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Working Paper 17-085

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First draft: July 2016

This draft: March 2017

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Abstract

This paper conducts a theoretical and empirical investigation of the risks of globally diversified portfolios of stocks and bonds and of optimal intertemporal global portfolio choice for long horizon investors in the presence of permanent cash flow shocks and transitory discount rate shocks to asset values. We show that an upward shift in cross-country one-period return correlations resulting from correlated cash flow shocks increases the risk of global portfolios and reduces investors' willingness to hold risky assets at all horizons. However, a similar upward shift in cross-country one-period return correlations resulting from correlated discount rate shocks has a much more muted effect on long-run portfolio risk and on the willingness to long horizon investors to hold risky assets. Correlated cash flow shocks imply that markets tend to move together at all horizons, thus reducing the scope for global diversification for all investors regardless of their investment horizon. By contrast, correlated discount rate shocks imply that markets tend to move together only transitorily and long-horizon investors can still benefit from global portfolios to diversify long-term cash flow risk. We document a secular increase in the cross-country correlations of stock and government bond returns since the late 1990ís. We show that for global equities this increase has been driven primarily by increased cross-country correlations of discount rate shocks, or global capital markets integration, while for bonds it has been driven by both global capital markets integration and increased cross-country correlations of inflation shocks that determine the real cash flows of nominal government bonds. Therefore, despite the significant increase in the short-run correlation of global equity markets, the benefits from global equity portfolio diversification have not declined nearly as much for long-horizon investors as they have for short-horizon investors. By contrast, increased correlation of inflation across markets implies that the benefits of global bond portfolio diversification have declined for long-only bond investors at all horizons. However, it also means that the scope for hedging liabilities using global bonds has increased, benefiting investors with long-dated liabilities. Finally, we show that the well documented negative stock-bond correlation in the U.S. since the late 1990's is a global phenomenon, suggesting that the benefits of stock-bond diversification have increased in all developed markets.

JEL classification: G12.

1 Introduction

The principle of portfolio diversification states that investors can significantly reduce their exposure to uncompensated risk by holding a well diversified portfolio of asset classes and securities (Markowitz, 1952). The enormous growth of assets under management in investment vehicles that emphasize portfolio diversification and index investing is arguably a direct result of the widespread adoption of this insight from academic Finance in investment practice (Sunderam et al. 2015, Viceira and Ciechanover 2016). However, although investors appear to have embraced the principle of diversification for their portfolios of domestic equities and bonds, they still seem reluctant to hold globally diversified portfolios despite the broad availability of inexpensive investment vehicles to invest globally and the well documented historical benefits of international portfolio diversification at short horizons (French and Poterba 1991, Philips, 2014, Bekaert et al., 2016).[2]

This paper conducts a theoretical and empirical investigation of the benefits of global portfolio diversification for long horizon investors. We argue that long-horizon considerations make the case for holding globally diversified portfolios even stronger in the sense that increases in short-run cross-country return correlations that imply a reduction of the benefits of global diversification for short horizon investors do not necessarily imply such a reduction for long horizon investors. Long-horizon investors can still benefit from diversifying their portfolios across equity and bond markets even when short-horizon investors see those benefits diminished.

The optimality of holding globally diversified portfolios has been examined under the assumption that investment opportunities are constant (Grubel 1968, Solnik 1974, French and Poterba 1991, De Santis and Gerard 1997). Under this assumption, investment horizon is irrelevant for portfolio decisions and, aside from investors risk tolerance, the main determinant for portfolio choice is the volatility, correlation, and expected return structure of asset returns (Samuelson 1968, Merton 1969). In such environment, an increase in cross-country return correlations, all else equal, reduces the benefits of international portfolio diversification for all long-only investors regardless of their investment horizon and risk tolerance.

Building on this logic, the traditional empirical argument for holding globally diver-

[2]The "home bias" in investors portfolios was first documented by French and Poterba (1991), who argued that it was hard to justify in light of the volatility and cross-country correlation of returns in global equity markets, plausible expected return assumptions, and the transaction costs of investing abroad. Vanguard (2014) reports that while U.S. equities accounted for 51% of global equity markets capitalization on December 31, 2013, U.S. mutual investors held, on average, only 27% of their total equity allocation in non-U.S. equity funds. Typical portfolio advice and actual portfolio construction by investment professionals also appear to be biased towards domestic assets. For exampe, in the same report, Vanguard suggests "a reasonable starting allocation to non-U.S. stocks of 20%, within an upper limit based on global market capitalization." Life-cycle funds, which have become the default allocation in defined-contribution pension plans, exhibit a domestic bias built into their equity and bond allocations. Bekaert et al. (2016) examine the international equity allocations of 3.8 million individuals over the period 2005-2011, and find that younger cohorts tend to be more internationally diversified than older cohorts, and that all cohorts seem to have increased their exposure to global stocks over time.

sified equity portfolios has relied primarily on the fact that global stocks has historically exhibited low enough cross-country return correlations that investors would need to have implausibly large return expectations on their own domestic markets relative to other markets to justify holding a domestically biased portfolio (French and Poterba 1991). Campbell, Serfaty-de-Medeiros, and Viceira (2010) document that historically currency-hedged long-term government bond returns exhibit cross-country correlations even lower than those of equities, suggesting that the case for holding globally diversified bond portfolios is also strong.

However, in recent decades as the cross-country correlations of global markets have experienced a significant increase (Quinn and Voth, 2008, Asness, Israelov, and Liew, 2011), suggesting a diminution of the benefits of global portfolio diversification. Figure 1 illustrates this empirical phenomenon. It plots the average 3-year moving correlation of monthly currency-hedged equity and bond returns across seven major markets (Australia, Canada, France, Germany, Japan, United Kingdom, and the United States) that make up most of global market capitalization for the 1986-2013 period. Figure 1 shows a secular increase in the cross-country correlations of stock and bond returns, likely driven by the phenomena of globalization of trade and capital flows.[3] The figure also shows that the global financial crisis of late 2008 and early 2009 also led to a further temporary significant increase in correlations.

Complementing Figure 1, Figure 2 plots the average 3-year moving stock-bond correlation both within countries and across countries. This figure shows a strong decline in the stock-bond correlation over the same period, including a reversal of its sign from positive to negative since the turn of the century. Figure 2 shows that this phenomenon, which has been well documented for the U.S. and the U.K., extends to a wide cross-section of developed economies (Campbell, Shiller, and Viceira 2009, Campbell, Sunderam, and Viceira, 2007). It suggests that, as the benefits of international portfolio diversification within stock and bond portfolios appear to have declined over time, the benefits of diversification across stocks and equities appear to have increased.

The traditional argument for global portfolio diversification assumes that discount rates are constant and that all variation in asset values and returns is driven by news about cash flows. However, research in Finance in recent decades has documented ample empirical evidence of predictable transitory variation in discount rates, both real interest rates and asset returns at the asset class level and at the individual security level (Campbell 1991, Cochrane, 2008 and 2011, Vuolteeenaho 2002). This evidence implies that realized asset returns and asset valuation vary over time as the result of both shocks to cash flows, which empirically appear to be permanent, and shocks to discount rates, which appear to be transitory (Campbell and Shiller, 1988, Campbell 1991, Campbell and Vuolteenaho 2004). Time variation in discount rates also implies a wedge between the optimal portfolios of long horizon investors and those of short horizon investors (see Campbell and Viceira 2002 for a textbook treat-

[3]Interestingly Figure 2, which plots the average 3-year moving stock-bond correlation both within countries and across countries, shows a strong decline in the stock-bond correlation from positive to negative. This suggests that, as the benefits of international portfolio diversification within stock and bond portfolios appear to have declined over time, the benefits of diversification between stocks and equities appear to have increased.

ment).

In light of this evidence and its implications for optimal portfolio choice, we revisit the case for global portfolio diversification in an environment with transitory variation in discount rates and permanent cash flow shocks and transitory discount rate shocks to asset valuations. In such environment, both types of shocks can drive the correlation of returns across assets or markets. We show that an increase in cross-country return correlations like the one documented in Figure 1 does not necessarily imply a decline of the benefits of global portfolio diversification for long-horizon investors all else equal. Whether such increase reduces the benefits of global portfolio diversification for long-horizon investors depends crucially on what type of cross-country news correlation drives it.

If it is driven by an increase in the cross-country correlations of cash flow news, the benefits of portfolio diversification decline for all investors regardless of their horizon. Intuitively, such an increase implies that permanent cash flow shocks to market valuations tend to happen simultaneously, thus reducing the scope for global diversification for all investors regardless of their investment horizon. If it is driven by an increase in the cross-country correlations of discount rate news, the benefits of portfolio diversification for all investors decline unambiguously for short horizon investors, but not for long horizon investors. Intuitively, such an increase implies that transitory discount rate shocks to valuations tend to happen simultaneously, driving up short-run cross-country correlations but not long-run correlations, which are driven by permanent cash flow shocks. Therefore, the scope for global diversification for long-term investors is not diminished.

To illustrate this result, we build a stylized symmetrical model of global capital markets and show that an upward shift in the cross-country correlations of cash flow news increases portfolio risk equally at all horizons, while an upward shift in the cross-country correlations of discount rate news increases portfolio risk relatively less or not at all at long horizons. We also examine in the context of this stylized model of identical markets calibrated to U.S stock returns the impact of increases in the cross-country correlation of news on optimal portfolio choice at long and short horizons, assuming investors maximize expected power utility of wealth at a finite horizon (Jurek and Viceira, 2011). We consider increases in the cross-country correlations of both cash flow news and discount rate news that result in an identical increase in the cross-country correlation of one-period returns. We show that an increase in cash flow news cross-country correlations leads to a reduction in the optimal equity holdings which is much larger at long-horizons that at short horizons. By contrast, an increase in discount rate news cross-country correlations leads to a much smaller reduction in optimal equity holdings at all horizons. Therefore, our results imply that similar increases in short-run cross-country return correlations have a much larger impact on optimal portfolios at long horizons when they are driven by increased correlation of cash flow news.

We explore the implications of these insights for global diversification in stocks and bonds. We start our empirical analysis by estimating the sources of cross-country return correlations in equity and sovereign bond markets in the 1986-2013 period for a cross section of seven developed economies representing most of global market capitalization. We also estimate the changes in these correlations between the first and the second half of the sample. We do not account in our analysis for the estimation uncertainty associated with our estimates of predictable variation in discount rates and the subsequent decomposition of realized returns into their news components, although we use simultaneously the whole cross-section of countries in our estimation.[4]

Our news estimates are based on the the return decomposition and news estimation framework of Campbell (1991). Following Ammer and Mei (1996), we interpret an increase in the cross-country correlations of discount rate news as an indicator of increased financial integration of markets, and an increase in the cross-country correlations of cash flow news as an indicator of increased real integration for stocks. For bonds, an increase in the cross-country correlations of cash flow news reflects an increase in the cross-country correlation of inflation news. This follows directly from the fact that the bonds we use in our analysis are nominal bonds, so their real cash flows vary inversely with inflation.

We document an economically and statistically significant increase in the average cross-country correlation of discount rate news in the 2000-2013 period relative to the 1986-1999 period for both stocks and bonds, and a significant increase in the average cross-country correlation of cash flow news (i.e., inflation news) for bonds. Our estimates suggest that the degree of real and financial integration of global stock and bond markets has increased in the most recent period, with capital market integration being the main driver of the increased co-movement of global equity and bond markets. Arguably the freedom of capital to flow across borders has drastically reduced capital market segmentation: Today the marginal investor in most developed markets is more likely to be a global investor, and investor sentiment and risk aversion in developed markets tend to move together more strongly than in the past.

Our results about increased real and financial global market integration are robust to alternative measures of market integration. In particular, we expand the R^2-based measure of market integration of Pukthuanthong and Roll (2009) to accommodate the cash flow and discount rate decomposition of realized returns. We also find strong evidence of increased real and financial market integration in the second subperiod, especially financial market integration, under this alternative measure.

Next we explore the implications of our empirical findings for global portfolio diversification in two different but related ways. First, following the methodology in Campbell and Viceira (2005), we compute the risk of global portfolios of stocks and bonds across investment horizons and across subsamples. For equities, we find that the long-run risk of internationally diversified stock portfolios has in fact declined in the late period relative to the early period, despite the significant increase in short-run cross-country return correlations in the late subperiod. We show that this decline in long-run portfolio equity risk is the result of both a cross-country covariance (or correlation) effect and a within-country variance effect.

[4] There is disagreement in the literature about how precisely one can estimate time variation in expected returns: See Campbell and Yogo (2006), Campbell and Thompson (2008), Goyal and Welch (2008), and Pastor and Stambaugh (2009 and 2012).

The cross-country correlation effect is that capital market integration, i.e., increased cross-country correlations of discount rate news, is the main driver of the increase in short-run cross-country return covariances in the late period. As our model shows, this effect increases short-run cross-country return covariances but not their long-run counterparts. The within-country variance effect is increased within-country stock return predictability over this period, which results in a decline of long-run within-country return variances.

By contrast, our estimates indicate that the risk of global bond portfolios has shifted upwards in the second subperiod across all horizons, consistent with our prior finding that, for bonds, the cross-country correlations of both discount rate news and cash flow news have increased in the second subperiod. This implies that the benefits of global diversification in bonds have declined for long-only investors in the most recent period, regardless of their investment horizon. Interestingly, it also implies that global bond portfolio diversification is beneficial to investors with long-term liabilities such as pension funds. Such investors can use global bonds to hedge their local pension liabilities. These benefits can be especially large to investors whose liabilities are large relative to the size of their domestic bond markets and are exposed to adverse price pressure when they try to hedge their liabilities in their local markets (Greenwood and Vayanos 2008, Hamilton and Wu 2012).

Second, we compute optimal intertemporal global equity portfolio allocations and expected utility implied by our estimates across periods under different assumptions about investor preferences. We consider an investor with power utility defined over terminal wealth at a finite horizon as in Jurek and Viceira (2011) and another with Epstein-Zin utility over instantaneous consumption and an infinite horizon as in Campbell and Viceira (1999) and Campbell, Chan, and Viceira (2003). Our findings suggest that the increase in the cross-country correlations of stock returns has not led to reduction in the benefits of global equity portfolio diversification at long horizons in the most recent period, even after we control for the increase in within-country stock return predictability. Because this increase results from correlated discount rate news, long-horizon investors still find that holding global equity portfolios helps diversify cash flow risk.

The paper is organized as follows. Section 2 introduces the basic asset return decomposition into cash flow news and discount rate news. Section 3 explores long-run portfolio risk and optimal intertemporal global portfolio diversification in a stylized symmetrical model of global markets. This section provides insights into the differential effects of each type of returns news on long-run global portfolio risk and portfolio choice. Section 4 conducts an empirical analysis of the changes in cross-country stock and bond return correlations over time and the sources of these changes. Section 5 explores the implications of those changes for the risk of globally diversified portfolios of stocks and bonds across investment horizons, and for optimal intertemporal portfolio choice. Finally, Section 6 concludes.

2 Asset Return Decomposition

The starting point of our analysis is the log-linear approximation to present value relations of Campbell and Shiller (1988) and the return decomposition of Campbell (1991). A log-linearization of the return on an asset around the unconditional mean of its dividend-price ratio—where dividend is a proxy for cash flow—implies

$$r_{t+1} \approx k + \rho p_{t+1} + (1-\rho)\, d_{t+1} - p_t, \tag{1}$$

and

$$p_t - d_t = \frac{k}{1-\rho} + \mathbb{E}_{t+1} \sum_{j=0}^{\infty} \rho^j \left[\Delta d_{t+1+j} - r_{t+1+j}\right], \tag{2}$$

where r_t denotes the natural log of the gross total return on the asset, p_t the log of its price P_t, and Δd_{t+1} the change in the log dividend (or cash flow). The constants ρ and k are log-linearization parameters, with $\rho \equiv 1/\left(1+\exp\left(\overline{d-p}\right)\right)$ and $k \equiv -\log(\rho) - (1-\rho)\log(1/\rho - 1)$, where $\overline{d-p}$ denotes the unconditional mean of the log dividend-price ratio. This log-linear approximation rules out bubbles by imposing $\lim_{j\to\infty} \rho^j p_{t+j} = 0$.

Substitution of (2) into (1) gives the following decomposition of realized returns (Campbell 1991):

$$r_{t+1} - \mathbb{E}_t\left[r_{t+1}\right] = \left(\mathbb{E}_{t+1} - \mathbb{E}_t\right) \sum_{j=0}^{\infty} \rho_i^j \Delta d_{t+1+j} - \left(\mathbb{E}_{t+1} - \mathbb{E}_t\right) \sum_{j=1}^{\infty} \rho^j r_{t+1+j}. \tag{3}$$

Equation (3) shows that the unexpected log return on an asset reflects changes in either its expected future cash flows or in its expected future returns (or discount rates). Following standard terminology in this literature, we will refer to the former as cash flow shocks or cash flow news, and to the latter as discount rate shocks or discount rate news, and write more succinctly

$$r_{t+1} - \mathbb{E}_t\left[r_{t+1}\right] \equiv N_{CF,t+1} - N_{DR,t+1}. \tag{4}$$

We can further decompose $N_{DR,t+1}$ into news about excess log returns—or risk premia—, and news about the return on the reference asset used to compute excess returns:

$$N_{DR,t+1} = N_{RR,t+1} + N_{RP,t+1}, \tag{5}$$

with

$$N_{RR,s,t+1} \equiv \left(\mathbb{E}_{t+1} - \mathbb{E}_t\right)\left[\sum_{j=1}^{\infty} \rho^j r^N_{f,t+1+j}\right],$$

$$N_{RP,s,t+1} \equiv \left(\mathbb{E}_{t+1} - \mathbb{E}_t\right)\left[\sum_{j=1}^{\infty} \rho^j x r_{t+1+j}\right],$$

where $xr_{t+1+j} = r_{t+1+j} + r^N_{t+1+j}$ and r^N_{t+1+j} is the log return on the benchmark asset.

In our empirical analysis we use cash (i.e., a short-term nominal bond like a T-bill in the US) as the reference asset, and measure both excess log returns and the return on the short-term bond in real terms. That is $r^N_{f,t+1} = y^N_{1,t} - \pi_{t+1}$, where $y^N_{1,t}$ denotes the yield on a one-period nominal bond at t, which is also its nominal return at $t+1$, and π_{t+1} denotes log inflation.

The preceding expressions assume the asset is a perpetual claim on cash flows, such as equities or a consol bond. In our empirical analysis we also consider nominal bonds, whose cash flows (i.e., coupons) are fixed in nominal terms—and thus in real terms they vary inversely with the price level—and have a fixed maturity. The Appendix shows that for a \$1-coupon nominal bond with maturity n,

$$r_{n,t+1} - \mathbb{E}_t \left[r_{n,t+1} \right] = N_{CF,n,t+1} - N_{RR,n,t+1} - N_{RP,n,t+1}, \tag{6}$$

with

$$\begin{aligned} N_{CF,n,t+1} = -N_{INFL,n,t+1} &\equiv - \left(\mathbb{E}_{t+1} - \mathbb{E}_t \right) \left[\sum_{j=1}^{n-1} \rho_b^j \pi_{t+1+j} \right], \\ N_{RR,n,t+1} &\equiv \left(\mathbb{E}_{t+1} - \mathbb{E}_t \right) \left[\sum_{j=1}^{n-1} \rho_b^j r^N_{f,t+1+j} \right], \\ N_{RP,n,t+1} &\equiv \left(\mathbb{E}_{t+1} - \mathbb{E}_t \right) \left[\sum_{j=1}^{n-1} \rho_b^j xr_{n-j,t+1+j} \right], \end{aligned}$$

and $\rho_b = 1 / \left(1 + \exp \left(-\bar{p_n} \right) \right)$.

The news components defined above are not directly observable, but we can infer them from a return generating model. We follow Campbell (1991) and assume that the asset return generating process follows a first-order vector autoregressive (VAR) model:

$$\tilde{\mathbf{z}}_{t+1} = \mathbf{a} + \mathbf{A} \tilde{\mathbf{z}}_t + \mathbf{u}_{t+1}, \tag{7}$$

where $\tilde{\mathbf{z}}_{t+1}$ is a state vector whose first elements are the excess log returns on the assets under consideration, and the rest are state variables that predict excess returns and variables that capture the dynamics of inflation and the short-term interest rate. The vector of innovations $\mathbf{u}_{t+1}$ is uncorrelated over time with conditional variance-covariance matrix $\mathbb{V}_t \left[\mathbf{u}_{t+1} \right]$.

The assumption of a first order for the VAR is not constraining because higher order vector autoregressions can be written as a VAR(1) through a straightforward change in the state vector. The return decomposition is sensitive to the particular specification of the components of the state vector (Chen and Zhao 2009). We specify our state vector to include variables for which there is wide consensus that capture time variation in risk premia, inflation, and real interest rates.

Given a specification for $\tilde{\mathbf{z}}_{t+1}$, it is straightforward to derive the components of the return decomposition as a function of the vector $\mathbf{u}_{t+1}$ of innovations to $\tilde{\mathbf{z}}_{t+1}$ and the parameters of the VAR(1). We perform this derivation in both Section 3 and Section 4.

3 Long-Run Portfolio Risk and Optimal Global Portfolio Diversification in a Symmetrical Model of Asset Returns

The return decomposition (4) implies that asset values and returns move over time in response to either changes in expected future cash flows or changes in discount rates. Therefore if asset returns are conditionally cross-sectionally correlated, it must be because either their cash flows or their discount rates, or both, are conditionally cross-sectionally correlated. This section shows that each type of correlation has a different impact on portfolio risk, portfolio choice, and the benefits of portfolio diversification at investment horizons beyond one period.

3.1 Model

To help fix ideas we consider a symmetrical model of investment opportunities, with N markets or assets (to which we will refer also as "countries") with identical return generating processes. This stylized model is particularly helpful because it allows to cleanly disentangle the effects of different types of cross-country news correlations on portfolio risk and portfolio choice at long horizons.

The dynamics of excess returns on each market i is given by the following single state variable VAR(1) model:

$$r_{i,t+1} = \mu_1 + \beta s_{i,t} + u_{i,t+1} \tag{8}$$

$$s_{i,t+1} = \mu_2 + \phi s_{i,t} + u_{si,t+1}, \tag{9}$$

where $r_{i,t+1}$ denotes the log return on country i, and $s_{i,t+1}$ denotes the state variable driving the time variation in the conditional expected return on country i, given by $\mathbb{E}_t[r_{i,t+1}] = \mu_1 + \beta s_{i,t}$. Without loss of generality we normalize $\beta > 0$. The parameters μ_1, μ_2, β, and ϕ are identical across countries, with $|\phi| < 1$ to preserve stationarity.

