
<i>Average</i>	1,988	0.582	0.945	0.023	<i>Average</i>	1,013	0.440	0.840	0.046

Table IA.4. Persistence of Choice Sets and of the Classification.

Panel A reports the share of an active global fund's current positions that the fund held at least once in the previous L quarters, for $L = 1$ to 11; the share is computed for each fund-quarter and the panel reports its median and mean across fund-quarters, following the convention of Kojien and Yogo (2019), over fund-quarters from 2003Q1, the first quarter with a full twelve-quarter history, to 2024Q4. Panel B reports the share of stock-quarters whose classification as institutional (coverage above zero) or retail is the same as L quarters earlier, over the stock panel of the 41 sample markets from 2002Q4 to 2024Q4.

L	1	2	3	4	5	6	7	8	9	10	11
Panel A: Choice-Set Persistence											
Median	0.924	0.927	0.930	0.932	0.933	0.935	0.936	0.938	0.939	0.939	0.940
Mean	0.858	0.863	0.869	0.871	0.875	0.876	0.879	0.880	0.882	0.883	0.884
Panel B: Classification Persistence											
Share unchanged	0.986	0.974	0.962	0.951	0.940	0.930	0.920	0.910	0.901	0.892	0.883

Table IA.5. Coverage versus Index Inclusion and Investability.

This table reports stock-level panel regressions of institutional holdings on lagged coverage, an MSCI ACWI membership indicator, and an investability indicator: coverage alone, each indicator alone and with coverage, and all three jointly. A stock is investable when its firm has a cross-listing in an open market outside its home market, following Karolyi and Wu (2018). Fund entries count the matched global funds that hold the stock in the quarter and did not in the previous quarter; global institutional ownership is aggregated across all global institutions in FactSet. All regressions include country-quarter fixed effects and control for log market capitalization, the FHT illiquidity measure, book-to-market, and momentum. Standard errors clustered by country-quarter in parentheses; ***, **, * denote significance at the 1%, 5%, 10% levels.

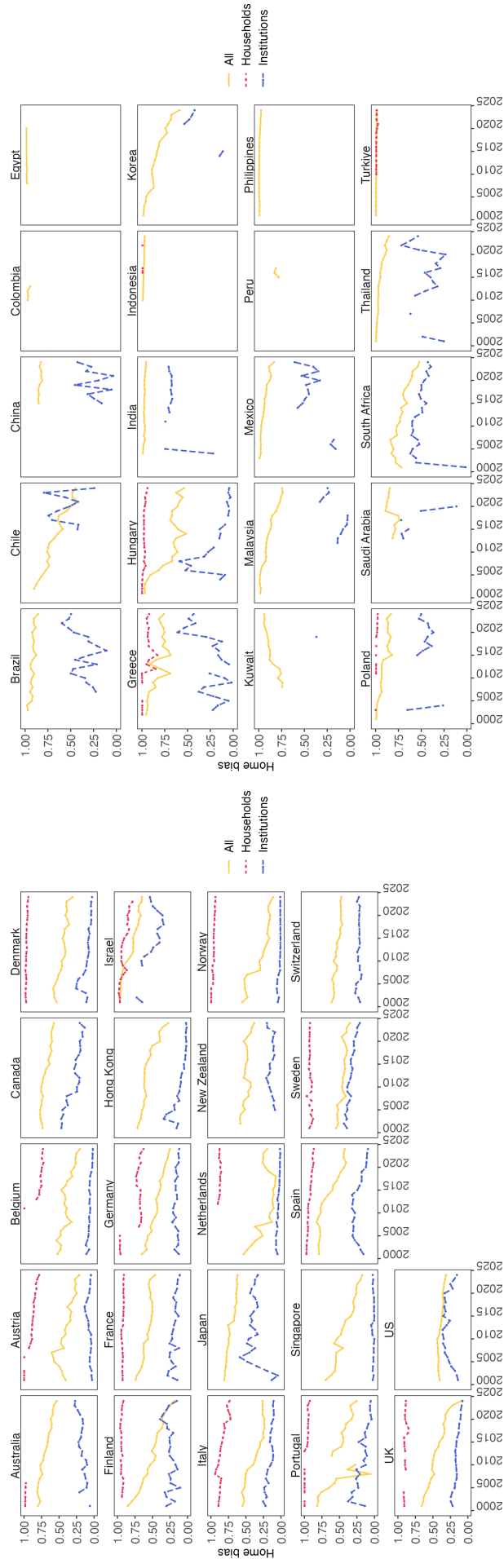
	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: Fund Entries						
Coverage	65.940*** (0.855)		53.228*** (0.824)		63.432*** (0.948)	51.175*** (0.857)
ACWI member		10.722*** (0.202)	6.805*** (0.194)			6.748*** (0.197)
Investable (KW)				4.827*** (0.141)	1.586*** (0.141)	1.366*** (0.111)
N	2,094,515	2,094,515	2,094,515	2,094,515	2,094,515	2,094,515
Within R^2	0.526	0.470	0.565	0.384	0.529	0.568
Panel B: Global Institutional Ownership						
Coverage	0.195*** (0.009)		0.179*** (0.009)		0.192*** (0.006)	0.176*** (0.007)
ACWI member		0.022*** (0.001)	0.009*** (0.001)			0.009*** (0.001)
Investable (KW)				0.012*** (0.002)	0.002 (0.002)	0.002 (0.002)
N	2,083,855	2,083,855	2,083,855	2,083,855	2,083,855	2,083,855
Within R^2	0.289	0.279	0.289	0.276	0.289	0.289