The within-country conditional variance-covariance matrix of the shocks to the VAR is also identical across countries and constant over time:

$$V_t\left[\mathbf{u}_{i,t+1}\right] = \begin{bmatrix} \sigma_{uu}^{wc} & \sigma_{us}^{wc} \\ \sigma_{us}^{wc} & \sigma_{ss}^{wc} \end{bmatrix}. \tag{10}$$

where $\mathbf{u}_{i,t+1} = (u_{i,t+1}, u_{si,t+1})'$ and the superscript wc denotes within-country quantities.

Finally, the conditional cross-country covariance matrix of VAR shocks between any pair of countries is also identical across country pairs and constant over time:

$$C_t\left[\mathbf{u}_{i,t+1}, \mathbf{u}_{j,t+1}\right] = \begin{bmatrix} \sigma_{uu}^{xc} & \sigma_{us}^{xc} \\ \sigma_{us}^{xc} & \sigma_{ss}^{xc} \end{bmatrix} \tag{11}$$

for all i and j. The superscript xc denotes cross-country quantities.

This stylized model of country returns defined by equations (8)-(11) implies that countries are identical and symmetrical with respect to the structure of their return dynamics and the cross-country correlation structure of returns and state variables. Of course the realized paths of returns and the state variable in each country will vary across countries. For example, in this model the expected excess return on country i is given by $\mu_1 + \beta s_{i,t}$, whose realizations depend on the realizations of the country-specific state variable $s_{i,t}$.

A straightforward application of the return decomposition (4) to the VAR(1) model (8)-(11) shows that the shocks to the model (8)-(9) are related to structural cash flow and discount rate shocks as follows:

$$N_{DR,i,t+1} = \lambda u_{si,t+1}, \tag{12}$$
$$N_{CF,i,t+1} = u_{i,t+1} + \lambda u_{si,t+1}, \tag{13}$$

with

$$\lambda = \frac{\rho\beta}{1-\rho\phi}.$$

(See Appendix.)

Therefore discount rate news are proportional to shocks to the state variable driving expected returns, with a proportionality constant λ which is increasing in the persistence (ϕ) of the state variable or expected returns, the loading of expected returns on the state variable (β), and the log-linearization parameter ρ. Note that when expected returns are constant, i.e., when $\beta = 0$, the constant λ is zero and all variation in returns is driven exclusively by cash flow news: $u_{i,t+1} = N_{CF,i,t+1}$.

Our assumptions about the conditional covariance structure of the innovations to the VAR (10)-(11), together with equations (12) and (13), imply that both within-country and cross-country conditional variances and covariances of news are constant over time and identical across countries. To fix notation, we write

$$C_t[N_{CF,i,t+l}, N_{CF,j,t+l}] \equiv \sigma^m_{CF,CF}, \tag{14}$$
$$C_t[N_{CF,i,t+l}, N_{DR,j,t+l}] \equiv \sigma^m_{CF,DR}, \tag{15}$$
$$C_t[N_{DR,i,t+l}, N_{DR,j,t+l}] \equiv \sigma^m_{DR,DR}, \tag{16}$$

where $m \equiv wc$ when $i = j$, and $m \equiv xc$ when $i \neq j$. For example, $\sigma^{xc}_{CF,CF}$ denotes both the conditional cross-country covariance of cash flows news.

3.2 Correlated Return News and the Portfolio Risk Across Investment Horizons

The symmetrical model of Section 3.1 provides a convenient framework to explore the impact of each type of return news on portfolio risk and portfolio choice across investment horizons.

Consider an equally-weighted portfolio of the N identical and symmetrical markets. This portfolio is optimal for a mean-variance investor who can invest only in these N risky markets. The risk of this portfolio at horizon k, defined as the conditional variance of the k-horizon log portfolio return normalized by the investment horizon, is a weighted average of the within-country conditional variance of k-horizon returns and the cross-country covariance of k-horizon returns:

$$\frac{1}{k}V_t[r_{p,t+k}^{(k)}] = \frac{1}{N}\frac{1}{k}V_t[r_{i,t+k}^{(k)}] + (1-\frac{1}{N})\frac{1}{k}C_t[r_{i,t+k}^{(k)}, r_{j,t+k}^{(k)}]. \tag{17}$$

where $r_{i,t+k}^{(k)} = \sum_{l=1}^{k} r_{i,t+l}$ is the log return at horizon k and

$$C_t[r_{i,t+k}^{(k)}, r_{j,t+k}^{(k)}] = \sum_{l=1}^{k} C_t[r_{i,t+l}, r_{j,t+l}] + 2\sum_{l=1}^{k-1}\sum_{m=1}^{k-l} C_t[r_{i,t+l}, r_{j,t+l+m}]. \tag{18}$$

A similar expression obtains immediately for $V_t[r_{i,t+k}^{(k)}]$ by noting that $V_t[r_{i,t+k}^{(k)}] = C_t[r_{i,t+k}^{(k)}, r_{i,t+k}^{(k)}]$. (Please refer to the Appendix for derivations of all expressions in this section.)[5]

We are interested in expressing the conditional within-country and cross-country moments of k-period returns as a function of the conditional moments of return news. A forward recursion of the dynamic equations of the VAR(1) model (8)-(9) shows that future one-period realized returns are given by

$$r_{i,t+l} - E_t[r_{i,t+l}] = N_{CF,i,t+l} - N_{DR,i,t+l} + \frac{\beta}{\lambda}\sum_{m=1}^{l-1}\phi^{m-1}N_{DR,i,t+l-m}, \tag{19}$$

where we have replaced the reduced-form shocks $u_{i,t+l}$ and $u_{si,t+l}$ with the structural shocks $N_{CF,i,t+l}$ and $N_{DR,i,t+l}$ using (12) and (13). Note that $\beta\lambda^{-1} = (1-\rho\phi)/\rho > 0$.

Equation (19) shows that conditional on information at time t, the realized return on an asset at time $t+l$ depends only on the contemporaneous cash flow shock, but it depends on the entire history of discount rate shocks between t and $t+l$ when expected returns are time varying and persistent, i.e., when both β and ϕ are not zero. Moreover, the contemporaneous discount rate shock impacts the return negatively, but the past history of discount rate news impacts the return positively. That is, a positive discount rate shock has an immediate negative impact on realized returns, but its effect reverses over time. This reflects the transitory nature of discount rate news: A positive shock to discount rates depresses asset valuations contemporaneously but, because it is a transitory shock, its impact eventually reverses back, driving future prices and returns up. The autoregressive coefficient ϕ determines the speed of this reversion.

[5] We normalize by k because $V_t[r_{p,t+k}^{(k)}]/k$ is a constant independent of investment horizon in the absence of return predictability. To see note from the definition of k-horizon log return that the moments on the right hand side of (17) are all proportional to k.

Using expression (19) it is straightforward to write the conditional moments of one-period returns in (18) as a function of the conditional moments of return news:

$$C_t[r_{i,t+l}, r_{j,t+l}] = [\frac{\beta^2}{\lambda^2}\frac{(1-(\phi^2)^{l-1})}{1-\phi^2}+1]\sigma^{xc}_{DR,DR} + \sigma^{xc}_{CF,CF} - 2\sigma^{xc}_{CF,DR}, \tag{20}$$

and

$$C_t[r_{i,t+l}, r_{j,t+l+m}] = \frac{\beta\phi^{m-1}}{\lambda}(\sigma^{xc}_{CF,DR} - \sigma^{xc}_{DR,DR}) + \frac{\beta^2\phi^m}{\lambda^2}\frac{1-(\phi^2)^{l-1}}{1-\phi^2}\sigma^{xc}_{DR,DR}, \tag{21}$$

for $l \geqslant 1$ and $m \geq 1$. Note that the moments of cash flow news enter only the contemporaneous covariance of returns and do so with a coefficient of one. The moments of discount rate news enter both the contemporaneous and the lead-lag covariances of returns, with coefficients that are a function of β and ϕ.

We can now compute the cross-country component of portfolio risk at horizon k in (17) as a function of the moments of news components of returns. Direct substitution of (20) and (21) into (18) gives:

$$\frac{1}{k}C_t[r^{(k)}_{i,t+k}, r^{(k)}_{j,t+k}] = \sigma^{xc}_{CF,CF} + \left[a(\mathrm{k})^2 + b(\mathrm{k})\right] \times \sigma^{xc}_{DR,DR} - 2 \times a(\mathrm{k}) \times \sigma^{xc}_{CF,DR}, \tag{22}$$

for $k > 1$ and

$$C_t[r_{i,t+1}, r_{j,t+1}] = \sigma^{xc}_{DR,DR} + \sigma^{xc}_{CF,CF} - 2\sigma^{xc}_{CF,DR}. \tag{23}$$

for $k = 1$. The coefficients $a(k) \equiv a(k;\beta,\phi,\rho)$ and $b(k) \equiv b(k;\beta,\phi,\rho)$ are given in the Appendix.

Equations (22) and (23) allow us to understand the impact of correlated cash flow news and discount rate news on portfolio risk across investment horizons. At a one period horizon ($k = 1$) we have that cross-country cash flow news covariances and cross-country discount rate news covariances have identical impact on the cross-country covariance of returns and portfolio risk at a one-period horizon. At horizons $k > 1$, equation (22) shows that each type of return news covariance has a different effect on portfolio risk.

In particular, the cross-country covariance of cash flow news $\sigma^{xc}_{CF,CF}$ has a coefficient of one at all horizons, implying that an increase in the cross-country covariance of cash flow news has identical impact on portfolio risk at all horizons. But the coefficient on $\sigma^{xc}_{DR,DR}$ and the coefficient on $\sigma^{xc}_{CF,DR}$ are a function of investment horizon. The Appendix shows that in the limit as investment horizon grows, the cross-country component of portfolio risk (22) converges to

$$lim_{k\to+\infty}\frac{1}{k}C_t[r^{(k)}_{i,t+k}, r^{(k)}_{j,t+k}] = \sigma^{xc}_{CF,CF} + \left(1 - \frac{1-\rho\phi}{\rho-\rho\phi}\right)^2 \times \sigma^{xc}_{DR,DR} - 2 \times \left(1 - \frac{1-\rho\phi}{\rho-\rho\phi}\right)^2 \times \sigma^{xc}_{CF,DR} \tag{24}$$

where the coefficient on $\sigma^{xc}_{DR,DR}$ is positive, smaller than one whenever $\rho > \phi$ and it is sufficiently close to one, and zero when $\rho = 1$.[6] These conditions hold in all the cases we consider in our empirical analysis.

Equation (24) shows that, unless discount rates are extremely persistent, the impact of a given increase in the cross-country covariance of discount rate news on portfolio risk at long horizons is smaller than a similar increase in the covariance of cash flow news. To see this, Figure 3 plots the coefficient on $\sigma^{xc}_{DR,DR}$ for values of β, ϕ, and ρ calibrated to U.S. data in our sample. The figure shows that, for this empirically relevant calibration, the coefficient on $\sigma^{xc}_{DR,DR}$ declines monotonically with investment horizon and rapidly approaches values well under 0.3 at horizons of 10 years or more, consistent with the intuition that correlated discount rate news matter less than correlated cash flow news for portfolio risk at long horizons.

A similar logic applies to the variation of the within-country component of portfolio risk across risk $V_t[r^{(k)}_{i,t+k}]$. From the symmetry of the model and $V_t[r^{(k)}_{i,t+k}] = C_t[r^{(k)}_{i,t+k}, r^{(k)}_{i,t+k}]$, it follows that:

$$\frac{1}{k} V_t[r^{(k)}_{i,t+k}] = \sigma^{wc}_{CF,CF} + \left[a(\mathrm{k})^2 + b(\mathrm{k})\right] \times \sigma^{wc}_{DR,DR} - 2 \times a(\mathrm{k}) \times \sigma^{wc}_{CF,DR}. \tag{25}$$

Of course, the within-country k-return portfolio variance (25) is also the k-horizon risk of a single-country portfolio. Campbell and Viceira (2005), Pastor and Stambaugh (2012), and others have studied the properties of this variance as a function of the moments of the VAR(1) shocks. Equation (25) writes it as a function of the moments of cash flow news and discount rate news. This derivation helps us gain intuition into why long-horizon portfolio risk per unit of time is declining in investment horizon when asset returns are predictable in empirical calibrations: Discount rate shocks are transitory shocks whose impact on long-run portfolio return variability is smaller than the impact of permanent cash flow shocks.

When returns are not predictable (i.e., $\beta = 0$), discount rate news are zero, and all return variation comes from cash flow news. In such case, (22) and (25) reduce to $\sigma^{xc}_{CF,CF}$ and $\sigma^{wc}_{CF,CF}$ respectively, which implies

$$\frac{1}{k} V_t[r^{(k)}_{p,t+k}] = \frac{1}{N} \sigma^{wc}_{CF,CF} + (1 - \frac{1}{N}) \sigma^{xc}_{CF,CF}.$$

That is, per-period portfolio risk is constant across investment horizons.

3.3 Illustrative Example

To illustrate the impact of each news component of unexpected returns on portfolio risk at different investment horizons, we have calibrated the VAR(1) model (8)-(9) to US excess stock returns, with the log dividend-price ratio as the state variable. We use these

[6]Note from that ρmeasures the importance of cash flow news and discount rate news far in the future for valuations and returns (see equation [3]), while ϕ determines the persistence of discount rate news. Therefore, the conditions $\rho > \phi$ and $\rho \to 1$ essentially say that correlated discount rate news do not matter for long-run portfolio risk when distant cash flows matter for valuation.

estimates to compute the volatility per period $\sqrt{V_t[r_{p,t+k}^{(k)}]/k}$ given in equation (17) of an equally-weighted portfolio of U.S. stock market clones under three different scenarios for the cross-country correlations of return news.

The first scenario, or baseline case, sets cross-country news correlations to zero so all markets are uncorrelated. The second scenario and the third scenario vary the cross-country correlation of cash flow news and discount rate news respectively while holding the cross-country correlation of one-period returns at the same value. The second scenario sets the cross-country correlation of cash flow news to its maximum admissible value and the cross-country correlation of discount rates to zero.[7] The third scenario sets the cross-country correlation of cash-flow news to zero, and the cross-country correlations of discount rate news to the value that implies the same one-period cross-country return correlation as in the second scenario.[8] All scenarios assume that discount rate news and cash flow news are uncorrelated, both within countries and across countries..[9]

Figure 4 plots annualized portfolio risk $\sqrt{V_t[r_{p,t+k}^{(k)}]/k}$ as a function of investment horizon for each of the three scenarios. Panel A plots portfolio risk for a portfolio of two countries, and Panel B for a portfolio of seven countries—the number of countries we consider in our empirical analysis. The figure shows that portfolio risk per unit of time declines as the investment horizon increases in each of the scenarios. This results from return predictability (Campbell and Viceira, 2005). In the absence of return predictability, the lines in each plot would be horizontal.

The figure shows that portfolio risk increases at all horizons when country returns become correlated as a result of correlated cash flow news. Moreover, the increase in portfolio risk is proportionally larger at long horizons or, equivalently, portfolio risk declines more slowly as investment horizon increases when cash flow news are correlated across markets. By contrast, portfolio risk increases proportionally less at long horizons when country returns become correlated as a result of correlated discount rate news. In fact, portfolio risk under correlated discount rate news converges rapidly to the risk under zero cross-country news correlations as the investment horizon increases.

Comparing across panels, Figure 4 shows that overall portfolio risk declines as the number of countries increases for all horizons. But the figure also shows that the differential

[7] In U.S. data, $\sigma_{DR,DR}^{wc}/\sigma_{CF,CF}^{wc} = 2.6$, that is, discount rate news are an order of magnitude more volatile than discount rate news. Holding this ratio to 2.6 for all countries and setting all cother cross-country correlations to zero, the maximum admissible value of the cross-country correlation of cash flow news that ensures that the overally variance-covariance matrix of shocks across all markets is positive semidefinite is 0.72. This in turn implies a cross-country correlation of returns of 0.09.

[8] This value is 0.11. It is much smaller because of the much larger volatiliy of discount rate news relative to cash flows news.

[9] In terms of the correlation structure to the innovations to the VAR, the first scenario implies zero cross-country correlations of unexpected returns and shocks to the state variables (see equations 12 and 13). The second scenario implies a positive cross-country correlation of unexpected stock returns and zero cross-country correlations of dividend-price ratio shocks. The Appendix provides the values of the coefficients.

effect on long-run portfolio risk of correlated cash flow news and correlated discount rate news also changes with the number of countries. That is, the reduction in long-run portfolio risk achieved by global portfolio diversification is much larger when the driver of return correlations is correlated discount rate news than when the driver is correlated cash flow news.

This stylized symmetrical model illustrates the main point of our argument. For a given increase in the short-run cross-country correlations of returns, the increase in overall portfolio risk is significantly smaller at long horizons when correlated discount rate news drives the increase than when correlated cash flow news drives it. Equivalently, the benefits of international portfolio diversification measured as a reduction on portfolio risk do not decline as much for long-horizon investors as they do for short-horizon investors when capital market integration (or correlated discount rates) is the source of increased cross-country return correlations. By contrast, the benefits of international portfolio diversification decline equally for all investors when real markets integration (or correlated cash flows) is the source of increased cross-country return correlations.

3.4 Optimal Global Portfolio Diversification Across Investment horizons

Our stylized symmetrical model is also helpful to understand the impact of financial and real market integration on optimal international portfolio diversification across investment horizons. We illustrate these effects using the model of optimal intertemporal portfolio choice of Jurek and Viceira (2011) in which an investor with power utility preferences over terminal wealth at a finite horizon faces a time-varying investment opportunity set described by a VAR(1) model for returns and state variables.

Formally, an investor with investment horizon k chooses the sequence of portfolio weights $\left\{\boldsymbol{\alpha}^{\boldsymbol{\tau}}_{\boldsymbol{t+k-\tau}}\right\}_{\tau=k}^{\tau=1}$ between time t and $(t+k-1)$ such that

$$\left\{\boldsymbol{\alpha}^{\boldsymbol{\tau}}_{\boldsymbol{t+k-\tau}}\right\}_{\tau=k}^{\tau=1} = argmax\ E_t\left[\frac{W^{1-\gamma}_{t+k}}{1-\gamma}\right] \tag{26}$$

subject to the intertemporal budget constraint

$$W_{t+1} = W_t\left(1+R_{p,t+1}\right), \tag{27}$$

$$R_{p,t+1} = \sum_{j=1}^{N} \alpha_{i,t}\left(R_{i,t+1} - R_{f,}\right) + R_{f,}, \tag{28}$$

where γ is the coefficient of relative risk aversion, $\boldsymbol{\alpha_t} = (\alpha_{1,t},\ldots,\alpha_{N,t})'$, $R_{i,t} = \exp\{r_{i,t}\} - 1$, and R_f is the risk-free rate, which we assume is constant. The dynamics of excess log returns in each market i follow the VAR(1) model (8)-(9), identical across markets.

This intertemporal portfolio optimization problem has an exact recursive solution up to a log-linear approximation to the log return on wealth (27)-(28) (Jurek and Viceira, 2011). The recursive solution is an affine function of the vector of states variables with coefficients that vary with investment horizon that has two components. The first component is equal to the optimal 1-period horizon allocation, which is the instantaneous mean-variance or "myopic" optimal portfolio. The second component, which is horizon dependent, reflects intertemporal hedging motives in optimal portfolio choice that arise only when investment opportunities are time varying. Therefore horizon effects enter portfolio choice exclusively through the optimal desire of the investor to hedge changes in investment opportunities.

Figure 5 plots the mean optimal portfolio allocation to risky assets (**??**) as a function of investment horizon for each of the scenarios we consider in Section 3.3. We add a risk free asset ("cash") to the menu of identical stock markets with returns calibrated to the U.S. stock market, and set the investor's coefficient of relative risk aversion to 5. Panel A presents results for two countries, and Panel B for seven. Note that the optimal portfolio allocation to risky assets is an equal weighted portfolio because all markets have identical return generating processes and cross-country correlations are identical for any pair of countries. Therefore we need to report only one mean optimal portfolio allocation for each scenario.

The intercepts in the figure reflect the one-period or instantaneously mean-variance efficient optimal allocation to risky assets. To facilitate interpretation, we set the unconditional expected returns and the risk-free rate such that the mean optimal allocation to cash is zero in the baseline scenario of zero cross-country return correlations. This implies a positive optimal allocation to cash, or equivalently a smaller optimal allocation to stocks, in the other two scenarios where the cross-country correlation of one-period stock returns is positive.[10]

Figure 5 shows that total portfolio demand for stocks is increasing in investment horizon in all three scenarios. This result is familiar from the literature that examines the optimal allocation to stocks at long horizons. Intertemporal portfolio choice is entirely driven by intertemporal hedging demand, which is positive in our calibration because shocks to the state variable—or equivalently expected excess returns—are negatively correlated with realized stock excess returns. That is, realized returns tend to be positive when expected excess returns fall, implying that a long position in the risky asset helps hedge against a fall in expected returns.

Figure 5 also shows that the intertemporal hedging demand for stocks is smaller in the scenarios with correlated market returns than in the benchmark case with zero-correlation. However, the extent to which this horizon effect is smaller depends crucially on the source of the cross-correlation of returns. Intertemporal hedging demands are significantly smaller when the source of short-run cross-country correlations of stocks returns is correlated cash flow news than then the source is correlated discount rate news.

This result is consistent with our results for portfolio risk across investment horizons. When discount rate news are correlated across markets but cash flow news are not, the

[10]It is also the same in both scenarios because recall that we set the cross-correlations of cash flow news and discount rate news in each scenario such that they imply identical one-period return cross-correlations.

scope for global diversification declines at short horizons, but much less so at long horizons. Long-horizon investors can still take advantage of a global portfolio to diversify cash flow risk, which is the most important risk at long horizons. At the same time, they can use global stocks and not just their local market to hedge against adverse changes in expected returns. By contrast, if cash flow news become correlated across markets, the scope for global diversification declines for all investors regardless of their investment horizon.

Figure 5 is also helpful to understand optimal intertemporal portfolio demand as the number of markets available for investing increases. The figure shows that the total optimal demand for stocks is independent of the number of markets available for investing at all horizons in the baseline case of uncorrelated markets. To see this, note that total portfolio demand obtains by multiplying by a factor of two the allocations in Panel A and by a factor of seven those in Panel B. As the number of markets increases, the investor distributes the total portfolio demand for stocks across more markets but the total portfolio demand for stocks, both myopic and hedging, remains unchanged. The scope for diversification of both discount rate risk and cash flow risk increases in the number of uncorrelated markets available for investing, as well as the ability to hedge adverse changes in expected returns.

Finally, Figure 5 shows that if returns are correlated across markets, the total optimal portfolio demand for stocks is a decreasing function of the number of markets available for investing at all horizons. The reduction in total portfolio demand is much larger when the source of cross-country return correlations is correlated cash flow news than when it is correlated discount rate news. The differential effect of each type of news comes through the intertemporal hedging demand, because the reduction in myopic demand is the same in both scenarios. It is straightforward to see from Figure 5 that total intertemporal hedging demand at the longest horizon in the plot declines from about 250% to about 220% in the correlated discount rate news scenario as we go from two to seven markets, and from about 190% to 70% in the correlated cash flow news scenario. The investor still distributes the total portfolio demand for stocks across more markets, but he does not see the increase in the number of markets as an opportunity to take on more overall portfolio risk as he is just adding more correlated—or less diversifiable—long run risk. But if the added risk is discount rate risk, the long-horizon investor understands this correlated risk has only a transitory impact on portfolio risk and he optimally reduces his overall risk exposure by much less than when the added risk is cash flow risk, which has a permanent impact on portfolio risk.

4 Sources of Return Correlation in Global Capital Markets

The stylized symmetrical model presented in Section 3 highlights the importance of understanding the sources of cross-country correlations of returns to evaluate the benefits of international portfolio diversification at long horizons. We now present an empirical analysis of the return news decomposition presented in Section 2 for stocks and government bond returns of seven major developed economies for the period January 1986 through December 2013. The countries included in our analysis are Australia, Canada, France, Germany, Japan, the U.K., and the U.S. These countries account for at least 80% of total global stock market capitalization throughout our sample period.

4.1 VAR Specification and Estimates

We estimate the VAR(1) model (7) for the seven countries using monthly data over the entire sample period 1986.01-2013.12. Our specification for the state vector for the VAR(1) includes the the log return on equities and bonds in excess of the return on their domestic T-bill,[11] variables that are known to predict excess returns—dividend-price ratios and yield spreads—, and variables that help capture the dynamics of real interest rates and inflation—nominal short-term interest rates and inflation (Campbell, Chan, and Viceira, 2003, Campbell and Viceira, 2005).

Specifically, we estimate a pooled VAR(1) model for the seven countries in our sample:

$$\tilde{\mathbf{z}}_{i,t+1} = \mathbf{a_i} + \mathbf{A}\tilde{\mathbf{z}}_{i,t} + \mathbf{u}_{i,t+1}, \tag{29}$$

where

$$\tilde{\mathbf{z}}_{i,t+1} = \left[xr_{s,i,t+1}, xr_{10,i,t+1}, d_{i,t+1} - p_{i,t+1}, \pi_{i,t+1}, y^N_{1,i,t+1}, y^N_{10,i,t+1} - y^N_{1,i,t+1}\right], \tag{30}$$

i denotes country, and $\mathbf{a_i}$ is a 6×1 vector of intercepts and $\mathbf{A}$ is a 6×6 slope coefficient matrix which is identical for all countries. We estimate a pooled VAR(1) model in an attempt to use as much cross-country information as possible to estimate the process for expected returns, since our sample is relatively short in the time series dimension. In practice, this procedure tempers the evidence of return predictability for those markets for which there is more in-sample evidence of return predictability, like the U.K. and the U.S.

In (30), $xr_{s,i,t+1}$ denotes the excess log return on equities in country i, $xr_{10,i,t+1}$ the excess log returns on the 10-maturity nominal government bond, $d_{i,t+1} - p_{i,t+1}$ the log of the dividend-price ratio, $\pi_{i,t+1}$ log inflation, $y^N_{1,i,t+1}$ the short-term nominal log interest rate, and $y^N_{10,i,t+1}$ the log yield on the 10-year nominal government bond. We measure excess log returns in each country as

$$xr_{i,t+1} = r^{\$}_{i,t+1} - y^N_{1,i,t} = \left(r^{\$}_{i,t+1} - \pi_{i,t+1}\right) - \left(y^N_{1,i,t} - \pi_{i,t+1}\right) \equiv r_{i,t+1} - r^N_{f,i,t+1}.$$

Finally, $\mathbf{u}_{i,t+1}$ is an i.i.d. 6×1 vector of shocks with within-country variance-covariance matrix $\sum_i^{wc}$ and cross-country covariance matrix $\sum_{i,j}^{xc}$, $i, j = 1, ..., 7$. We obtain monthly data for the state variables in all seven countries from a variety of sources. The Appendix provides a detailed description of the data and its sources.

4.2 Summary Statistics and VAR Estimates

Table 1, Table 2, and Table 3 present summary statistics for stock and bond returns over the entire sample period and for two subperiods of equal length, 1986.01-1999.12 and 2000.01-2013.12. This partition of the sample is motivated by our interest in exploring the sources

[11] This ensures that the return decomposition is currency independent (Campbell, Sefarty de Medeiros, and Viceira, 2010).

of the changes in cross-country stock and bond return correlations that have occurred during our sample period, illustrated in Figure 1, and their impact on international portfolio diversification across investment horizons.

Table 1 shows that the Sharpe Ratio of bonds in every country is significantly larger than the Sharpe Ratio of equities, both in the whole sample and in each subperiod, with the sole exception of the U.K. and the U.S. during the 1986-1999 period. The superior performance of bonds reflects a common long-term downward trend in nominal interest rates that has pushed bond prices higher throughout the entire period in all countries; by contrast, equity valuations have gone through periods of expansion and contraction, including the run-up in valuations in the 1990's and the sharp drop in valuations during the financial crisis of 2008 and 2009.

Across subsamples, the cross-country average excess bond return remained stable at 3.1% per annum, while the average excess stock return declined from 5.1% to 1.0% p.a. between the first and the second half of the sample period. Return volatility in each market and in each country has been fairly stable between subperiods—around 6% p.a. for bonds and 18% p.a. for stocks.

Table 2 reports the cross-country correlation matrix of bond and stock excess returns over the entire sample period and the two subperiods. Table 3 summarizes Table 2 and reports within-country and cross-country average excess return correlations. These tables complement Figure 1 and Figure 2. They show that cross-country return correlations have increased significantly from the early to the late subperiod for both stocks and bonds, and that the stock-bond correlation has switched sign from positive to negative.

Appendix E reports the estimates of the pooled VAR(1) model as well as estimates for each country. The top panel in each table reports coefficient estimates with t-statistics in parentheses and the R^2 statistic for each equation in the model. The bottom panel reports the correlation matrix of residuals, with the diagonal elements showing monthly standard deviations multiplied by 100 and the off-diagonal elements showing correlations.

We summarize here the estimation results. Our estimates reproduce the well-known result that the dividend-price ratio forecasts stock excess returns positively and that the short-term nominal rate forecasts stock excess returns negatively. Our estimates also reproduce the well-known result that yield spreads and short-term nominal interest rates have predictive power for bond excess returns, with positive coefficients. The equations for excess log returns exhibit the lowest R^2, which demonstrates the difficulty of predicting returns.

The estimates for the equations corresponding to the log dividend-price ratio, log inflation, the nominal short-term interest rate, and the log yield spread show that each variable is generally well-described by a univariate AR(1) process. The dividend-price ratio and the nominal short rate follow persistent processes. The yield spread and inflation follow less persistent processes, with inflation exhibiting the lowest persistence. As we will see, this has important implications for the benefits of global diversification of bond portfolios.

The correlation matrix of residuals shows a large negative average correlation between unexpected excess stock returns and shocks to the dividend-price ratio. We also estimate a negative but smaller average correlation between unexpected excess bond returns and shocks to the yield spread.[12] Because the dividend-price ratio and the yield spread are the main predictors of excess stock and bond returns, respectively, these negative correlations imply that shocks to expected excess returns are negatively correlated with realized excess returns. That is, stocks and bonds tend to do well when expected excess returns fall, thus providing investors with a hedge against deterioration in investment opportunities.

4.3 Estimates of News Components of Stock and Bond Excess Returns

We obtain estimates of the news components of stock and bond excess returns for each country implied by the estimates of the VAR(1) system (29)-(30).

Following standard practice in this literature, we have specified the state vector (30) such that we can explicitly identify unexpected stock excess returns and discount rate news—real rate and stock excess returns—from equations in the VAR, and obtain cash flow news obtain as the sum of unexpected excess returns and discount rate news. Specifically, Appendix A shows that the news components for stock returns given in (4)-(5) obtain from the VAR system as follows:

$$\begin{aligned}
xr_{s,t+1} - \mathbb{E}_t\left[xr_{s,t+1}\right] =& \mathbf{e1}'\mathbf{u}_{t+1}, \\
N_{RP,s,t+1} =& \mathbf{e1}'\left(\sum_{j=1}^{\infty}\rho_s^j\mathbf{A}^j\right)\mathbf{u}_{t+1}, \\
N_{RR,s,t+1} =& \mathbf{e5}'\left(\sum_{j=1}^{\infty}\rho_s^j\mathbf{A}^{j-1}\right)\mathbf{u}_{t+1} - \mathbf{e4}'\left(\sum_{j=0}^{\infty}\rho_s^j\mathbf{A}^j\right)\mathbf{u}_{t+1}, \\
N_{CF,s,t+1} =& xr_{s,t+1} - \mathbb{E}_t\left[xr_{s,t+1}\right] + N_{RR,s,t+1} + N_{RP,s,t+1},
\end{aligned}$$

where we omit the country subscript i for simplicity, and where $\mathbf{eL}$ denotes a column vector with a 1 in the $\mathbf{L}$ position and 0's in the rest.

We follow a different identification strategy for estimating the news components of bond excess returns. We explicitly identify bond cash flow news from the inflation equation in the VAR and obtain the risk premium or future expected excess returns component as the residual. Appendix A shows that the news components for excess bond returns given in (6) obtain from the VAR system as follows:

[12] Campbell, Chan, and Viceira (2003) and Campbell and Viceira (2005) report a positive estimate of this correlation for the U.S. in the postwar period up to the early 2000's.

$$
\begin{aligned}
xr_{n,t+1} - \mathbb{E}_t\left[xr_{n,t+1}\right] =& \mathbf{e2}'\mathbf{u}_{t+1}, \\
N_{CF,n,t+1} =& -\mathbf{e4}'\left(\sum_{j=1}^{n-1}\rho_b^j\mathbf{A}^j\right)\mathbf{u}_{t+1}, \\
N_{RR,n,t+1} =& \mathbf{e5}'\left(\sum_{j=1}^{n-1}\rho_b^j\mathbf{A}^{j-1}\right)\mathbf{u}_{t+1} - \mathbf{e4}'\left(\sum_{j=1}^{n-1}\rho_b^j\mathbf{A}^j\right)\mathbf{u}_{t+1}, \\
N_{RP,n,t+1} =& N_{CF,n,t+1} - N_{RR,n,t+1} - \left(xr_{n,t+1} - \mathbb{E}_t\left[xr_{n,t+1}\right]\right).
\end{aligned}
$$

We extract the news components of stock and bond excess returns for the entire sample period as well as for the subperiods 1986-1999 and 2000-2013. To compute news components for each subperiod, we use the full sample coefficient estimates $\mathbf{A}$ and split the residuals $\mathbf{u}_{t+1}$ into two subsamples. That is, we assume that the dynamics of returns and the state variables have been stable over the sample period. This assumption is motivated by the fact that the the state variables that capture expected excess returns, inflation, and the nominal short-rate follow highly persistent processes that require long samples to be precisely estimated. Additionally, we don't have strong priors about why the slopes of the VAR system might have changed throughout our sample period, while we do have priors about why the correlation structure of the shocks might have changed across subperiods. Accordingly, we use our entire sample period to estimate the slope coefficients.

4.4 News Decomposition of Cross-Country Correlations of Global Stock and Bond Excess Returns

Following Ammer and Mei (1996), we use our estimates of the news components of returns to explore the sources of the cross-country correlations of returns and their changes between the 1986-1999 subperiod and the 2000-2013 subperiod. We adopt their terminology and refer to cross-country cash flow news correlations as a measure of real economic integration, and to cross-country discount rate news—excess return news and real rate news—correlations as a measure of financial integration.

To get intuition for these two types of integration, consider a world in which capital markets are perfectly integrated, so there is a unique marginal investor pricing all assets. Since discount rates are determined by investors, we would expect discount rates to move synchronously in that world. Alternatively, we can also think of a world with integrated capital markets as a world in which shocks to investor risk aversion or investor sentiment propagate rapidly across markets. In either case, we expect discount rate news to be highly correlated across markets. By contrast, cash flows need not be perfectly correlated in such world, just like we don't expect the cash flows on two different stocks in the same market to be perfectly correlated, as they can be subject to idiosyncratic shocks in addition to aggregate shocks.

Now consider a world with a high degree of economic integration. In that case we expect aggregate shocks to affect all countries. To the extent that national stock markets are large enough to reflect their national economies and allow for diversification of company-specific idiosyncratic shocks, we expect country stock returns to reflect global shocks and country-wide shocks. Consequently, we expect a high degree of cross-country correlation of cash flow news.

Finally, we also expect local inflation to react in similar ways to global demand shocks in a world with a high degree of real economic integration, particularly if central banks react to those shocks in a similar manner. In that case we expect a high degree of cross-country correlation of innovations in inflation, i.e., in bond cash flow news.

Table 4 reports the average cross-country correlations of the news components of excess stock returns for each subperiod (upper left panel), p-values of the differences based on bootstrap and Fisher transformation methods (lower left panel), and the contribution of each component to the average cross-correlation of unexpected excess stock returns (right panel). Table 5 reports similar figures for bond excess returns. (See Appendix G for a description of the statistical tests and the calculation of the contribution of each component to total correlation.)

Table 4 shows that the source of the large increase in correlations in global cross-country correlations of stock excess returns from the early subperiod to the late subperiod has been an increase in the cross-country correlations of discount rate news. The cross-country correlations of both the real rate news component and the risk premium component of discount rate news have experienced increases which are significant both economically and statistically, from 0.33 to 0.59 and from 0.39 to 0.59, respectively. By contrast, the cross-country correlations of cash flow news have stayed fairly stable from one subperiod to another. The upper right panel of Table 4 shows that the the risk-premium component of stock returns is the main contributor to the total cross-country correlation of stock returns in both subperiods, with a large increase of its relative importance in the second subperiod, from 35% to 68%.

The results in Table 4 suggest that financial integration has been the main driver of the increase in cross-country correlations of stock returns since the turn of the century. These results also highlight the importance of accounting for time variation in discount rates to understand the second moments of returns. They also add to the evidence on time variation in expected returns, since in a world with constant discount rates, cash flow news is the only source of cross-country return correlations and their change.

Table 5 shows that, in contrast to stocks, the cross-country correlation of cash flow news—i.e., the negative of inflation news–has also been a significant contributor to the increase in cross-country bond return correlations from the early subperiod to the late subperiod. The increase in cross-country inflation news is significant both economically and statistically, from 0.29 to 0.50. Consistent with our findings for stock returns, the average cross-country correlation of real rate news has also risen significantly, from 0.30 to 0.57. The cross-country correlation of risk premium news has also increased, although more modestly from 0.21 to 0.37. This increase is nonetheless statistically significant under the bootstrap test and it is borderline significant under the Fisher r-to-z method. The upper right panel of the table shows that real rate news and risk premium news are the main contributors to the cross-country correlation of bonds returns in both subperiods, with a large increase in the second subperiod.

Overall, Tables 4 and 5 present strong evidence that financial integration has been a powerful driver of the increase in the cross-country correlation of stock and bond returns between 1986-1999 and 2000-2013. In the case of bonds, increased cross-country correlations of cash flow (or inflation) news has also been an important contributor to increased cross-country return correlations.

Table 6 and Table 7 explore the sources of the stock-bond correlation within countries (Table 6) and across countries (Table 7). Both tables show a switch from positive to negative in the sign of the stock-bond cash flow news correlation from the early period to the late period. This is one of the main drivers of the switch in the sign of the stock-bond return correlation shown in Figure 2 and Table 3. Since bond cash flow news is the negative of inflation news, this switch in correlation implies that inflation news has switched from behaving countercyclically in the early period to behaving procyclically in the late subperiod in our sample of countries.

The tables also show a significant increase in the correlation of bond risk premium news with stock cash flow news in the most recent subperiod, which is the second main driver of the switch in the sign of the stock-bond correlation. It suggests that investors demand lower risk premia on bonds in recessions—when stock cash flow news are negative—, consistent with bonds being considered by investors as safe assets in the most recent period. Both sets of results are consistent with the evidence shown in Viceira (2012) and Campbell, Sunderam and Viceira (2013) for U.S. stocks and bonds, and the economic drivers of bond risk explored in Campbell, Pflueger and Viceira (2015) for the U.S.

4.5 Robustness Check: Pukthuanthong and Roll (2009) Measure of Market Integration

Thus far we have used only cross-country correlations of returns and their news components as our metric for financial and real integration. Pukthuanthong and Roll (2009) argue that small cross-country correlations do not necessarily imply a lack of integration. For example, they argue that cross-country return correlations can be small even when countries are highly integrated if returns are explained by a global multifactor model and each country return loadings on these global factors differ. They propose using an alternative metric of integration: the R^2 from regressing returns on global factors estimated from a principal component analysis. A larger R^2 then corresponds to greater integration.

We apply the Pukthuanthong-Roll methodology to realized returns and the news components of returns. For a given return or news series, we find the first three principal components every year and the R^2 from a simple least squares regression. This methodology is particularly helpful to determine if the relatively low degree of cross-country correlations of cash flow news of stocks could be the result of a multifactor structure underlying these shocks instead of evidence of lack of integration in stock cash flows. For all other news components in stocks and bonds, the pair-wise correlations are already large in both subperiods and have increased significantly from the early to the late subperiod, suggesting that pair-wise correlations help capture integration for these components.

Table 8 reports average R^2 over the two subperiods for each series. Panel A corresponds to stocks, and Panel B corresponds to bonds. When excess returns $xr_{s,t+1}$ and $xr_{n,t+1}$ are the return series of interest, the R^2 increases from 0.60 to 0.74 and 0.59 to 0.75, respectively. This result suggests that the overall level of integration has risen between the two subperiods. Not surprisingly, a similar result holds when conducting the analysis using innovations in excess returns. The results for news terms lead to the same conclusions we have a achieved from the cross-country correlation analysis: We find a substantial increase in R^2 in each case except for stocks cash flow news, for which the increase in R^2 is the smallest in magnitude. This suggests a significant increase in the level of financial integration in the stock market and in the bond market from the early subperiod to the late subperiod.

5 The Impact of Real and Financial Integration on Portfolio Risk and Optimal Global Portfolio Diversification at Long Horizons

Section 4 presents robust empirical evidence of an economically and statistically significant increase in the cross-country correlations of stock and bond excess returns between 1986-1999 and 2000-2013, driven primarily by an increase in the cross-country correlations of discount rate news in both markets. It also documents a substantive increase in the cross-country correlation of inflation innovations that determine cash flow news of nominal bonds.

We have shown in Section 3 that an increase in cross-country return correlations affects portfolio risk and portfolio choice at long horizons differently depending on the source of such increase—correlated cash flow news or correlated discont rate news. Accordingly, we now explore the implications of our empirical results for portfolio risk and optimal international portfolio diversification at long horizons.

5.1 The Risk of Globally Diversified Stock and Bond Portfolios Across Investment Horizons

We start with an analysis of risk across investment horizons of all-equity and all-bond portfolios. We consider both equal-weighted (EW) and value-weighted (VW) portfolios of the seven markets in our sample. We set the weights for both the all-equity and the all-bond value-weighted portfolios equal to the relative stock market capitalization values at the inception of our sample in January 1986: 1.51% (Australia), 2.83% (Canada), 5.22% (France), 5.07% (Germany), 16.09% (Japan), 10.38% (U.K.), and 58.88% (U.S.). This choice of weights implies that the results for the value-weighted portfolios are largely dominated by the U.S.

market experience.

Figure 6 presents results for the VW and EW all-equity portfolios. Panel A in Figure 6 plots the percent annualized standard deviation of portfolio excess returns, $\sqrt{(12/k)\, V_t[xr^{(k)}_{p,t+k}]}$ times 100, implied by our VAR estimates for each sample as a function of investment horizon. The panel shows that the short-run risk of a globally diversified equity portfolio is about the same in both subperiods but that the long-run risk is systematically lower in the late sample than in the early sample. For example, at a 1-month horizon, the risk of the VW portfolio is similar in both samples—14.3% p.a. in the early sample and 14.8% in the late sample; at a 25-year horizon, the risk of the portfolio is 9.3% p.a. in the early sample and 7.8% p.a. in the late sample. This is an economically significant difference, especially when compounded over 25 years.

The portfolio risk decomposition (17) shows that to understand the changes in portfolio risk across investment horizons we need to separate cross-country return correlation effects from within-country return volatility effects. Panel B and Panel C report the results from performing this decomposition. Panel B in Figure 6 plots the VAR-implied percent annualized average within-country volatility of k-horizon excess returns, calculated as

$$ave\_std(k) = 100 \times \sqrt{\frac{12}{k}} \times \sum_{i=1}^{N} \left(\frac{w_i}{\sum_{i=1}^{N} w_i} \right) \sigma_i^k,$$

where $\sigma_i^k = V_t[xr^{(k)}_{i,t+k}]$, and w_i is equal to either 1/7 (EW portfolio) or market i capitalization weight (VW portfolio).

Panel B shows that the average within-country excess return volatility declines as the investment horizon increases in both samples. This reflects the well-known effect of stock return predictability at the country level. The panel shows that the average within-country return volatility is lower in the late sample across all investment horizons and that the difference in excess return volatility between the late sample and the early sample is larger at long horizons. This pattern reflects both a slightly lower short-run return volatility and, most importantly, a higher degree of stock return predictability (or mean-reversion) in the late sample relative to the early sample.[13] This is not surprising, since the early sample includes the second half of the 1990's, a period of a sharp rise in valuations relative to fundamentals that weakened the empirical evidence on return predictability, while the late sample includes the subsequent correction that strengthened the empirical evidence on stock return predictability (Cochrane, 2008).