Table IA.6. Stock-Level Global Institutional Ownership and Factor Exposures.

This table reports, for each market, the number of stocks (N) for which stock-level global institutional ownership percentile rank is selected in the conditional factor loading specification and its median coefficient (*Median*) in the stock-specific instrument coefficient matrix C_i in Eq. (24). The coefficients are estimated from the conditional estimation in Section 3.2, for 21 developed markets (Panel A) and 19 emerging markets (Panel B). The last row of each panel counts the markets whose median coefficient on the retail local factor is negative, the one sign the model implies.

Market	World		Inst. Local		Retail Local	
	N	Median	N	Median	N	Median
Panel A: Developed Markets						
Australia	682	0.048	708	0.084	715	-0.393
Austria	25	0.717	27	0.045	27	-0.329
Belgium	59	0.335	58	-0.451	63	-0.208
Canada	311	-0.113	336	-0.372	316	-0.167
Denmark	111	0.082	109	0.104	111	-0.154
Finland	80	0.071	88	0.179	85	0.103
France	316	0.069	324	-0.050	317	-0.185
Germany	337	0.349	352	0.261	351	0.293
Hong Kong	551	-0.093	545	-0.132	568	-0.608
Israel	313	0.006	316	0.036	316	-0.460
Italy	105	-0.170	132	0.094	116	-0.272
Japan	1,545	0.143	1,547	0.075	1,579	-0.045
Netherlands	35	0.349	37	-0.385	38	-0.027
New Zealand	55	-0.060	58	0.202	62	-0.257
Norway	54	0.548	63	0.219	72	-0.251
Singapore	237	-0.145	244	-0.105	246	-0.324
Spain	69	0.209	76	0.552	71	-0.413
Sweden	185	0.118	181	-0.063	180	0.103
Switzerland	112	0.233	114	-0.050	113	-0.810
UK	337	-0.238	380	0.518	378	0.278
US	2,258	0.042	2,305	0.213	2,304	-0.050
<i>No. markets, median < 0</i>						17
Panel B: Emerging Markets						
Chile	54	-0.366	54	0.088	57	-0.402
China	2,121	0.011	2,170	0.018	2,173	-0.658
Egypt	99	-0.796	98	-0.298	96	-0.307
Greece	182	0.242	181	0.012	179	0.418
India	771	-0.299	751	-0.061	753	0.210
Indonesia	209	-0.013	210	-0.128	206	-0.568
Korea	775	0.201	817	-0.134	820	-0.182
Kuwait	76	-0.289	76	-0.198	82	-1.143
Malaysia	364	0.146	365	-0.086	377	-0.343
Mexico	43	0.483	43	0.061	44	0.082
Philippines	77	0.333	75	0.385	74	0.068
Poland	203	-0.187	205	0.214	208	0.437
Qatar	37	0.250	32	0.480	35	0.103
Saudi Arabia	128	-0.168	119	-0.020	127	-0.193
South Africa	154	-0.047	154	0.034	153	0.025
<i>Continue on the next page</i>						

Market	World		Inst. Local		Retail Local	
	N	Median	N	Median	N	Median
Taiwan	805	-0.128	810	-0.421	842	-0.443
Thailand	272	0.058	268	-0.099	272	-0.465
Turkiye	147	0.041	143	-0.069	152	-0.096
UAE	46	-0.572	49	-0.528	49	-0.965
<i>No. markets, median < 0</i>						12



A. Developed markets

B. Emerging markets

Figure IA.2. Home Bias by Investor Group, 2001–2024.

This figure plots the time series of the raw home bias for three investor groups — all residents, global institutions, and households — in developed markets (Panel A) and emerging markets (Panel B) from 2001 to 2024. The figure includes all economies with available home-bias data and is broader than the sample used in the main asset pricing tests. Home bias is defined as the difference between each group’s actual domestic portfolio weight and the country’s weight in the world market portfolio, using portfolio holdings from the IMF Coordinated Portfolio Investment Survey (CPIS), OECD National Accounts, and FactSet. Appendix A provides details on variable construction and data sources.

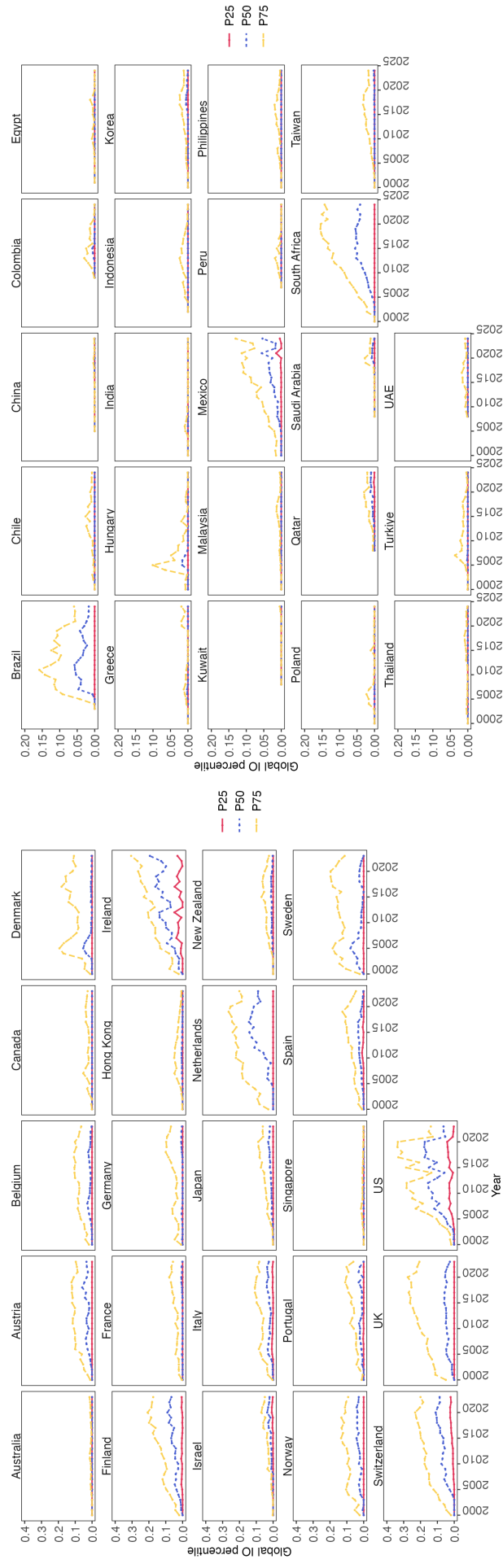


Figure IA.3. Global Institutional Ownership Quartiles, 2001 – 2024.