Panel C in Figure 6 plots the percent average cross-country correlation of k-horizon excess

[13] Note that we keep the slope coefficients of the VAR the same across samples. Therefore this effect on within-country return volatility is essentially the result of the correlation of unexpected excess stock returns and shocks to expected excess stock returns becoming more strongly negative in the late sample. We also know from Table 1 that the volatility of one-period stock returns is also somewhat smaller in the late sample.

returns for each subperiod, calculated as

$$ave\_corr(k) = 100 \times \sum_{i=1}^{N} \sum_{j=i}^{N} \left(\frac{w_i w_j}{\sum_{i=1}^{N} \sum_{j=i}^{N} w_i w_j} \right) \gamma_{ij}^{k},$$

where $\gamma_{ij}^{k} = Corr_t[xr_{i,t+k}^{(k)}, xr_{j,t+k}^{(k)}]$, and w_i is equal to either 1/7 (EW portfolio) or market i capitalization weight (VW portfolio).

Panel C shows a highly significant increase in short-run cross-country excess return correlations in the late sample relative to the early sample from about 52% to about 72%, consistent with the evidence shown in Table 3 and Figure 1. However, the average cross-country excess return correlation declines as investment horizon increases, and it does so more rapidly for the late sample. At long horizons, both samples exhibit a similar average cross-country excess return correlation of about 35%.

The symmetric model of Section 3 is helpful to understand the patterns of cross-country excess return correlations within each sample and across samples. This model shows that both cash flow news correlations and discount rate news correlations contribute similarly to cross-county return correlations at short horizons, but that cash flow news correlations dominate cross-country correlations at long horizons. Cross-country excess return correlations decline at long horizons in each sample because, as Table 4 shows, discount rate news explain a large fraction of total cross-country excess return correlations, and the effect of correlated discount rate news on cross-country return correlations subsides at long horizons. Table 4 also shows that discount rate news are significantly more correlated across stock markets in the late sample, while cross-country cash flow news correlation are stable across periods. This has resulted in a sharp increase in short-horizon cross-country correlations of excess returns in the late sample relative to the early sample. However, as this effect subsides at long horizons, the average cross-country correlation of excess returns converges to that in the early sample.

In summary, both within-country effects (increased mean reversion in stock returns) and cross-country effects (increased cross-country correlations of discount rate news) explain the observed patterns in the risk of global equity portfolios across investment horizons and across periods shown in Panel A of Figure 6.

There results raise the question of what change in long-run equity portfolio risk we might have observed if only cross-country correlations of return news had changed from the early sample to the late sample, while within-country volatilities had remained stable. To address this question re-estimate the overall variance-covariance matrix of VAR innovations across all countries in the late subperiod subject to the constraint that the variance-covariance matrix of innovations for each country remains at the same values as in the early subperiod.[14]

[14]It is important to note that direct substitution of the within-country covariance matrices of VAR innovations in the late sample with those in the early sample does not guarantee that the resulting overall variance-covariance matrix is properly defined in the sense that it is a positive-definite matrix. To ensure

We then re-compute the same objects as in the unconstrained analysis.

Figure 7, which has a structure identical to that of Figure 6, reports the results from this exercise. Panel A in the figure shows that, holding within-country k-horizon return volatility the same across both subperiods (as shown in Panel B), internationally diversified equity portfolios have become riskier in the late period relative to the early period at horizons up to 12 years. Panel C shows that this is exclusively the result of a significant increase in cross-country return correlations at those horizons. At longer horizons, cross-country k-horizon return correlations are lower in the late subperiod, resulting in lower overall portfolio risk, consistent with the fact we have documented that discount rate news accounts for most of the increase in short-run cross-country return correlations. We interpret these results as evidence that the increase in short-run cross-country return correlations in the late sample has not resulted in increased risk of internationally diversified portfolios at long horizons.

Figure 8 presents results for global bond portfolios. Panel A in the figure shows that, similar to equities, the percent annualized standard deviation of portfolio excess returns is decreasing in investment horizon for both sample periods. However, unlike that of equities, the risk of internationally diversified bond portfolios is larger in the late sample at all investment horizons. Panel B shows that within-country effects cannot explain this increase, as the annualized average within-country volatility of k-horizon excess returns on bonds is very similar in both samples across all investment horizons. Thus the effect has to be entirely the result of changes in the cross-country correlations of the news components of excess bond returns. Panel C shows that indeed the average cross-country correlation of k-horizon excess returns on bonds is much larger in the late sample than in the early sample for all horizons.

Once again, the symmetric model in Section 3 and the empirical results in Table 5 help understand the patterns shown in Figure 8. Table 5 shows that increased cross-country correlations of cash flow news—i.e., inflation—are even more important than increased cross-country correlations of discount rate news in explaining the increase in short-run cross-correlations of bond returns. The model in Section 3 shows that cross-country correlations of news on the persistent component of returns impactt k-horizon return correlations similarly across all investment horizons. Therefore, the increase in the cross-country correlation of inflation has exacerbated the risk of internationally diversified bond portfolios at all horizons in the late period. Arguably the benefits of global portfolio diversification in bonds has declined in the late sample relative to the early sample.

5.2 Optimal Global Equity Portfolio Diversification at Long Horizons

Section 5.1 shows that the long-run risk of globally diversified equity portfolios, measured as the EW and VW portfolio return variance at long horizons, has not increased in the

this basic property of variance-covariance matrices we use semidefinite programming methods to re-estimate the cross-country components of the overall variance-covariance matrix subject to the constraint the within-country components take values equal those of the early sample. See Appendix G for a description of the semidefinite programming method we use.

2000-2013 period relative to the earlier 1986-1999 period, despite a significant increase in the cross-country correlations of one-period excess stock returns. This result holds even after controlling for the effects on long-run portfolio risk of declining long-run within-country return variances. By contrast, the long-run risk of globally diversified bond portfolios have increased. These results suggest that the changes in the correlations of global equity and bond markets that have taken place in the early XXI century have not diminished the benefits of global portfolio diversification for long-horizon equity investors, although they have diminished the benefits of global bond portfolio diversification.

We now explore this insight in the context of models of intertemporal portfolio choice. Specifically, we compute the optimal intertemporal portfolio allocations and expected utility implied by our estimates of return dynamics in each sample for two types of investors who can invest in cash and global equities. The first one is the investor we consider in Section 3.4, that is, an investor with power utility preferences over terminal wealth at a finite horizon to whom we will refer to as the "JV investor" (Jurek and Viceira, 2011). For calibration purposes we set the investment horizon of the JV investor to 20 years and the value of the coefficient of relative risk aversion to 5. The second investor is an infinitely lived investor with Epstein-Zin utility over intermediate consumption to whom we will refer to as the "CCV investor" (Campbell and Viceira 1999, and Campbell, Chan, and Viceira 2003). We set the elasticity of intertemporal substitution of consumption of this investor to one, the coefficient of relative risk aversion to 5, and the time discount factor to 0.92. This choice of parameters implies the the investor consumes optimally a constant fraction of his wealth every month equal to 8% annually.[15]

In order to explore optimal portfolio allocations we need to take a stand on unconditional expected returns and the risk free rate. In the spirit of the approach pioneered by Black and Litterman (1992), we set the vector of unconditional expected excess returns and the risk free rate such that the myopic or one-period mean-variance optimal portfolio allocation in the early sample equals either the EW equity portfolio or the VW equity portfolio described in Section 5.1, given the estimated variance-covariance of one-period returns. This assumption allows us to understand how optimal portfolio allocations change across investment horizons within each period, and across periods, for reasons related exclusively to changes in risk.

Table 9 reports optimal global equity portfolio allocations and Table 10 reports expected utility for the two investors for each of the subperiods. The first numerical column in each panel of Table 9 reports the mean optimal one-period (or mean-variance) allocation to stocks, which is the same for both investors. The second column and the third column report the vector of mean intertemporal hedging demands for the JV investor and the CCV investor respectively. Table 10 reports expected utility expressed as a certainty equivalent of wealth for a JV investor at two horizons (10 and 20 years), and expected utility per unit of wealth

[15]We solve for the optimal intertemporal portfolio allocation of this investor building on the approximate solution methods of Campbell and Viceira (1999) and Campbell, Chan, and Viceira (2003). They show that the optimal intertemporal portfolio policy for this investor is an affine function of the vector of state variables similar to the solution in Jurek and Viceira (2011) that has two componets, a myopic or one-period component and an intertemporal hedging component.

for the CCV investor. Each panel in Table 10 reports two values of certainty equivalent of wealth and expected utility for each investor. The first one is expected utility when the investor can invest only in the U.S. stock market. The second one is expected utility when the investor have access to all seven equity markets.

Panel A in Table 9 reports results for the early sample. By construction, the myopic allocation is 100% invested in either the EW equity portfolio or the VW equity portfolio. The total intertemporal hedging demand for stocks is positive and large for both investors, at about 104% for the JV investor and 67% for the CCV investor. The intertemporal hedging demand for the CCV investor is smaller than that for the JV investor because, although this investor is infinitely lived, his investment horizon is effectively shorter: The CCV investor consumes a fraction of his wealth every period, while the JV investor only consumes at the end of his long horizon of 20 years. Given our parametric assumptions, the duration of the consumption liabilities the CCV investor is funding out of his wealth is about 13.5 years, significantly shorter than that of the JV investor, which is 20 years. The relative composition of the intertemporal hedging allocation across markets is similar for both investors, with the U.S. stock market absorbing the largest portfolio share for both the EW and the VW portfolio.

Panel A in Table 10 shows large gains in expected utility for long-horizon investors from the ability to invest globally. The certainty equivalent of wealth for the JV investor and the expected utility of consumption per unit for the CCV investor are both an order of magnitude larger when these long-horizon investors have access to global equity markets than when they are able to invest only in the U.S. stock market. Moreoer, for the JV investor, the gains increase exponentially with investment horizon: At a 20-year horizon, the gain from having access to seven markets is proportionally much larger than at a 10-year horizon. These large benefits of diversification are consistent with those reported in Jarek and Viceira (2011, Tables VI and VIII) and Campbell, Chan, and Viceira (2003, Table 5) for U.S. investors who gain access to bonds when they can invest only in U.S. equities and cash.

Panel B in Table 9 reports the optimal equity portfolio allocations and expected utility implied by our estimates of the return generating process in the late sample, holding unconditional expected excess returns and the risk free rate at the same values as in the early sample. The increase in the cross-country correlations of one-period returns generates a one-period myopic allocation with long and short positions. For example, the VW portfolio shows a significant increase of the short-run allocation to U.K. and Australian equities, funded by a short position in German equities and cash. The investors optimally choose levered myopic equity portfolios, illustrating the fact that increased correlations do not necessarily imply less willingness to hold risky assets in a portfolio in the absence of borrowing and short-selling constraints. Panel B also shows a significant increase in intertemporal hedging demands for stocks in the late sample, at 140% for the JV investor and 90% for the CCV investor in the EW case. The corresponding panel in Table 10 shows that expected utility also increases dramatically for both investors relative to the early sample.

The portfolio risk decomposition of Section 5.1 is helpful to understand the changes in intertemporal hedging demand and in expected utility in the late period with respect to the early period. We have shown that the late sample is characterized by both a significant increase in cross-country correlations of one-period stock excess returns and a significant decrease in within-country volatility of stock excess returns at all horizons resulting from increased mean reversion. The second factor implies more willingness to hold risky assets in a portfolio for intertemporal hedging reasons and, as shown in Campbell and Viceira (1999), it also implies large increases in expected utility. Therefore, within-country effects could explain the changes in intertemporal hedging demands and in expected utility across periods.

To test this hypothesis, Panel C in Table 9 reports optimal equity portfolio allocations in the late sample holding constant within-country stock return predictability across samples. Specifically, as in Section 5.1, we use semidefinite programming methods to re-estimate the overall variance-covariance matrix of VAR innovations across all countries in the late subperiod subject to the constraint that the elements of the within-country variance-covariance matrix of innovations for each country remain at the same values as in the early subperiod.

Panel C in Table 9 shows that, under this constrained estimation, optimal myopic portfolio demand in the late sample is somewhat smaller than that in the early sample, at 96% and 84% for the VW and EW portfolios respectively. Moreover, intertemporal hedging demands stay about the same across both subperiods. The corresponding panel in Table 10 shows that expected utility also remains at about the same values. This implies that the increases reported in Panel B are exclusively a result of within-country effects, that is, of increased mean reversion in stock returns at the country level in the late subperiod.

The results in Panel C of Table 9 and Table 10 also imply that the increase in cross-country return correlations of stock returns in the late sample has not diminished the benefits of global equity diversification for long-term investors. This is so because this increase has been driven primarily by financial integration or increased correlation of discount rates, and we have shown in Section 3 that increased correlations of transitory discount rate news do not imply an increase in portfolio risk at long horizons or a decline in the willingness of long-term investors to hold risky assets. Because permanent cash flows have not become significantly more correlated across countries in the late sample, long-run investors can still benefit from global investing to diversify cash flow risk.

6 Conclusions

We have documented a substantial secular increase in the cross-country correlations of global stock and bond returns since the turn of the 21st century and explored its implications for long-run portfolio risk, optimal intertemporal global portfolio choice, and the benefits of global portfolio diversification in a framework with time-varying discount rates in which asset valuations and returns vary over time in response to cash flow news and to discount rate news, both of which can be correlated across markets. We find that although this increase implies a reduction in the benefits of global portfolio diversification for short-horizon investors, it does not imply a reduction of these benefits for long-horizon equity investors.

Long run portfolio risk has not increased, optimal long-horizon equity portfolios are as globally diversified and invest in equities as much as in the prior period with lower cross-country return correlations, and the expected utility from holding global equity portfolios has not declined for long-horizon investors.

To understand these results, we have built a stylized symmetrical model of global capital markets, and shown that an increase in the cross-country correlation of discount rate news has only a small impact on the long-run risk of a globally diversified portfolio and in the willingness of long-horizon investors to hold risky assets in their portfolios, while an increase in the correlation of cash flow news has a very significant impact on both long-run portfolio risk and willingness to hold risky assets. This differential impact on long-run portfolio risk derives from the fact that cash flow shocks are highly persistent shocks that affect valuations and returns at all horizons, while discount rate shocks are transitory shocks whose impact on valuations and returns dissipates at long horizons. We have shown that the effect of an increase in the cross-country correlation of cash flow news leads to a similar increase in the cross-country correlations of returns at all horizons, while the effect of an increase in the cross-country correlations of discount rate news declines as investment horizon increases. A similar argument explains why an increase in the correlation of cash flow news has a much stronger negative impact on optimal portfolio demands at long horizons than a similar increase in the correlation of discount rate news. When discount rate shocks become more correlated across markets but cash flow shocks don't, long horizon investors can still use global diversification to attenuate long-run cash flow risk, which is the most relevant risk to them. By contrast, short-horizon investors care equally about both about discount rate risk and cash flow risk.

We have shown that empirically the increase in cross-country correlations of global equity markets since the turn of the century has been driven primarily by an increase in the correlation of discount rate news, which we attribute to the integration of global capital markets. By contrast, we don't find strong evidence that the globalization of trading flows has led to an increase in the cross-country correlation of cash flow shocks and through that to a corresponding increase in the cross-country correlations of stock returns. Consistent with our stylized model of global capital markets, we have shown that this evidence implies that the globalization of capital markets has not lead to a significant increase in the long-run risk of globally diversified equity portfolios nor to a decline in the expected utility and the willingness of long-horizon investors to hold globally diversified equity portfolios.

Moreover, we have also documented a decline in the long-run volatility of stock returns at the country level which has resulted in a reduction in long-term portfolio risk and an increase in the benefits of holding equities for long-term investors. Both factors together suggest that, if anything, the benefits of holding global equity portfolios for long-horizon investors have increased since the turn of the century. Long-term investors can still use global portfolios to diversify away long-term equity cash flow risk.

By contrast, we find that the significant increase in the cross-country correlation of bond returns has been driven by both increased correlation of discount rate news resulting from global capital markets integration, and increased correlation of nominal bond cash flow news resulting from increased correlation of inflation across monetary areas. Long-run cross-country bond return correlations have increased as much as short-run correlations, implying that the benefits of international bond portfolio diversification have declined as much for long-horizon long-only bond investors as for short-horizon investors.

However, the increased correlation of global bond markets at short and long horizons is beneficial to investors with long-dated liabilities. The scope for hedging liabilities using global bonds has increased. This can be particularly beneficial to investors with large long-dated liabilities whose own domestic bond markets are small.

Finally, we have shown that the well documented negative stock-bond correlation in the U.S. since the turn of the century is a global phenomenon. We have shown that this correlation is negative not only within countries but also across countries, suggesting that the benefits of stock-bond diversification have increased in all developed markets in recent times.

7 References

References

[1] Ammer, J. and Mei, J. (1996). Measuring international economic linkages with stock market data. *The Journal of Finance*, 51(5):1743–1763.

[2] Asness, C. S., Israelov, R., and Liew, J. M. (2011). International diversification works (eventually). *Financial Analysts Journal*, 67(3):24–38.

[3] Bekaert, G., Hoyem, K., Hu, W.-Y., and Ravina, E. (2016). Who is internationally diversified? evidence from the 401 (k) plans of 296 firms. *Journal of Financial Economics*.

[4] Black, F. and Litterman, R. (1992). Global portfolio optimization. *Financial analysts journal*, 48(5):28–43.

[5] Campbell, J. Y. (1991). A variance decomposition for stock returns. *The Economic Journal*, 101(405):157–179.

[6] Campbell, J. Y., Chan, Y. L., and Viceira, L. M. (2003). A multivariate model of strategic asset allocation. *Journal of financial economics*, 67(1):41–80.

[7] Campbell, J. Y., Pflueger, C., and Viceira, L. M. (2014). Monetary policy drivers of bond and equity risks. Technical report, National Bureau of Economic Research.

[8] Campbell, J. Y., Serfaty-De Medeiros, K., and Viceira, L. M. (2010). Global currency hedging. *The Journal of Finance*, 65(1):87–121.

[9] Campbell, J. Y. and Shiller, R. J. (1988). The dividend-price ratio and expectations of future dividends and discount factors. *Review of financial studies*, 1(3):195–228.

[10] Campbell, J. Y., Shiller, R. J., Viceira, L. M., Mishkin, F. S., and Wright, J. H. (2009). Understanding inflation-indexed bond markets/comments and discussion. *Brookings Papers on Economic Activity*, (1):79.

[11] Campbell, J. Y., Sunderam, A., and Viceira, L. M. (forthcoming). Inflation bets or deflation hedges? the changing risks of nominal bonds. *Critical Finance Review*.

[12] Campbell, J. Y. and Viceira, L. M. (1999). Consumption and portfolio decisions when expected returns are time varying. *The Quarterly Journal of Economics*, 114(2):433–495.

[13] Campbell, J. Y. and Viceira, L. M. (2002). *Strategic asset allocation: portfolio choice for long-term investors*. Oxford University Press, USA.

[14] Campbell, J. Y. and Viceira, L. M. (2005). The term structure of the risk–return trade-off. *Financial Analysts Journal*, 61(1):34–44.

[15] Campbell, J. Y. and Vuolteenaho, T. (2004). Bad beta, good beta. *The American Economic Review*, 94(5):1249–1275.

[16] Cochrane, J. H. (2008). The dog that did not bark: A defense of return predictability. *Review of Financial Studies*, 21(4):1533–1575.

[17] Cochrane, J. H. (2011). Presidential address: Discount rates. *The Journal of Finance*, 66(4):1047–1108.

[18] French, K. R. and Poterba, J. M. (1991). Investor diversification and international equity markets. *American Economic Review*, 81.2:222–226.

[19] Greenwood, R. and Vayanos, D. (2014). Bond supply and excess bond returns. *Review of Financial Studies*, 27(3):663–713.

[20] Grubel, H. G. (1968). Internationally diversified portfolios: welfare gains and capital flows. *The American Economic Review*, 58(5):1299–1314.

[21] Hamilton, J. D. and Wu, J. C. (2012). The effectiveness of alternative monetary policy tools in a zero lower bound environment. *Journal of Money, Credit and Banking*, 44(s1):3–46.

[22] Jurek, J. W. and Viceira, L. M. (2011). Optimal value and growth tilts in long-horizon portfolios. *Review of Finance*, 15(1):29–74.

[23] Markowitz, H. (1952). Portfolio selection. *The journal of finance*, 7(1):77–91.

[24] Merton, R. C. (1969). Lifetime portfolio selection under uncertainty: The continuous-time case. *The review of Economics and Statistics*, pages 247–257.

[25] Philips, C. B. (2014). Global equities: Balancing home bias and diversification. *Vanguard research working paper, The Vanguard Group*.

[26] Pukthuanthong, K. and Roll, R. (2009). Global market integration: An alternative measure and its application. *Journal of Financial Economics*, 94(2):214–232.

[27] Quinn, D. P. and Voth, H.-J. (2008). A century of global equity market correlations. *The American Economic Review*, 98(2):535–540.

[28] Samuelson, P. A. (1969). Lifetime portfolio selection by dynamic stochastic programming. *The review of economics and statistics*, pages 239–246.

[29] Santis, G. and Gerard, B. (1997). International asset pricing and portfolio diversification with time-varying risk. *The Journal of Finance*, 52(5):1881–1912.

[30] Solnik, B. H. (1974). An equilibrium model of the international capital market. *Journal of economic theory*, 8(4):500–524.

[31] Sunderam, A., Viceira, L., and Ciechanover, A. (2015). The vanguard group, inc. in 2015: Celebrating 40. *Harvard Business School Case 216-026*.

[32] Viceira, L. M. and Ciechanover, A. (2016). The wealthfront generation. *Harvard Business School Case 216-085*.

[33] Vuolteenaho, T. (2002). What drives firm-level stock returns? *The Journal of Finance*, 57(1):233–264.

Figure 1: Stock and Bond Correlations Across Countries

This figure plots average correlations of stock returns across countries and bond returns across countries. Monthly averages are computed using pairwise return correlations across seven different countries over 3-year rolling windows (Australia, Canada, France, Germany, Japan, United Kingdom, and United States). Returns are in U.S. Dollar currency-hedged terms in excess of the three-month U.S. Treasury bill rate. The sample is from Jan 1986 to Dec 2013.