This figure plots the quartiles of stock-level global institutional ownership for each country in developed markets (Panel A) and emerging markets (Panel B) from 2001 to 2024. The figure includes all economies with available FactSet ownership coverage and is broader than the sample used in the main asset pricing tests. Stock-level ownership is calculated using FactSet Ownership following the method of Ferreira and Matos (2008). Appendix A provides details on variable construction and data sources.

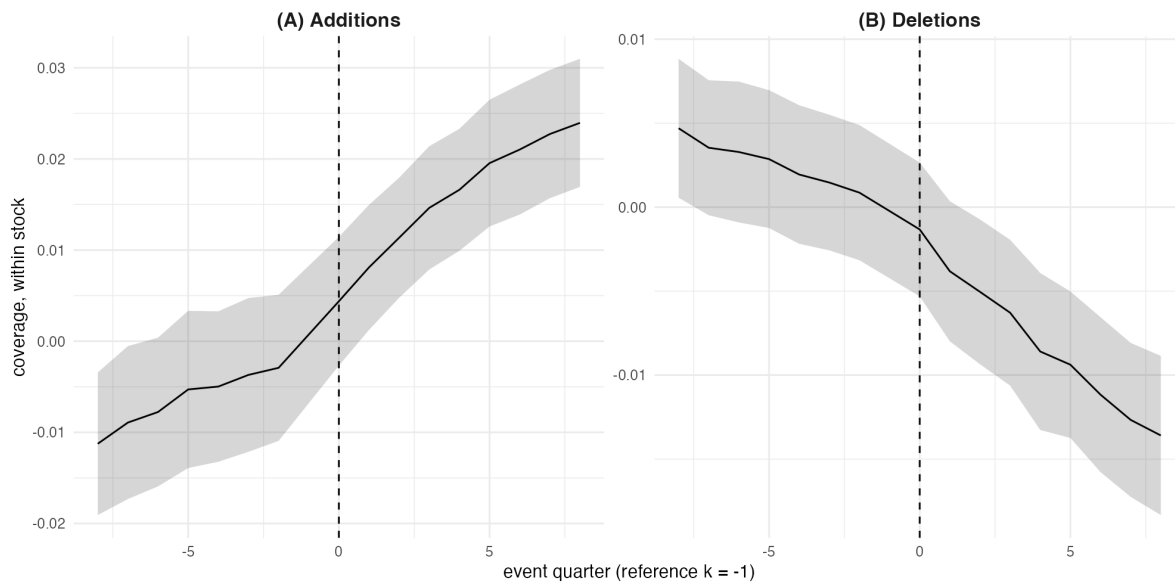


Figure IA.4. Coverage around MSCI ACWI Additions and Deletions.

This figure plots coefficients on event-time dummies from regressions of coverage on quarters relative to a stock's first addition to (left) or first deletion from (right) the MSCI ACWI, with stock and country-quarter fixed effects and controls for log market capitalization, the FHT illiquidity measure, book-to-market, and momentum; the reference quarter is $k = -1$ and non-switching stocks serve as controls. A deletion counts only when no addition occurred in the previous 20 quarters. Shaded areas are 95% confidence bands from standard errors clustered by country-quarter. Coverage is the AUM of active global funds whose choice set contains the stock, as a share of the total AUM of active global funds participating in the stock's market.