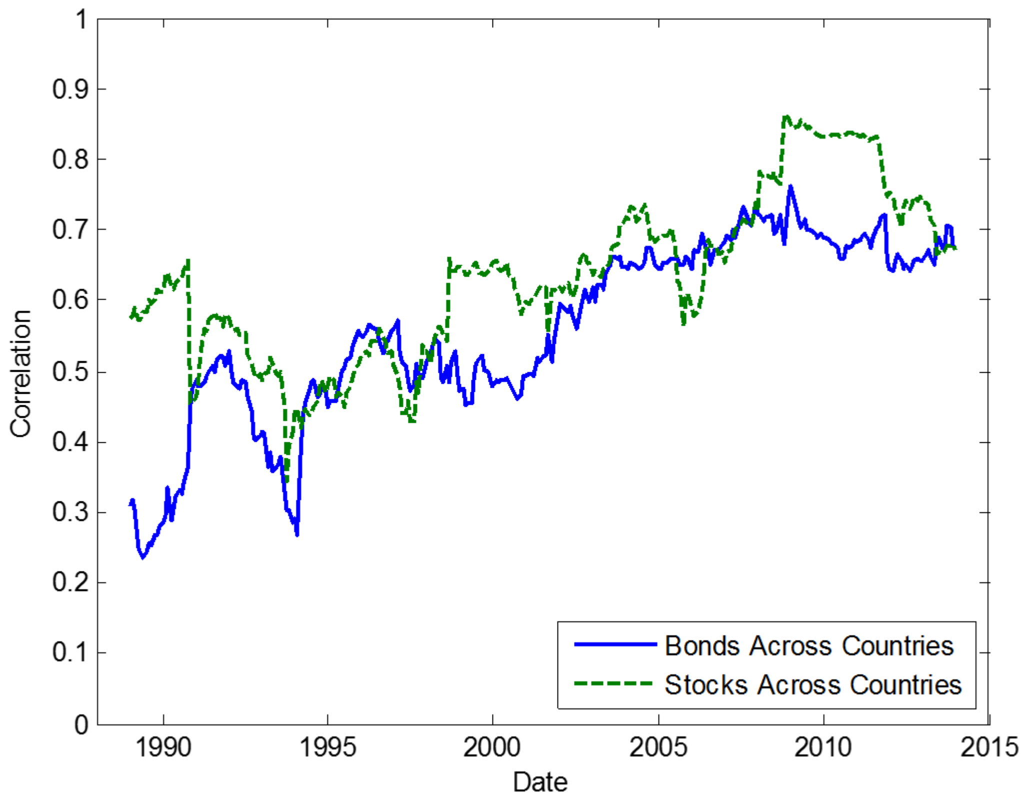


Figure 2: Stock-Bond Correlation Across and Within Countries

This figure plots average stock-bond correlations across countries and within countries. Monthly averages are computed using pairwise return correlations within and across seven different countries over 3-year rolling windows (Australia, Canada, France, Germany, Japan, United Kingdom, and United States). Returns are in U.S. Dollar currency-hedged terms in excess of the three-month U.S. Treasury bill rate. The sample is from Jan 1986 to Dec 2013.

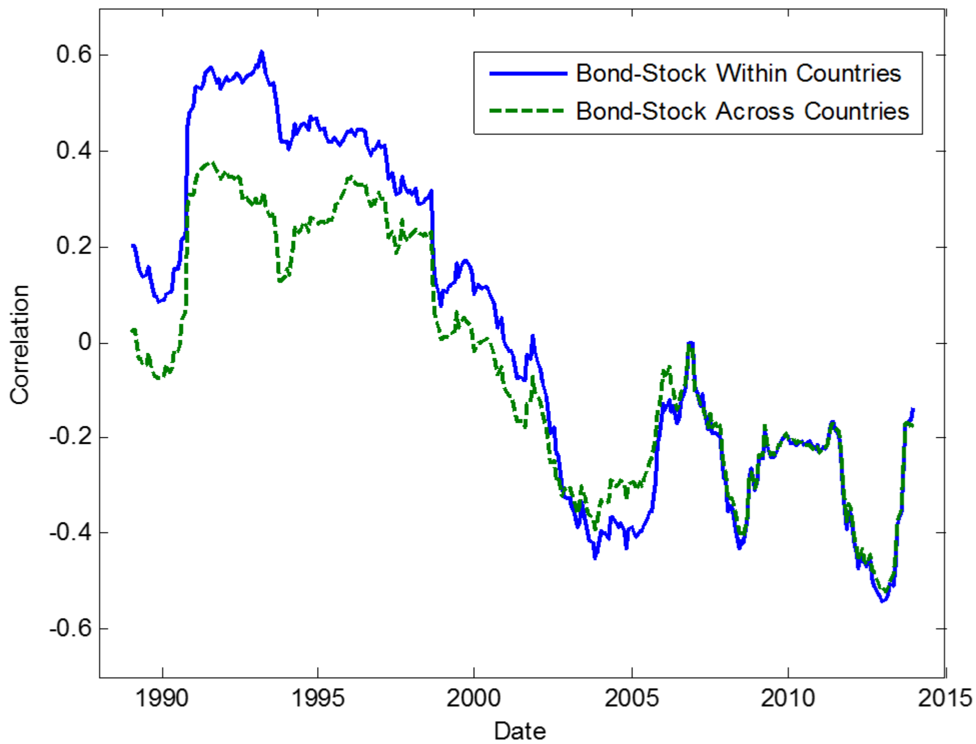


Figure 3: Coefficient on $\sigma^{xc}_{DR,DR}$ as a function of investment horizon k

The figures plots the coefficient on $\sigma^{xc}_{DR,DR} = a(k;\beta,\phi,\rho)^2 + b(k;\beta,\phi,\rho)$ as a function of investment horizon k. We use parameters estimated from U.S. data for calibration: $\beta = 0.0121$, $\phi = 0.9864$, $\rho = 0.9982$. The expressions for $a(k;\beta,\phi,\rho)$ and $b(k;\beta,\phi,\rho)$ are given in the Appendix.

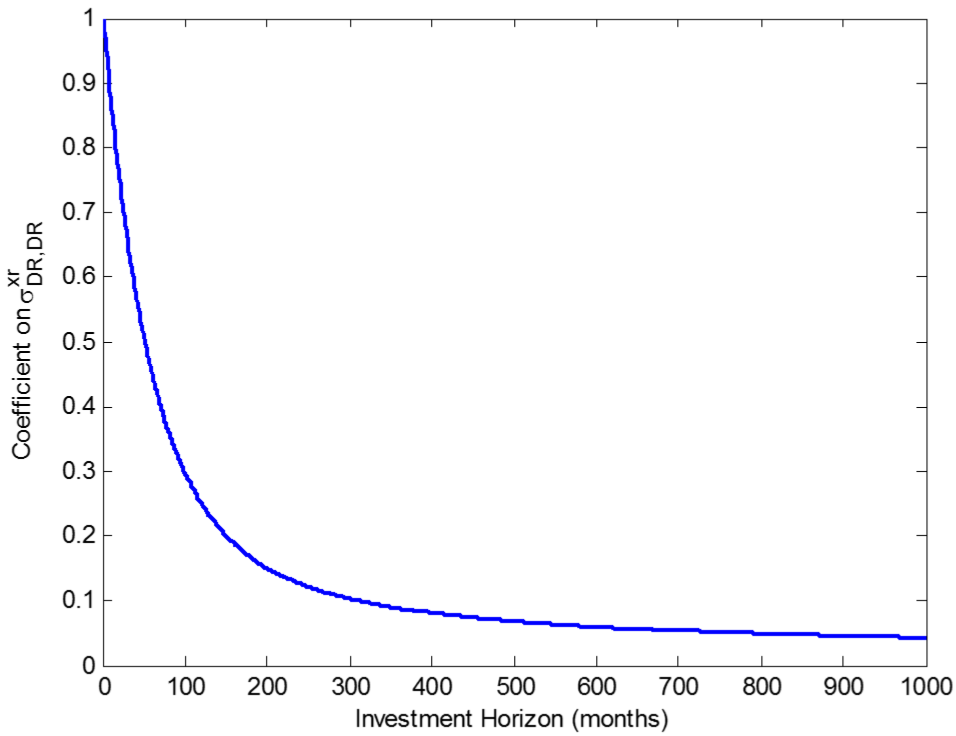
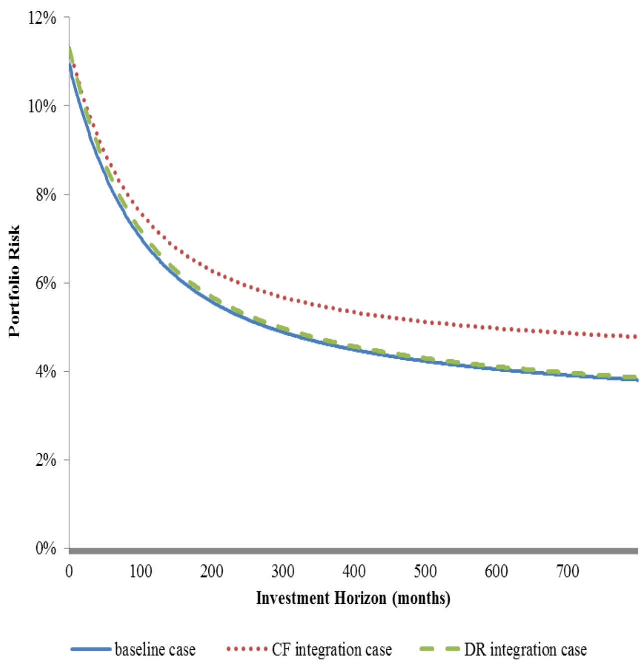
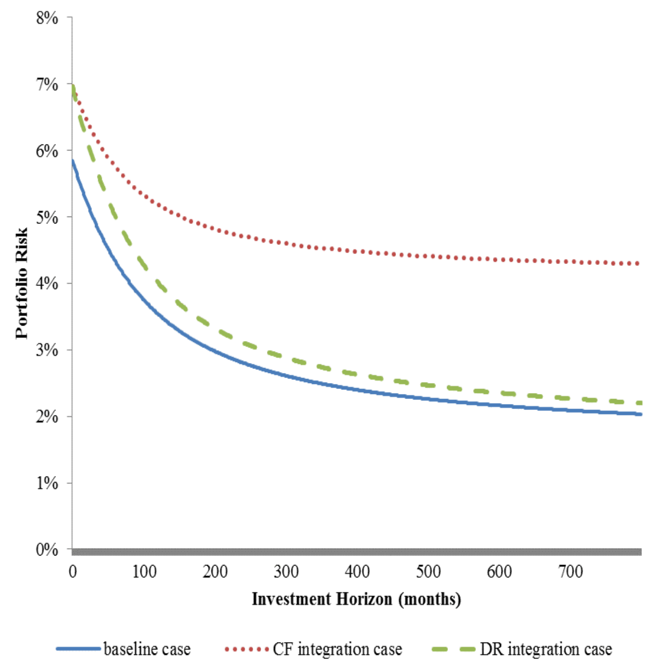


Figure 4: Annualized portfolio risk as a function of investment horizon

The figure plots $\sqrt{V_t(r^{(k)}_{p,t+k})/k}$ as a function of investment horizon k (months). We compare the term structure of this conditional standard deviation for 3 scenarios: (1) Baseline case with zero cross-country return news correlations, both for CF news and DR news. (2) CF news integration case, where cross-country return correlations come from positive cross-country CF news correlations; cross-country correlations of DR news are zero. (3) DR integration case, where cross-country return correlations come from positive cross-country DR news correlation; cross-country correlations of CF news are zero. To make Scenarios 2 and 3 comparable, we set the cross-country correlation of one-period returns at the same value (0.07). Panel A plots portfolio risk in each scenario for a portfolio of two symmetric countries, and Panel B for a portfolio of seven countries.



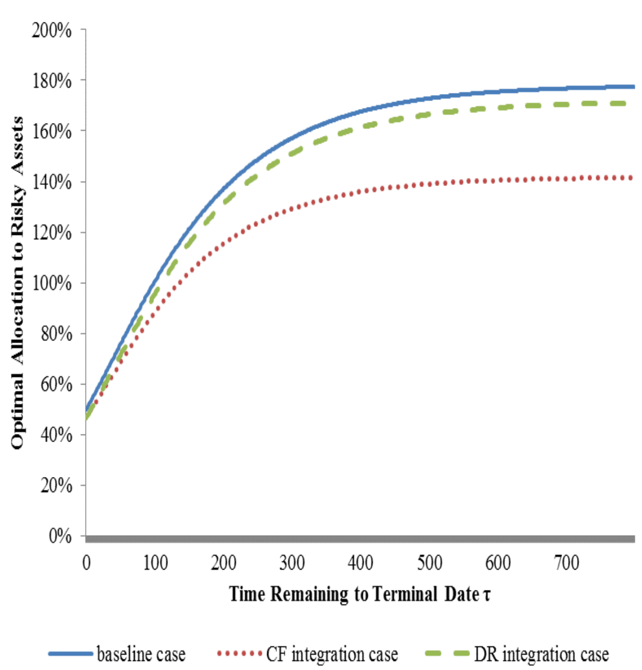
(a) Panel A: two symmetric countries



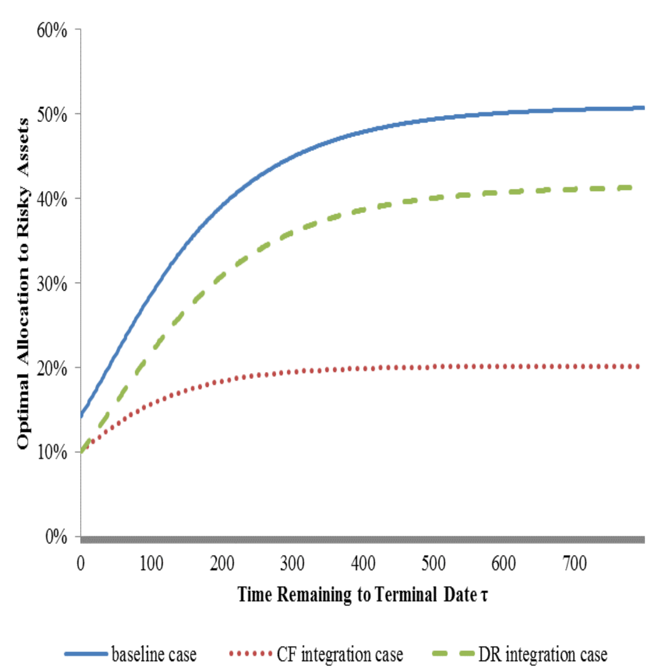
(b) Panel B: seven symmetric countries

Figure 5: Optimal allocation to risky assets as a function of time remaining to terminal date

The figure plots optimal allocation to risky assets as a function of time remaining to terminal date. The total optimal allocation is the sum of two parts: myopic allocation (equals the intercept at $\tau = 1$) and hedging allocation. The investor has horizon of $k = 800$ and rebalance his allocation each period. The x-axis τ is the time remaining to the terminal date. We compare the term structure of optimal allocation to risky assets across the same 3 scenarios described in Figure 4. We set the expected excess returns so that in the benchmark case, the myopic investor ($\tau = 1$) allocate $1/N$ to each risky asset (50% for $N = 2$ and 14.3% for $N = 7$) and zero to cash. The expected excess returns are kept the same across the three cases to make them comparable. Panel A plots optimal allocation in each case for a portfolio of two symmetric countries, and Panel B for a portfolio of seven countries.



(c) Panel A: two symmetric countries

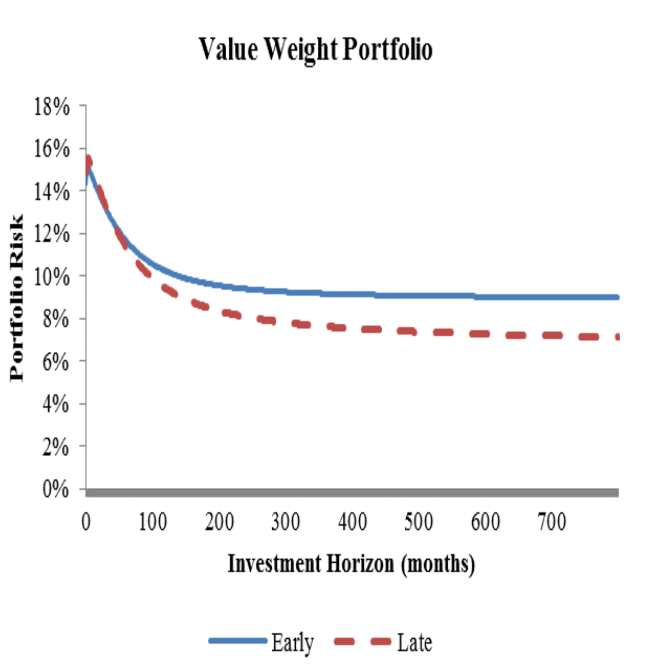


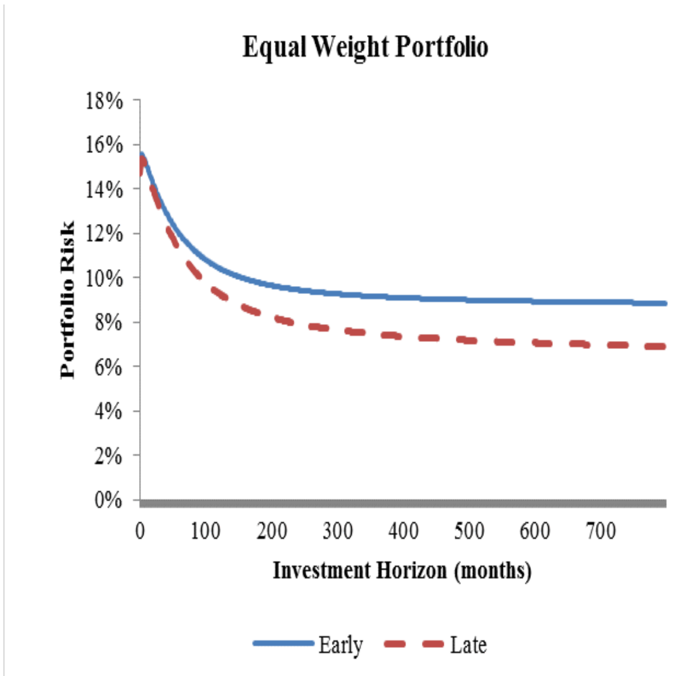
(d) Panel B: seven symmetric countries

Figure 6: All-equity portfolio risk as a function of investment horizon

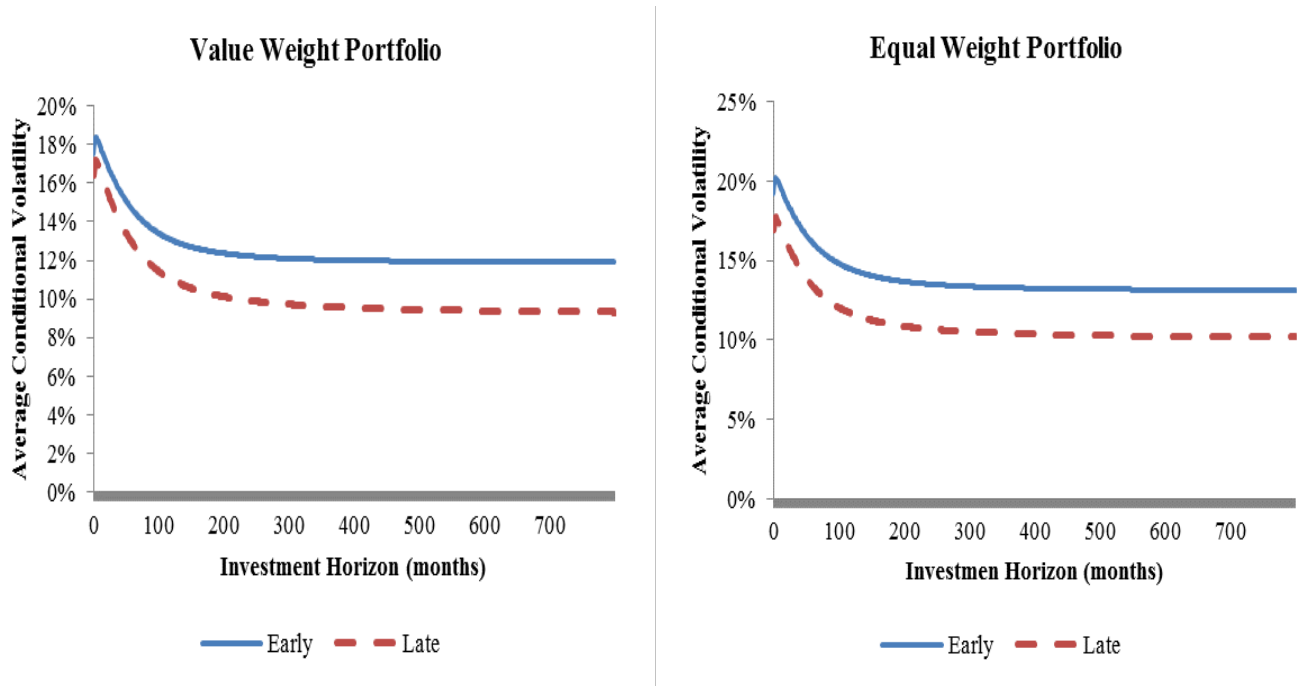
The figure compares the early sample (1986.01-1999.12) and late sample (2000.01-2013.12) all-equity portfolio risk across investment horizons. Panel A plots the annualized conditional standard deviation of portfolio excess returns. Panel B plots the average (across N countries) annualized k-period conditional volatility of excess returns. Panel C plots the pairwise average of cross-country k-period excess returns conditional correlation. Each panel includes the results for value-weighted and equal-weighted portfolios.

Panel A: Annualized conditional standard deviation of portfolio excess returns





Panel B: Average annualized k-period conditional volatility of excess returns



Panel C: Pairwise average of cross-country k-period excess returns conditional correlation

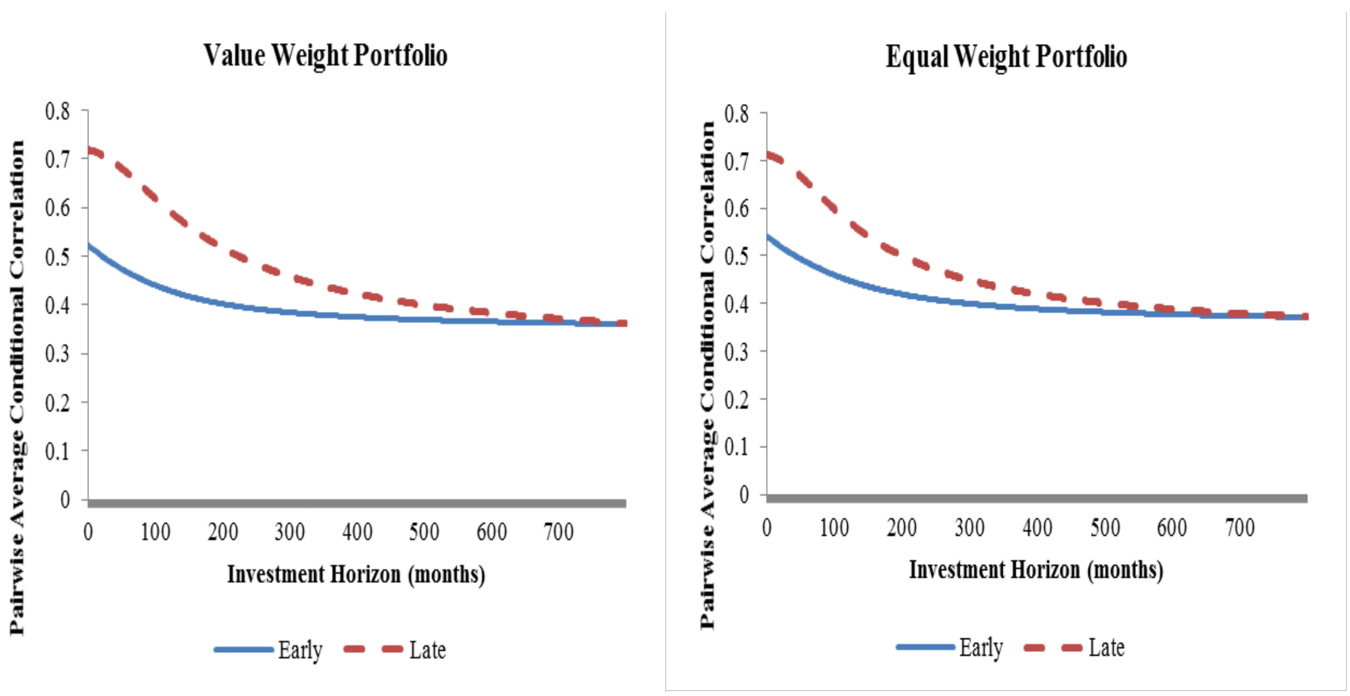
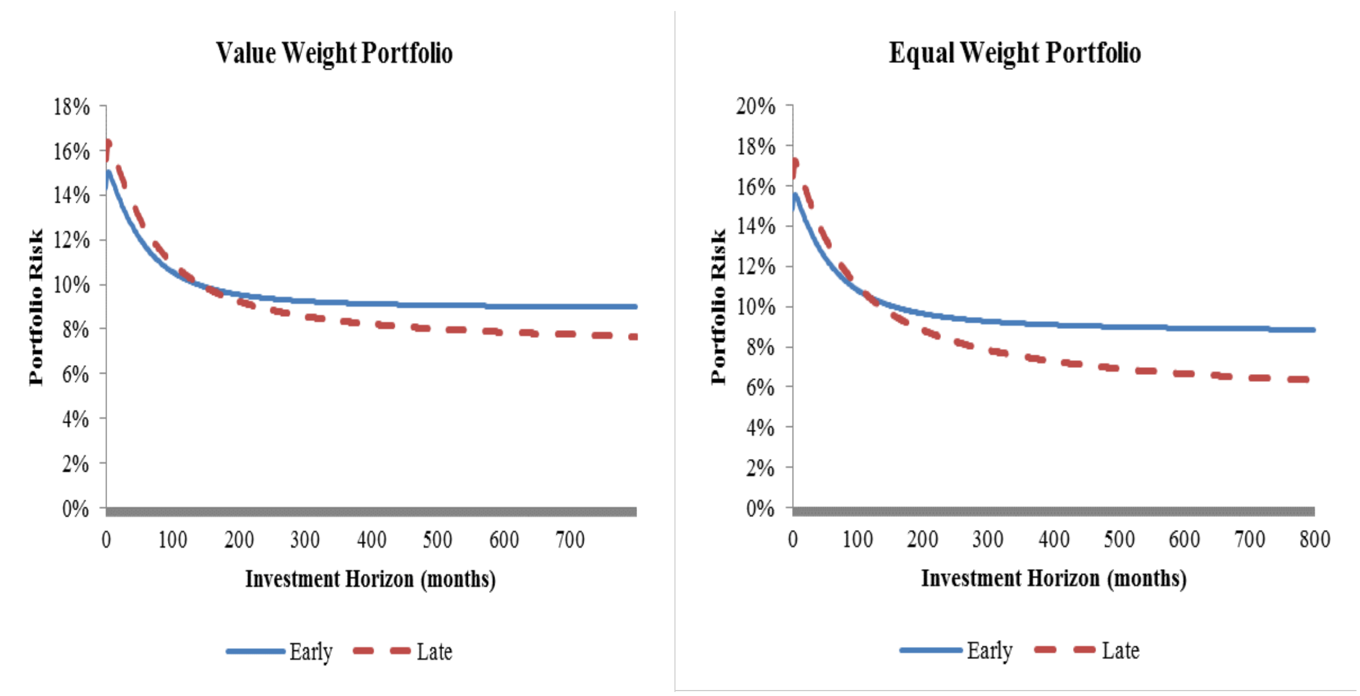


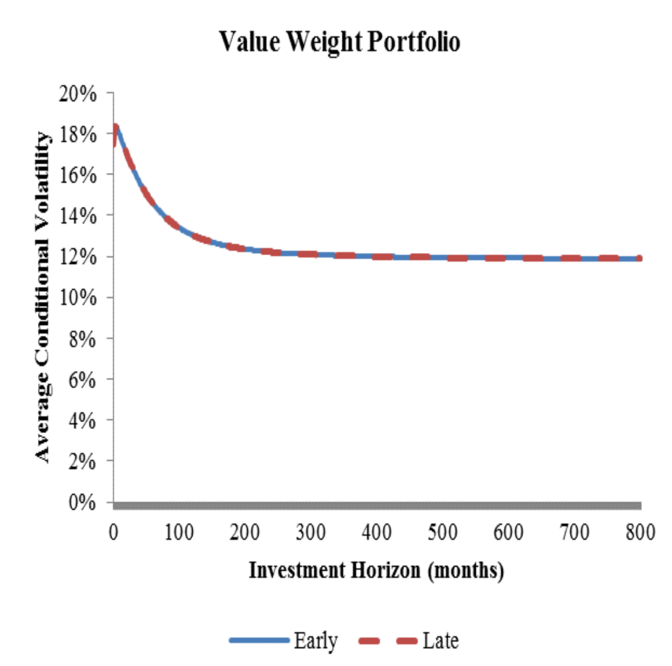
Figure 7: All-equity portfolio risk as a function of investment horizon, controlling for within-country volatility

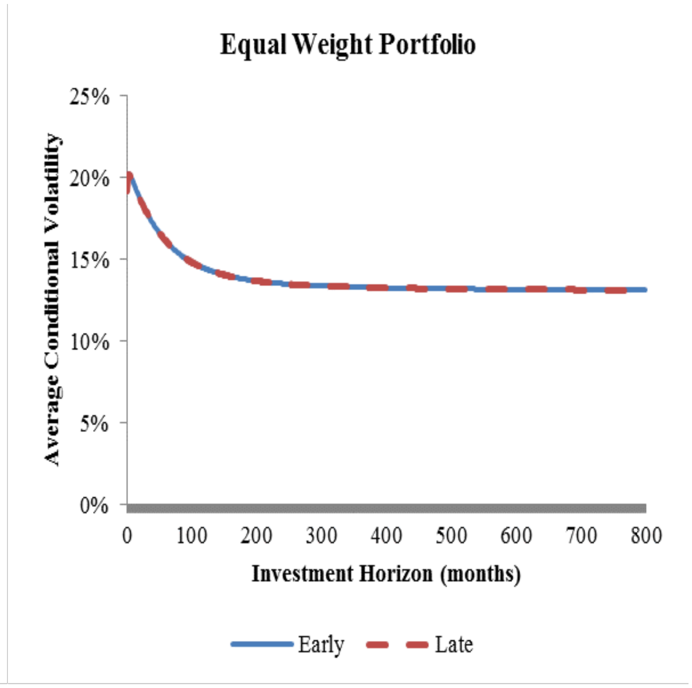
The figure compares the early sample (1986.01-1999.12) and late sample (2000.01-2013.12) all-equity portfolio risk across investment horizons. It differs from Figure 6 in that it estimates a hypothetical late sample covariance matrix to study the effect of an increase in cross-country return correlations controlling for within-country effects. The estimation imposes two constraints: (a) all state variables have the same volatility in the early and late samples; and (b) within-country correlations are the same in the early and late samples. Given these two constraints, we estimate the cross-country correlations in the late sample covariance matrix, by minimizing the distance between late sample hypothetical covariance matrix and actual covariance matrix. To guarantee the estimated hypothetical covariance matrix is well behaved, we use the semidefinite programming (SDP) methodology. Panel A plots the annualized conditional standard deviation of portfolio excess returns. Panel B plots the average (across N countries) annualized k-period conditional volatility of excess returns. Panel C plots the pairwise average of cross-country k-period excess returns conditional correlation. Each panel includes the results for value weighted and equal weighted portfolios.

Panel A: Annualized conditional standard deviation of portfolio excess returns

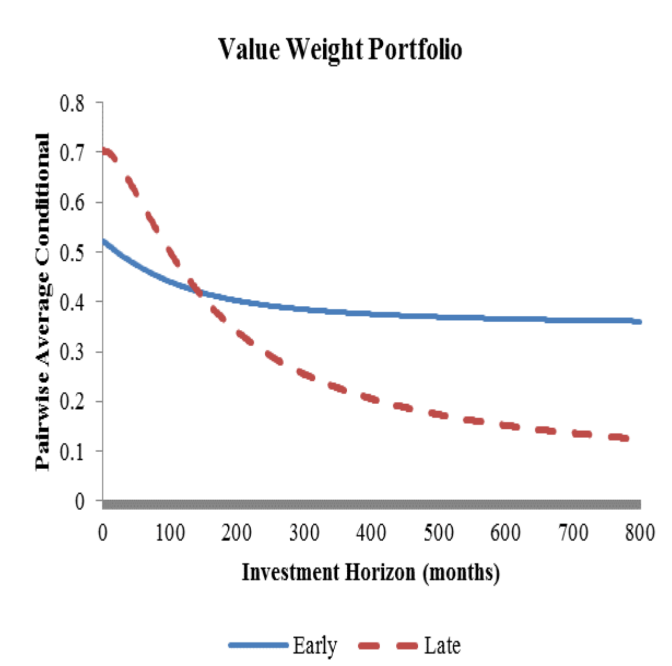


Panel B: Average annualized k-period conditional volatility of excess returns





Panel C: Pairwise average of cross-country k-period excess returns conditional correlation



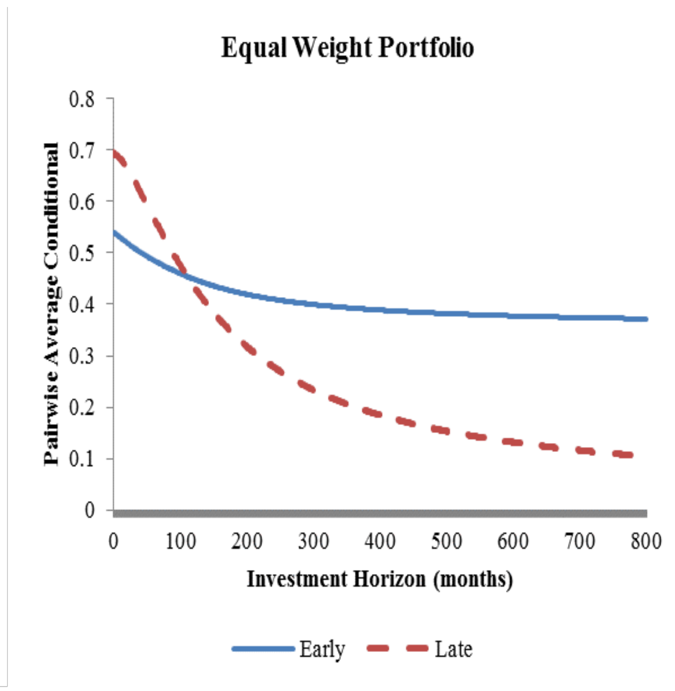
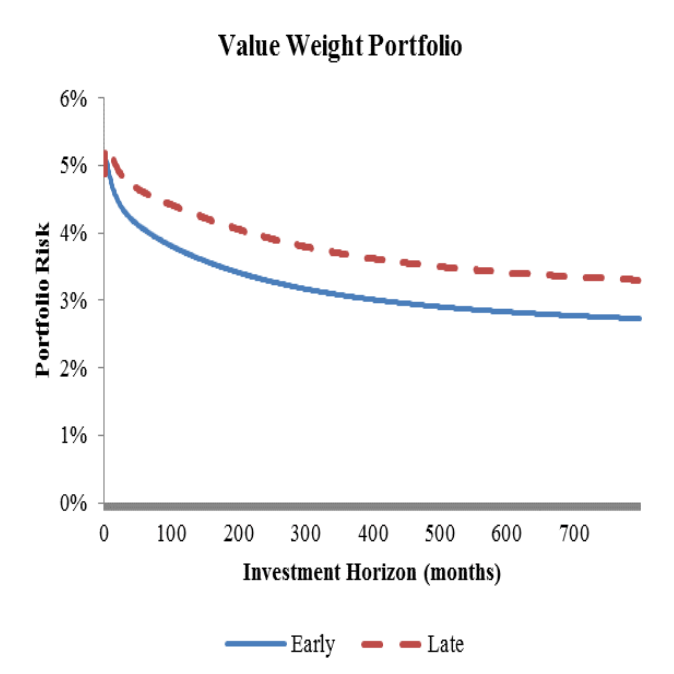
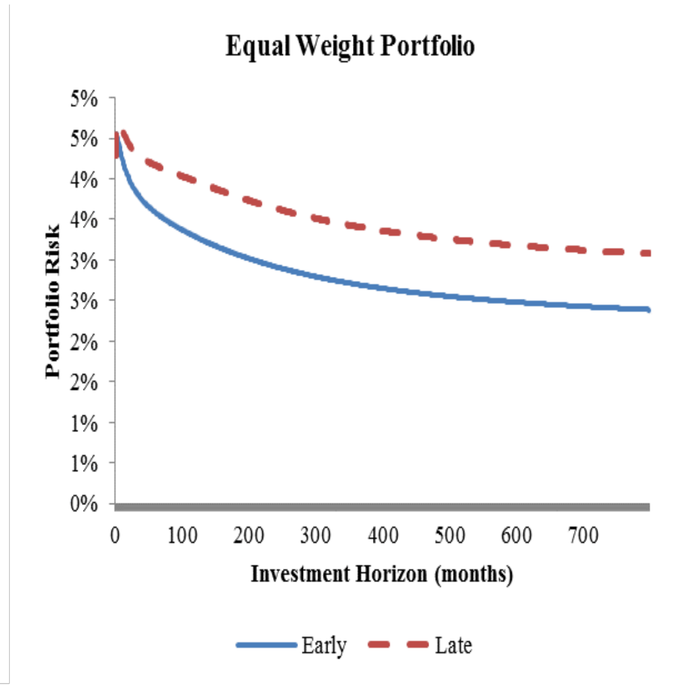


Figure 8: Bond portfolio risk as a function of investment horizon

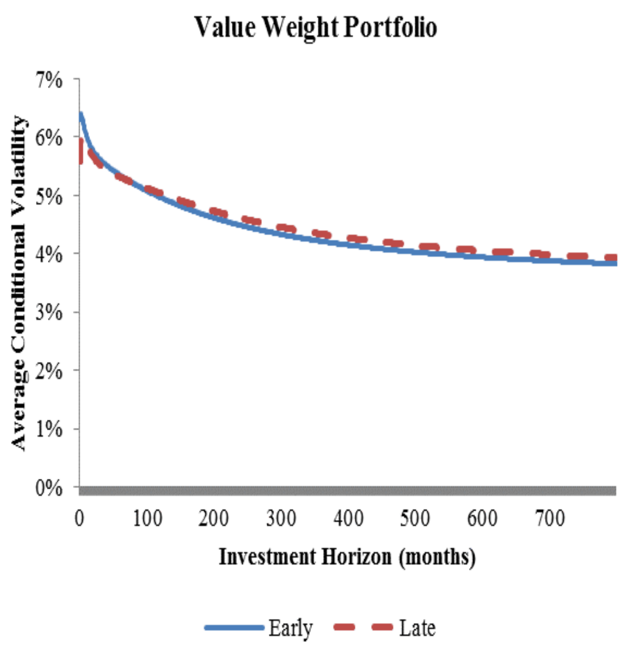
The figure compares the early sample (1986.01-1999.12) and late sample (2000.01-2013.12) all-bond portfolio risk across investment horizons. Panel A plots the annualized conditional standard deviation of portfolio excess returns. Panel B plots the average (across N countries) annualized k-period conditional volatility of excess returns. Panel C plots the pairwise average of cross-country k-period excess returns conditional correlation. Each panel includes the results for value weighted and equal weighted portfolios.

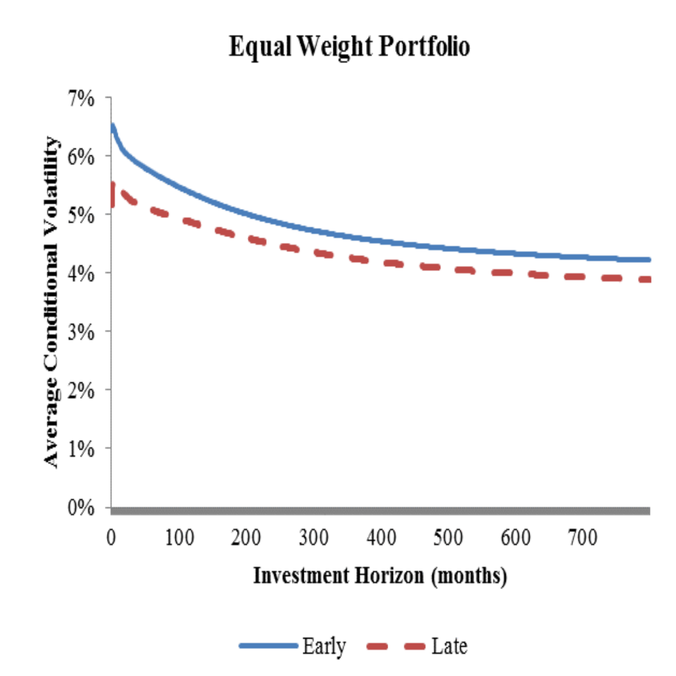
Panel A: annualized conditional standard deviation of portfolio excess returns



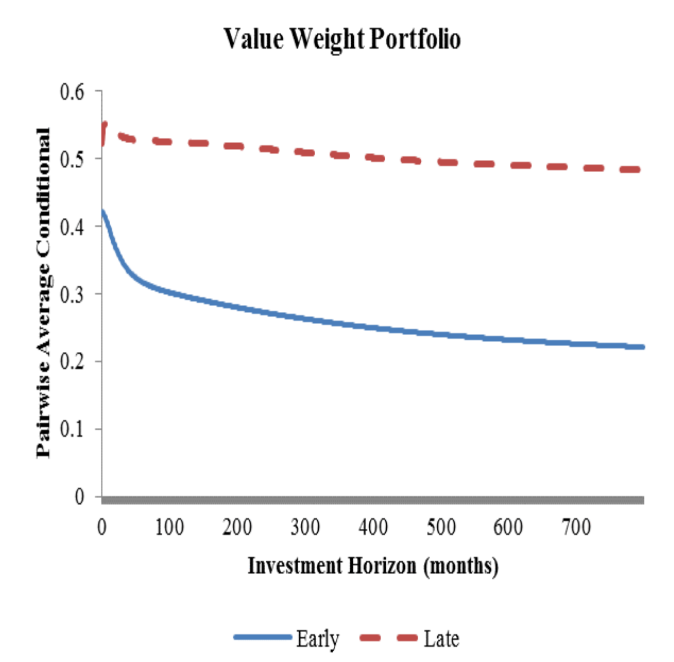


Panel B: Average annualized k-period conditional volatility of excess returns





Panel C: Pairwise average of cross-country k-period excess returns conditional correlation



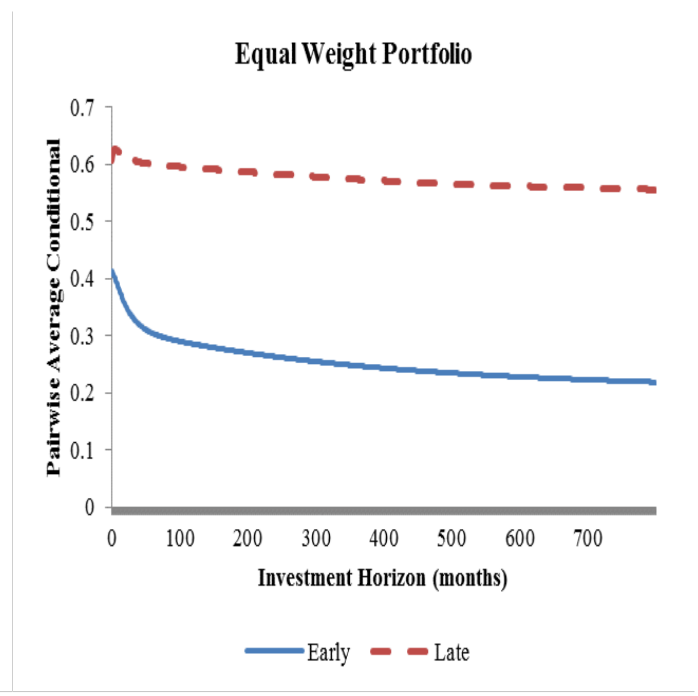


Table 1: **Summary Statistics**

Whole Sample: January 1986 to December 2013							
Stocks							
	AUS	CAN	FRA	GER	JPN	UKI	USA
Mean	5.8%	3.8%	4.0%	2.4%	-2.0%	5.2%	5.9%
Volatility	17.7%	15.8%	20.0%	22.2%	20.4%	16.1%	15.7%
Sharpe Ratio	0.32	0.24	0.20	0.11	-0.10	0.32	0.37
Bonds							
	AUS	CAN	FRA	GER	JPN	UKI	USA
Mean	5.8%	3.9%	3.9%	2.5%	0.7%	4.4%	3.6%
Volatility	6.8%	6.3%	5.6%	5.2%	5.5%	6.5%	6.4%
Sharpe Ratio	0.85	0.63	0.71	0.48	0.13	0.67	0.57
Early Sample: January 1986 to December 1999							
Stocks							
	AUS	CAN	FRA	GER	JPN	UKI	USA
Mean	6.0%	4.2%	9.0%	4.3%	-1.3%	8.7%	10.5%
Volatility	21.2%	15.5%	21.1%	21.5%	21.9%	17.4%	15.3%
Sharpe Ratio	0.28	0.27	0.43	0.20	-0.06	0.50	0.69
Bonds							
	AUS	CAN	FRA	GER	JPN	UKI	USA
Mean	7.1%	4.2%	4.3%	1.6%	1.6%	5.2%	3.2%
Volatility	7.6%	7.1%	5.9%	5.2%	6.8%	7.5%	6.4%
Sharpe Ratio	0.93	0.59	0.74	0.30	0.23	0.70	0.50
Late Sample: January 2000 to December 2013							
Stocks							
	AUS	CAN	FRA	GER	JPN	UKI	USA
Mean	5.5%	3.5%	-1.0%	0.5%	-2.7%	1.6%	1.3%
Volatility	13.6%	16.1%	18.7%	23.0%	18.8%	14.7%	16.0%
Sharpe Ratio	0.40	0.21	-0.05	0.02	-0.14	0.11	0.08
Bonds							
	AUS	CAN	FRA	GER	JPN	UKI	USA
Mean	4.5%	3.7%	3.6%	3.4%	-0.1%	3.5%	4.1%
Volatility	5.9%	5.2%	5.3%	5.3%	3.8%	5.4%	6.5%
Sharpe Ratio	0.75	0.70	0.67	0.64	-0.03	0.65	0.63

This table reports summary statistics of monthly bond and stock returns for the whole sample (January 1986 to December 2013), early sample (January 1986 to December 1999) and late sample (January 2000 to December 2013). Estimates of means, volatilities, and Sharpe Ratios are all scaled to annualized units. Returns are in U.S. Dollar currency-hedged terms in excess of the three-month U.S. Treasury bill rate.

Table 2: **Return Correlations**

Table 2.A - Correlations (Jan. 1986 - Dec. 2013)

		Bonds							Stocks						
		AUS	CAN	FRA	GER	JPN	UKI	USA	AUS	CAN	FRA	GER	JPN	UKI	USA
Bonds	AUS	1.00													
	CAN	0.55	1.00												
	FRA	0.45	0.52	1.00											
	GER	0.46	0.58	0.86	1.00										
	JPN	0.25	0.35	0.34	0.41	1.00									
	UKI	0.46	0.60	0.67	0.72	0.37	1.00								
	USA	0.54	0.73	0.60	0.64	0.34	0.59	1.00							
Stocks	AUS	0.22	-0.03	-0.05	-0.10	-0.09	-0.04	-0.16	1.00						
	CAN	0.08	0.12	-0.07	-0.11	-0.03	-0.03	-0.07	0.63	1.00					
	FRA	-0.02	0.00	0.11	-0.01	0.05	0.06	-0.14	0.56	0.64	1.00				
	GER	-0.04	-0.04	-0.03	-0.08	-0.03	-0.05	-0.20	0.55	0.61	0.83	1.00			
	JPN	-0.06	0.04	0.01	-0.05	0.03	-0.02	-0.11	0.44	0.47	0.51	0.44	1.00		
	UKI	0.12	0.09	0.04	-0.03	0.02	0.16	-0.07	0.66	0.69	0.74	0.69	0.46	1.00	
	USA	0.04	0.11	-0.02	-0.10	0.02	0.01	-0.03	0.62	0.79	0.71	0.70	0.48	0.79	1.00

Table 2.B - Correlations (Jan. 1986 - Dec. 1999)

		Bonds							Stocks						
		AUS	CAN	FRA	GER	JPN	UKI	USA	AUS	CAN	FRA	GER	JPN	UKI	USA
Bonds	AUS	1.00													
	CAN	0.44	1.00												
	FRA	0.33	0.39	1.00											
	GER	0.31	0.46	0.78	1.00										
	JPN	0.17	0.34	0.32	0.43	1.00									
	UKI	0.32	0.51	0.58	0.63	0.36	1.00								
	USA	0.41	0.68	0.50	0.52	0.38	0.47	1.00							
Stocks	AUS	0.44	0.01	0.02	-0.01	-0.12	0.04	-0.11	1.00						
	CAN	0.38	0.30	0.07	0.06	0.02	0.12	0.11	0.64	1.00					
	FRA	0.18	0.12	0.40	0.30	0.09	0.27	0.10	0.48	0.56	1.00				
	GER	0.24	0.13	0.23	0.24	-0.03	0.12	0.06	0.50	0.55	0.75	1.00			
	JPN	0.09	0.17	0.13	0.10	0.13	0.11	0.06	0.34	0.39	0.42	0.32	1.00		
	UKI	0.37	0.20	0.21	0.17	0.03	0.36	0.11	0.64	0.66	0.62	0.58	0.37	1.00	
	USA	0.34	0.35	0.19	0.12	0.05	0.21	0.26	0.57	0.79	0.59	0.55	0.36	0.73	1.00

Table 2.C - Correlations (Jan. 2000 - Dec. 2013)

		Bonds							Stocks						
		AUS	CAN	FRA	GER	JPN	UKI	USA	AUS	CAN	FRA	GER	JPN	UKI	USA
Bonds	AUS	1.00													
	CAN	0.74	1.00												
	FRA	0.62	0.72	1.00											
	GER	0.67	0.76	0.95	1.00										
	JPN	0.41	0.40	0.40	0.43	1.00									
	UKI	0.70	0.78	0.81	0.87	0.40	1.00								
	USA	0.72	0.82	0.70	0.75	0.32	0.77	1.00							
Stocks	AUS	-0.22	-0.12	-0.18	-0.24	-0.01	-0.21	-0.26	1.00						
	CAN	-0.29	-0.11	-0.21	-0.26	-0.12	-0.21	-0.24	0.68	1.00					
	FRA	-0.33	-0.20	-0.25	-0.33	-0.05	-0.26	-0.40	0.72	0.73	1.00				
	GER	-0.37	-0.25	-0.30	-0.36	-0.05	-0.28	-0.43	0.66	0.67	0.93	1.00			
	JPN	-0.29	-0.17	-0.14	-0.21	-0.18	-0.23	-0.30	0.63	0.57	0.62	0.57	1.00		
	UKI	-0.26	-0.09	-0.18	-0.24	-0.02	-0.18	-0.28	0.73	0.73	0.88	0.81	0.60	1.00	
	USA	-0.32	-0.19	-0.24	-0.30	-0.05	-0.25	-0.30	0.74	0.80	0.85	0.83	0.61	0.87	1.00

This table reports sample correlations of monthly bond and stock returns for the whole sample (January 1986 to December 2013), early sample (January 1986 to December 1999) and late sample (January 2000 to December 2013). Returns are in U.S. Dollar currency-hedged terms in excess of the three-month U.S. Treasury bill rate.

Table 3: **Correlation Summary Statistics**

		Bonds	Stocks		Bonds	Stocks
Full Period	Bonds	1.00		Bonds	0.52	
	Stocks	0.07	1.00	Stocks	-0.03	0.62
		Bonds	Stocks		Bonds	Stocks
Early Sample	Bonds	1.00		Bonds	0.44	
	Stocks	0.30	1.00	Stocks	0.13	0.54
		Bonds	Stocks		Bonds	Stocks
Late Sample	Bonds	1.00		Bonds	0.65	
	Stocks	-0.23	1.00	Stocks	-0.22	0.72
		Bonds	Stocks		Bonds	Stocks
Difference	Bonds	0.00		Bonds	0.21	
	Stocks	-0.53	0.00	Stocks	-0.35	0.18

This table summarizes the individual country-pair correlations found in Tables 2.A, 2.B, and 2.C. Overall average correlations are computed within and across countries for the full period as well as for each subperiod.

Table 4: **Return Correlation Decomposition (Stocks Across Countries)**

		Component Correlations			Component Contributions			
		CF (s)	RR (s)	ER (s)		CF (s)	RR (s)	ER (s)
Subperiod 1	CF (s)	0.20			CF (s)	0.10	0.02	0.10
	RR (s)	-0.05	0.33		RR (s)	0.02	0.04	-0.10
	ER (s)	-0.16	-0.26	0.39	ER (s)	0.10	-0.10	0.35
Subperiod 2	CF (s)	0.24			CF (s)	0.09	-0.02	0.16
	RR (s)	0.13	0.59		RR (s)	-0.02	0.07	-0.21
	ER (s)	-0.23	-0.55	0.59	ER (s)	0.16	-0.21	0.68
Difference	CF (s)	0.04			CF (s)	-0.01	-0.04	0.07
	RR (s)	0.18	0.26		RR (s)	-0.04	0.03	-0.11
	ER (s)	-0.07	-0.29	0.21	ER (s)	0.07	-0.11	0.33
		CF (s)	RR (s)	ER (s)				
p-values	CF (s)	0.18						
(bootstrap)	RR (s)	0.00	0.00					
	ER (s)	0.11	0.00	0.01				
p-values	CF (s)	0.34						
(Fisher r-to-z)	RR (s)	0.05	0.00					
	ER (s)	0.26	0.00	0.01				

This table decomposes the sources of global stock return correlations. Correlations among individual stock return components (i.e., cash-flow, real-rate, and expected-return news) across countries are shown in the leftmost columns. Contributions of these components to unexpected stock return correlations across countries (using stock return innovations from the estimated VARs) are broken down in the rightmost columns. Note that values in the total column should approximately equal their corresponding correlations found in Table 3. Estimates are reported for each subperiod as well as the difference between the two subperiods. Tests for significant correlation differences between subperiods are based on bootstrap and Fisher r-to-z methods for calculating p-values.

Table 5: **Return Correlation Decomposition (Bonds Across Countries)**

		Component Correlations			Component Contributions			
		CF (b)	RR (b)	ER (b)		CF (b)	RR (b)	ER (b)
Subperiod 1	CF (b)	0.29			CF (b)	0.02	0.06	0.00
	RR (b)	0.30	0.30		RR (b)	0.06	0.13	0.01
	ER (b)	-0.03	0.01	0.21	ER (b)	0.00	0.01	0.12
Subperiod 2	CF (b)	0.50			CF (b)	0.06	0.12	-0.04
	RR (b)	0.52	0.57		RR (b)	0.12	0.26	-0.04
	ER (b)	-0.12	-0.08	0.37	ER (b)	-0.04	-0.04	0.20
Difference	CF (b)	0.21			CF (b)	0.03	0.06	-0.03
	RR (b)	0.23	0.26		RR (b)	0.06	0.12	-0.06
	ER (b)	-0.09	-0.09	0.15	ER (b)	-0.03	-0.06	0.08
		CF (b)	RR (b)	ER (b)				
p-values	CF (b)	0.00						
(bootstrap)	RR (b)	0.00	0.00					
	ER (b)	0.13	0.11	0.01				
p-values	CF (b)	0.01						
(Fisher r-to-z)	RR (b)	0.01	0.00					
	ER (b)	0.21	0.21	0.06				

This table decomposes the sources of global bond return correlations. Correlations among individual bond return components (i.e., cash-flow, real-rate, and expected-return news) across countries are shown in the leftmost columns. Contributions of these components to unexpected bond return correlations across countries (using bond return innovations from the estimated VARs) are broken down in the rightmost columns. Note that values in the total column should approximately equal their corresponding correlations found in Table 3. Estimates are reported for each subperiod as well as the difference between the two subperiods. Tests for significant correlation differences between subperiods are based on bootstrap and Fisher r-to-z methods for calculating p-values.

Table 6: **Return Correlation Decomposition (Bonds vs. Stocks Within Countries)**

		Component Correlations			Component Contributions			
		CF(s)	RR(s)	ER(s)		CF(s)	RR(s)	ER(s)
Subperiod 1	CF (b)	-0.08	-0.21	-0.39	CF (b)	0.03	0.09	-0.18
	RR (b)	0.89	0.98	-0.22	RR (b)	0.14	0.24	-0.45
	ER (b)	-0.58	-0.61	0.46	ER (b)	0.19	-0.07	0.32
Subperiod 2	CF (b)	0.14	0.01	-0.38	CF (b)	-0.03	0.10	-0.28
	RR (b)	0.87	0.99	-0.18	RR (b)	-0.01	0.24	-0.63
	ER (b)	-0.78	-0.85	0.37	ER (b)	0.17	-0.08	0.32
Difference	CF (b)	0.22	0.22	0.01	CF (b)	-0.06	0.00	-0.10
	RR (b)	-0.02	0.01	0.03	RR (b)	-0.15	0.01	-0.18
	ER (b)	-0.19	-0.24	-0.09	ER (b)	-0.02	-0.01	0.00
		CF(s)	RR(s)	ER(s)				
p-values	CF (b)	0.00	0.00	0.48				
(bootstrap)	RR (b)	0.17	0.00	0.32				
	ER (b)	0.00	0.00	0.13				
p-values	CF (b)	0.02	0.02	0.47				
(Fisher r-to-z)	RR (b)	0.22	0.01	0.37				
	ER (b)	0.00	0.00	0.16				

This table decomposes the sources of global bond-stock return correlations within countries. Correlations among individual return components (i.e., cash-flow, real-rate, and expected-return news) within countries are shown in the leftmost columns. Contributions of these components to unexpected bond-stock return correlations within countries (using return innovations from the estimated VARs) are broken down in the rightmost columns. Note that values in the total column should approximately equal their corresponding correlations found in Table 3. Estimates are reported for each subperiod as well as the difference between the two subperiods. Tests for significant correlation differences between subperiods are based on bootstrap and Fisher r-to-z methods for calculating p-values.

Table 7: **Return Correlation Decomposition (Bonds vs. Stocks Across Countries)**

		Component Correlations			Component Contributions			
		CF(s)	RR(s)	ER(s)		CF(s)	RR(s)	ER(s)
Subperiod 1	CF (b)	-0.03	-0.05	-0.19	CF (b)	0.01	0.03	-0.08
	RR (b)	0.31	0.32	0.01	RR (b)	0.03	0.08	-0.17
	ER (b)	-0.26	-0.24	0.17	ER (b)	0.11	0.01	0.10
Subperiod 2	CF (b)	0.15	0.14	-0.28	CF (b)	-0.03	0.06	-0.19
	RR (b)	0.53	0.58	-0.07	RR (b)	-0.05	0.14	-0.41
	ER (b)	-0.51	-0.54	0.21	ER (b)	0.13	-0.02	0.17
Difference	CF (b)	0.18	0.19	-0.09	CF (b)	-0.04	0.03	-0.11
	RR (b)	0.22	0.26	-0.08	RR (b)	-0.08	0.06	-0.23
	ER (b)	-0.26	-0.30	0.04	ER (b)	0.01	-0.03	0.07
		CF(s)	RR(s)	ER(s)				
p-values	CF (b)	0.00	0.00	0.03				
(bootstrap)	RR (b)	0.00	0.00	0.13				
	ER (b)	0.00	0.00	0.34				
p-values	CF (b)	0.05	0.04	0.20				
(Fisher r-to-z)	RR (b)	0.01	0.00	0.22				
	ER (b)	0.00	0.00	0.35				

This table decomposes the sources of global bond-stock return correlations across countries. Correlations among individual return components (i.e., cash-flow, real-rate, and expected-return news) across countries are shown in the leftmost columns. Contributions of these components to unexpected bond-stock return correlations across countries (using return innovations from the estimated VARs) are broken down in the rightmost columns. Note that values in the total column should approximately equal their corresponding correlations found in Table 3. Estimates are reported for each subperiod as well as the difference between the two subperiods. Tests for significant correlation differences between subperiods are based on bootstrap and Fisher r-to-z methods for calculating p-values.

Table 8: **Average R^2 using PCs as global factors**

Panel A: Stocks				
	All	Sub-period 1	Sub-period 2	Difference
Currency Hedged Stock Returns	0.67	0.60	0.74	0.14
Unexpected Stock Returns	0.67	0.58	0.76	0.18
CF News (Stocks)	0.39	0.36	0.43	0.06
RR News (Stocks)	0.55	0.42	0.67	0.25
RP News (Stocks)	0.55	0.44	0.66	0.21

Panel B: Bonds				
	All	Sub-period 1	Sub-period 2	Difference
Currency Hedged Bond Returns	0.67	0.59	0.75	0.16
Unexpected Bond Returns	0.62	0.55	0.70	0.15
CF News (Bonds)	0.45	0.31	0.59	0.28
RR News (Bonds)	0.54	0.42	0.65	0.23
RP News (Bonds)	0.41	0.38	0.45	0.07

This table applies the Pukthuanthong-Roll methodology to realized returns, unexpected returns and the three news components of returns. For a given return or news component series, we find the first three principal components every year and obtain the R^2 from a simple least squares regression using PCs as global factors. The table reports average R^2. Panel A corresponds to stocks, and Panel B corresponds to bonds.

Table 9: **Optimal global equity portfolio allocations**

		Value Weight Portfolio			Equal Weight Portfolio		
	Country	Myopic demand	JV hedging demand at 20 yr	CCV hedging demand	Myopic demand	JV hedging demand at 20 yr	CCV hedging demand
Panel A: Early Sample	AUS	1.51%	12.82%	6.38%	14.29%	21.87%	12.66%
	CAN	2.83%	11.49%	7.65%	14.29%	15.98%	10.91%
	FRA	5.22%	-5.57%	-3.19%	14.29%	-5.11%	-3.26%
	GER	5.07%	19.46%	12.09%	14.29%	27.37%	16.95%
	JPN	16.09%	17.07%	11.20%	14.29%	17.36%	11.25%
	UKI	10.38%	-0.35%	0.42%	14.29%	-1.15%	0.72%
	USA	58.88%	48.96%	32.21%	14.29%	26.38%	16.94%
	Total	100.00%	103.89%	66.74%	100.00%	102.71%	66.16%
Panel B: Late Sample	AUS	31.54%	33.61%	24.94%	93.35%	80.30%	58.49%
	CAN	2.59%	18.20%	10.19%	15.19%	33.63%	21.54%
	FRA	4.74%	27.21%	19.55%	22.12%	32.75%	23.93%
	GER	-36.15%	-17.83%	-15.02%	-8.99%	-3.48%	-3.68%
	JPN	3.91%	9.80%	3.81%	-5.46%	5.79%	0.14%
	UKI	49.40%	10.75%	7.60%	52.94%	9.31%	5.57%
	USA	59.83%	57.76%	38.73%	-43.17%	-0.66%	-4.08%
	Total	115.85%	139.50%	89.81%	125.97%	157.64%	101.90%
Panel C: Late Sample (Hypothetical Covariance Matrix)	AUS	-9.02%	6.58%	3.81%	11.52%	5.87%	3.79%
	CAN	25.22%	5.05%	3.87%	28.37%	3.66%	2.73%
	FRA	-1.72%	16.50%	9.02%	14.67%	15.23%	8.58%
	GER	-20.46%	22.30%	11.11%	7.28%	23.30%	12.69%
	JPN	1.34%	10.52%	6.14%	-0.35%	8.99%	5.03%
	UKI	18.80%	0.45%	-0.34%	16.98%	0.44%	-0.41%
	USA	82.46%	44.33%	32.04%	6.18%	59.10%	39.18%
	Total	96.63%	105.74%	65.65%	84.64%	116.59%	71.59%

The table reports optimal global equity portfolio allocations by "JV" investor and "CCV" investor. The myopic demand is the allocation of those two investors at investment horizon 1. An investor's allocation is the sum of myopic demand and hedging demand. We report the JV hedging demand for an investor at horizon of 20 years (240 months). We compare across 3 scenarios: optimal allocation in early sample (Panel A), late sample (Panel B) and late sample with hypothetical covariance matrix that controls for within-country correlation (Panel C). To make it comparable, we fix the monthly implied excess returns across these 3 scenarios. We set implied excess returns for value weight portfolio such that investor hold the myopic demand equal to market cap weight, and for equal weight portfolio such that investor hold the myopic demand equal to 1/N in each country. "Total" allocation is the sum of the allocations to each country.

Table 10: **Expected Utility**

		Value Weight Portfolio			Equal Weight Portfolio		
	Number of Countries	JV W_{CE} K=120	JV W_{CE} K=240	CCV $E[V_t]$	JV W_{CE} K=120	JV W_{CE} K=240	CCV $E[V_t]$
Panel A: Early Sample	7	82.08	11861.95	0.078	108.62	11668.84	0.083
	1	4.00	34.52	0.011	3.24	22.48	0.009
Panel B: Late Sample	7	111.84	2380750.84	0.171	160.45	3272444.57	0.222
	1	3.88	37.30	0.013	3.08	22.93	0.010
Panel C: Late Sample (Hypothetical Covaricnce Matrix)	7	79.92	11291.24	0.076	116.33	15737.17	0.079
	1	4.00	34.52	0.011	3.24	22.48	0.009

The table reports the expected utility by "JV" investor and "CCV" investor, with the same optimal portfolio allocation as reported in Table 9 (across the 3 scenarios). The CCV investor has Epstein-Zin preference and the expected value function defined as $E[V_t] \equiv \frac{U_t}{W_t} = (1-\delta)^{-\psi/(1-\psi)} \left(\frac{C_t}{W_t}\right)^{1/(1-\psi)}$. We report the expected value function for CCV investor in the Table across the 3 scenarios (with EIS $\psi \to 1$ and RRA$\gamma = 5$). The JV investor's utility is power utility defined on terminal wealth $E_t[\frac{1}{1-\gamma} W_{t+K}^{1-\gamma}]$. We assume investor has initial wealth of one dollar and look at investment horizons of 10 years (120 months) and 20 years (240 months). We report the certainty equivalent for the JV investor (with RRA$\gamma = 5$). The results are obtained by Monte Carlo simulation using 2,000 VAR paths sampled using the method of antithetic variates. The certainty equivalent of wealth is computed by evaluating the mean utility realized across the simulated paths and computing, $W_{CE} = u^{-1}\left(E[u(\widetilde{W_{t+K}})]\right)$.

Appendix

Appendix A. Calculating News Components

A.1 Excess Bond Returns (3 News Components)

Define the log one-period nominal return on a nominal n-period coupon bond as

$$\begin{aligned} r^{\$}_{n,t+1} &= \log\left(1 + R^{\$}_{n,t+1}\right) = \log\left(P_{n-1,t+1} + C\right) - \log\left(P_{n,t}\right) \\ &= p_{n-1,t+1} - p_{n,t} + \log\left(1 + \exp\left(c - p_{n-1,t+1}\right)\right) \\ &\approx k + \rho_b p_{n-1,t+1} + \left(1 - \rho_b\right) c - p_{n,t}, \end{aligned} \tag{1}$$

where $\rho_b = \frac{1}{1+\exp\left(\overline{c-p}\right)}$ and $k = -\log\left(\rho_b\right) - \left(1 - \rho_b\right)\log\left(\frac{1}{\rho_b} - 1\right)$. Solving forward and imposing the terminal condition that $p_{n-j,t+j}|_{j=n} = 0$, we get that

$$p_{n,t} = \left(k + \left(1 - \rho_b\right) c\right)\left(\sum_{j=0}^{n-1} \rho_b^j\right) - \sum_{j=0}^{n-1} r^{\$}_{n-j,t+1+j}\rho_b^j.$$

Plugging this expression in to the unexpected bond return from Eq. (1), we get that

$$\begin{aligned} \left(\mathbb{E}_{t+1} - \mathbb{E}_t\right)\left[r^{\$}_{n,t+1}\right] &= \left(\mathbb{E}_{t+1} - \mathbb{E}_t\right)\left[\rho_b p_{n-1,t+1}\right] - \left(\mathbb{E}_{t+1} - \mathbb{E}_t\right)\left[p_{n,t}\right] \\ &= \left(\mathbb{E}_{t+1} - \mathbb{E}_t\right)\left[\rho_b p_{n-1,t+1}\right] \\ &= -\left(\mathbb{E}_{t+1} - \mathbb{E}_t\right)\left[\sum_{j=1}^{n-1} r^{\$}_{n-j,t+1+j}\rho_b^j\right]. \end{aligned} \tag{2}$$

We can write $r^{\$}_{n,t+1} = xr_{n,t+1} + r^{\$}_{f,t+1}$, where $xr_{n,t+1}$ is the excess log 1-period return on a nominal n-period coupon bond and $r^{\$}_{f,t+1}$ is the realized nominal return of the 1-period nominal bond, which is the same as the yield of the 1-period nominal bond $y_{1,t}$.

Decomposing the surprise bond return in Eq. (2) gives

$$\left(\mathbb{E}_{t+1} - \mathbb{E}_t\right)\left[xr_{n,t+1} + r^{\$}_{f,t+1}\right] = -\left(\mathbb{E}_{t+1} - \mathbb{E}_t\right)\left[\sum_{j=1}^{n-1} \rho_b^j xr_{n-j,t+1+j}\right] - \left(\mathbb{E}_{t+1} - \mathbb{E}_t\right)\left[\sum_{j=1}^{n-1} \rho_b^j r^{\$}_{f,t+1+j}\right].$$

The LHS can be simplified by noting that

$$\left(\mathbb{E}_{t+1} - \mathbb{E}_t\right)\left[r^{\$}_{f,t+1}\right] = \left(\mathbb{E}_{t+1} - \mathbb{E}_t\right)\left[y_{1,t}\right] = 0.$$

To simplify the RHS, we simply note that the realized nominal return of the 1-period nominal bond is the realized real return of the 1-period nominal bond plus realized inflation: $r^{\$}_{f,t+1} = r_{f,t+1} + \pi_{t+1}$. The second term on the RHS is then

$$\left(\mathbb{E}_{t+1} - \mathbb{E}_t\right)\left[\sum_{j=1}^{n-1} \rho_b^j r^{\$}_{f,t+1+j}\right] = \left(\mathbb{E}_{t+1} - \mathbb{E}_t\right)\left[\sum_{j=1}^{n-1} \rho_b^j r_{f,t+1+j}\right] + \left(\mathbb{E}_{t+1} - \mathbb{E}_t\right)\left[\sum_{j=1}^{n-1} \rho_b^j \pi_{t+1+j}\right]. \tag{3}$$

Putting together the simplified LHS and RHS, we have the following 3 news component decomposition for unexpected excess bond returns:

$$r_{n,t+1} - \mathbb{E}_t\left[r_{n,t+1}\right] = \left(\mathbb{E}_{t+1} - \mathbb{E}_t\right)\left[xr_{n,t+1}\right] = N_{CF,n,t+1} - N_{RR,n,t+1} - N_{RP,n,t+1}$$

where

$$N_{CF,n,t+1} = -N_{INFL,n,t+1} \equiv -\left(\mathbb{E}_{t+1} - \mathbb{E}_t\right)\left[\sum_{j=1}^{n-1} \rho_b^j \pi_{t+1+j}\right],$$

$$N_{RR,n,t+1} \equiv \left(\mathbb{E}_{t+1} - \mathbb{E}_t\right)\left[\sum_{j=1}^{n-1} \rho_b^j r_{f,t+1+j}\right], \text{ and}$$

$$N_{RP,n,t+1} \equiv \left(\mathbb{E}_{t+1} - \mathbb{E}_t\right)\left[\sum_{j=1}^{n-1} \rho_b^j xr_{n-j,t+1+j}\right]. \tag{4}$$

To extract the news components from the VAR, consider the vector of state variables

$$\tilde{\mathbf{z}}_{t+1} = \left[xr_{s,t+1}, xr_{n,t+1}, d_{t+1} - p_{t+1}, \pi_{t+1}, y_{1,t+1}, y_{10,t+1} - y_{1,t+1}\right]. \tag{5}$$

The main VAR equation is $\tilde{\mathbf{z}}_{t+1} = a + \mathbf{A}\tilde{\mathbf{z}}_t + \mathbf{u}_{t+1}$, which leads to $\mathbb{E}_t\left[\tilde{\mathbf{z}}_{t+j}\right] = \mathbf{A}^j \tilde{\mathbf{z}}_t$ and $\left(\mathbb{E}_{t+1} - \mathbb{E}_t\right)\left[\tilde{\mathbf{z}}_{t+j}\right] = \mathbf{A}^{j-1}\mathbf{u}_{t+1}$. It is then straightforward to see how the decomposition can be written in VAR notation:

$$\begin{aligned}
\left(\mathbb{E}_{t+1} - \mathbb{E}_t\right)\left[xr_{n,t+1}\right] =& \mathbf{e2}'\mathbf{u}_{t+1}, \\
N_{CF,n,t+1} =& -\mathbf{e4}'\left(\sum_{j=1}^{n-1} \rho_b^j \mathbf{A}^j\right)\mathbf{u}_{t+1}, \\
N_{RR,n,t+1} =& \mathbf{e5}'\left(\sum_{j=1}^{n-1} \rho_b^j \mathbf{A}^{j-1}\right)\mathbf{u}_{t+1} - \mathbf{e4}'\left(\sum_{j=1}^{n-1} \rho_b^j \mathbf{A}^j\right)\mathbf{u}_{t+1}, \text{ and} \\
N_{RP,n,t+1} =& N_{CF,n,t+1} - N_{RR,n,t+1} - \left(\mathbb{E}_{t+1} - \mathbb{E}_t\right)\left[xr_{n,t+1}\right].
\end{aligned}$$

We get $N_{RR,n,t+1}$ by using Eq. (3) to express real rate news in terms of nominal rate news and inflation news. Finally, we back out $N_{RP,n,t+1}$ as the residual.

A.2 Real Bond Returns (2 News Components)

With the excess bond returns decomposition in hand, the 2 news component decomposition follows.

$$\left(\mathbb{E}_{t+1} - \mathbb{E}_t\right)\left[r_{n,t+1}\right] = N_{CF,n,t+1} - N_{DR,n,t+1},$$

where

$$N_{CF,n,t+1} = -N_{INFL,n,t+1} \equiv -\left(\mathbb{E}_{t+1} - \mathbb{E}_t\right)\left[\sum_{j=0}^{n-1} \rho_b^j \pi_{t+1+j}\right] \text{ and}$$

$$N_{DR,n,t+1} \equiv \left(\mathbb{E}_{t+1} - \mathbb{E}_t\right)\left[\sum_{j=1}^{n-1} \rho_b^j r_{n-j,t+1+j}\right]. \tag{6}$$

$r_{n,t+1}$ is the realized real return of the nominal n-period coupon bond.

We can relate the 2 news component decomposition to the 3 news component decomposition as follows. The unexpected excess bond return differs from the unexpected real bond return by an inflation innovation term that we adjust for by indexing $N_{CF,n,t+1}$ from $j = 0$ instead of $j = 1$:

$$\begin{aligned}
\left(\mathbb{E}_{t+1} - \mathbb{E}_t\right)\left[xr_{n,t+1}\right] &= \left(\mathbb{E}_{t+1} - \mathbb{E}_t\right)\left[r_{n,t+1} - r_{f,t+1}\right] \\
&= \left(\mathbb{E}_{t+1} - \mathbb{E}_t\right)\left[r_{n,t+1}\right] - \left(\mathbb{E}_{t+1} - \mathbb{E}_t\right)\left[r^{\$}_{f,t+1} - \pi_{t+1}\right] \\
&= \left(\mathbb{E}_{t+1} - \mathbb{E}_t\right)\left[r_{n,t+1}\right] - \left(\mathbb{E}_{t+1} - \mathbb{E}_t\right)\left[y_{1,t} - \pi_{t+1}\right] \\
&= \left(\mathbb{E}_{t+1} - \mathbb{E}_t\right)\left[r_{n,t+1}\right] + \left(\mathbb{E}_{t+1} - \mathbb{E}_t\right)\left[\pi_{t+1}\right].
\end{aligned}$$

We can go from the penultimate line to the last line because the nominal short rate is in the information set at time t. Furthermore, it is obvious from Eqs. (4) and (6) that $N_{RR,n,t+1} + N_{RP,n,t+1} = N_{DR,n,t+1}$. Thus, keeping the same vector of state variables $\mathbf{z}_{t+1}$ as in Eq. (5), we write the decomposition in VAR notation:

$$\begin{aligned}
(\mathbb{E}_{t+1} - \mathbb{E}_t)\left[r_{n,t+1}\right] =& \mathbf{e2}'\mathbf{w}_{t+1} - \mathbf{e4}'\mathbf{u}_{t+1}, \\
N_{CF,n,t+1} =& -\mathbf{e4}'\left(\sum_{j=0}^{n-1} \rho_b^j \mathbf{A}^j\right)\mathbf{u}_{t+1} \text{ and} \\
N_{DR,s,t+1} =& N_{CF,n,t+1} - (\mathbb{E}_{t+1} - \mathbb{E}_t)\left[r_{n,t+1}\right].
\end{aligned}$$

A.3 Excess Stock Returns (3 News Components)

We know the equation for the 3 news component decomposition for unexpected excess stock returns:

$$(\mathbb{E}_{t+1} - \mathbb{E}_t)\left[xr_{s,t+1}\right] = N_{CF,s,t+1} - N_{RR,s,t+1} - N_{RP,s,t+1},$$

where

$$\begin{aligned}
N_{CF,s,t+1} &\equiv (\mathbb{E}_{t+1} - \mathbb{E}_t)\left[\sum_{j=0}^{\infty} \rho_s^j \Delta d_{t+1+j}\right], \\
N_{RR,s,t+1} &\equiv (\mathbb{E}_{t+1} - \mathbb{E}_t)\left[\sum_{j=0}^{\infty} \rho_s^j r_{f,t+1+j}\right], \text{ and} \\
N_{RP,s,t+1} &\equiv (\mathbb{E}_{t+1} - \mathbb{E}_t)\left[\sum_{j=1}^{\infty} \rho_s^j xr_{s,t+1+j}\right].
\end{aligned} \tag{7}$$

With the same vector of state variables $\mathbf{z}_{t+1}$ as in Eq. (5), we write the decomposition in VAR notation:

$$\begin{aligned}
(\mathbb{E}_{t+1} - \mathbb{E}_t)\left[xr_{s,t+1}\right] =& \mathbf{e1}'\mathbf{u}_{t+1}, \\
N_{CF,s,t+1} =& (\mathbb{E}_{t+1} - \mathbb{E}_t)\left[xr_{s,t+1}\right] + N_{RR,s,t+1} + N_{RP,s,t+1}, \\
N_{RR,s,t+1} =& \mathbf{e5}'\left(\sum_{j=1}^{\infty} \rho_s^j \mathbf{A}^{j-1}\right)\mathbf{u}_{t+1} - \mathbf{e4}'\left(\sum_{j=0}^{\infty} \rho_s^j \mathbf{A}^j\right)\mathbf{u}_{t+1}, \text{ and} \\
N_{RP,s,t+1} =& \mathbf{e1}'\left(\sum_{j=1}^{\infty} \rho_s^j \mathbf{A}^j\right)\mathbf{u}_{t+1}.
\end{aligned}$$

Similar to the case with bonds, we get $N_{RR,n,t+1}$ by using an infinite-sum version of Eq. (3) to express real rate news in terms of nominal rate news and inflation news. Note that the first term in $N_{RR,s,t+1}$ starts from $j = 1$ instead of $j = 0$ because $(\mathbb{E}_{t+1} - \mathbb{E}_t)\left[y_{1,t}\right] = 0$. Finally, we back out $N_{CF,s,t+1}$ as the residual.

A.4 Real Stock Returns (2 News Components)

The main equation here is simply

$$(\mathbb{E}_{t+1} - \mathbb{E}_t)\left[r_{s,t+1}\right] = N_{CF,s,t+1} - N_{DR,s,t+1},$$

where

$$\begin{aligned}
N_{CF,s,t+1} &\equiv (\mathbb{E}_{t+1} - \mathbb{E}_t)\left[\sum_{j=0}^{\infty} \rho_s^j \Delta d_{t+1+j}\right] \text{ and} \\
N_{DR,s,t+1} &\equiv (\mathbb{E}_{t+1} - \mathbb{E}_t)\left[\sum_{j=1}^{\infty} \rho_s^j r_{s,t+1+j}\right].
\end{aligned} \tag{8}$$

We can relate the 2 news component decomposition to the 3 news component decomposition as follows. The unexpected excess stock return differs from the unexpected real stock return by an innovation term for the realized real return of the 1-period nominal bond:

$$
\begin{aligned}
\left(\mathbb{E}_{t+1}-\mathbb{E}_{t}\right)\left[xr_{s,t+1}\right] &= \left(\mathbb{E}_{t+1}-\mathbb{E}_{t}\right)\left[r_{s,t+1}-r_{f,t+1}\right] \\
&= \left(\mathbb{E}_{t+1}-\mathbb{E}_{t}\right)\left[r_{s,t+1}\right]-\left(\mathbb{E}_{t+1}-\mathbb{E}_{t}\right)\left[r_{f,t+1}\right].
\end{aligned}
$$

By adding this extra term back to $N_{RR,s,t+1}$ and $N_{RP,s,t+1}$, we get that $-\left(\mathbb{E}_{t+1}-\mathbb{E}_{t}\right)\left[r_{f,t+1}\right]+N_{RR,s,t+1}+N_{RP,s,t+1}=N_{DR,s,t+1}$. Since we also have that

$$
\begin{aligned}
\left(\mathbb{E}_{t+1}-\mathbb{E}_{t}\right)\left[r_{f,t+1}\right] &= \left(\mathbb{E}_{t+1}-\mathbb{E}_{t}\right)\left[r_{f,t+1}^{\$}-\pi_{t+1}\right] \\
&= \left(\mathbb{E}_{t+1}-\mathbb{E}_{t}\right)\left[y_{1,t}-\pi_{t+1}\right] \\
&= -\left(\mathbb{E}_{t+1}-\mathbb{E}_{t}\right)\left[\pi_{t+1}\right]
\end{aligned}
$$

because the nominal short rate is in the information set at time t, we see that

$$
\left(\mathbb{E}_{t+1}-\mathbb{E}_{t}\right)\left[r_{s,t+1}\right]=\left(\mathbb{E}_{t+1}-\mathbb{E}_{t}\right)\left[xr_{s,t+1}\right]-\left(\mathbb{E}_{t+1}-\mathbb{E}_{t}\right)\left[\pi_{t+1}\right].
$$

Similarly,

$$
\left(\mathbb{E}_{t+1}-\mathbb{E}_{t}\right)\left[\pi_{t+1}\right]+N_{RR,s,t+1}+N_{RP,s,t+1}=N_{DR,s,t+1}.
$$

Keeping the same vector of state variables $\mathbf{z}_{t+1}$ as in Eq. (5), we write the decomposition in VAR notation:

$$
\begin{aligned}
\left(\mathbb{E}_{t+1}-\mathbb{E}_{t}\right)\left[r_{s,t+1}\right] =& \mathbf{e1}'\mathbf{u}_{t+1}-\mathbf{e4}'\mathbf{u}_{t+1}, \\
N_{CF,s,t+1} =& \left(\mathbb{E}_{t+1}-\mathbb{E}_{t}\right)\left[r_{s,t+1}\right]+N_{DR,s,t+1} \text{ and} \\
N_{DR,s,t+1} =& \mathbf{e5}'\left(\sum_{j=1}^{\infty}\rho_s^j\mathbf{A}^{j-1}\right)\mathbf{u}_{t+1}-\mathbf{e4}'\left(\sum_{j=1}^{\infty}\rho_s^j\mathbf{A}^{j}\right)\mathbf{u}_{t+1}+\mathbf{e1}'\left(\sum_{j=1}^{\infty}\rho_s^j\mathbf{A}^{j}\right)\mathbf{u}_{t+1}.
\end{aligned}
$$

Appendix B. Derivation of Results in Section 3.2

We want to derive the general formula for k period portfolio return variance, where the portfolio is constructed by holding equal weight on N identical countries. The starting point is from our toy model

$$\begin{cases} r_{i,t+1} = \mu_{i1} + \beta_i s_{i,t} + u_{i,t+1} \\ s_{i,t+1} = \mu_{i2} + \phi_i s_{i,t} + u_{s_i,t+1} \end{cases} \tag{9}$$

and we could also write the VAR residual in terms of news terms $u_{i,t+1} = N_{CF,i,t+1} - N_{DR,i,t+1}$ and $u_{s_i,t+1} = \frac{1}{\lambda_i} N_{DR,i,t+1}$, where $\lambda_i = \frac{\rho\beta_i}{1-\rho\phi_i}$. The log portfolio return is

$$r_{p,t+k}^{(k)} = r_0^{(k)} + \alpha_t'(r_{t+k}^{(k)} - r_0^{(k)} l) + \frac{1}{2}\alpha_t(k)^2\sigma_t(k)^2 - \frac{1}{2}\alpha_t(k)\Sigma_t(k)\alpha_t(k) \tag{10}$$

and the variance of k period portfolio return is

$$V_t[r_{p,t+k}^{(k)}] = \frac{1}{N} V_t[r_{i,t+k}^{(k)}] + (1 - \frac{1}{N}) C_t[r_{i,t+k}^{(k)}, r_{j,t+k}^{(k)}] \tag{11}$$

The term of interest in the expression is the cross-country covariance. Let's now derive the general expression for the covariance term. Note that the 1 period return at $t+l$ could be written as

$$r_{i,t+l} = \mu_{i1} + \beta_i s_{i,t+l-1} + u_{i,t+l}$$

$$= \mu_{i1} + \beta_i(\phi_i s_{i,t+l-2} + u_{s_i,t+l-1}) + u_{i,t+l}$$

$$\cdots$$

$$= \mu_{i1} + \beta_i \phi_i^{l-1} s_{i,t} + \beta_i \sum_{m=1}^{l-1} \phi_i^{m-1} u_{s_i,t+l-m} + u_{i,t+l} \tag{12}$$

and

$$C_t[r_{i,t+l}, r_{j,t+l}] = C_t[\beta_i \sum_{m=1}^{l-1} \phi_i^{m-1} u_{s_i,t+l-m} + u_{i,t+l}, \beta_j \sum_{m=1}^{l-1} \phi_j^{m-1} u_{s_j,t+l-m} + u_{j,t+l}]$$

$$= C_t[\frac{\beta_i}{\lambda_i} \sum_{m=1}^{l-1} \phi_i^{m-1} N_{DR,i,t+l-m} + N_{CF,i,t+l} - N_{DR,i,t+l}, \frac{\beta_j}{\lambda_j} \sum_{m=1}^{l-1} \phi_j^{m-1} N_{DR,j,t+l-m} + N_{CF,j,t+l} - N_{DR,j,t+l}] \tag{13}$$

note that this is a symmetric model, thus we could ignore the subscript i and j in the expression. We make the assumption that $(for\ \forall l \geqslant 1, i \neq j)$

$$C_t[N_{CF,i,t+l}, N_{CF,j,t+l}] \equiv \sigma_{CF,CF}^{xc}$$

$$C_t[N_{CF,i,t+l}, N_{DR,j,t+l}] \equiv \sigma_{CF,DR}^{xc}$$

$$C_t[N_{DR,i,t+l}, N_{DR,j,t+l}] \equiv \sigma_{DR,DR}^{xc}$$

Thus we have

$$C_t[r_{i,t+l}, r_{j,t+l}] = [\frac{\beta^2}{\lambda^2} \frac{(1-(\phi^2)^{l-1})}{1-\phi^2} + 1]\sigma_{DR,DR}^{xc} + \sigma_{CF,CF}^{xc} - 2\sigma_{CF,DR}^{xc} \tag{14}$$

For the cross-period & cross-country covariance, we have

$$C_t[r_{i,t+l}, r_{j,t+l+p}] = C_t[\beta_i \sum_{m=1}^{l-1} \phi_i^{m-1} u_{s_i,t+l-m} + u_{i,t+l}, \beta_j \sum_{m=1}^{l+p-1} \phi_j^{m-1} u_{s_j,t+l+p-m} + u_{j,t+l+p}]$$

$$= C_t[u_{i,t+l} + \beta_i u_{s_i,t+l-1} + \beta_i \phi_i u_{s_i,t+l-2} + \cdots + \beta_i \phi_i^{l-2} u_{s_i,t+1}, \beta_j \phi_j^{p-1} u_{s_j,t+l} + \beta_j \phi_j^p u_{s_j,t+l-1} + \beta_j \phi_j^{p+1} u_{s_j,t+l-2} + \cdots + \beta_j \phi_j^{l+p-2} u_{s_j,t+1}]$$

$$= \beta\phi^{p-1} C_t[u_{i,t+l}, u_{s_j,t+l}] + \beta^2\phi^p C_t[u_{s_i,t+l-1}, u_{s_j,t+l-1}] + \beta^2\phi^{p+2} C_t[u_{s_i,t+l-2}, u_{s_j,t+l-2}] + \cdots + \beta^2\phi^{p+2(l-2)} C_t[u_{s_i,t+1}, u_{s_j,t+1}]$$

$$= \frac{\beta\phi^{p-1}}{\lambda}(\sigma^{xc}_{CF,DR} - \sigma^{xc}_{DR,DR}) + \frac{\beta^2\phi^p}{\lambda^2}\frac{1-(\phi^2)^{l-1}}{1-\phi^2}\sigma^{xc}_{DR,DR} \tag{15}$$

with $p \geqslant 1$. Using the results above, we could get the k period cross-country return covariance

$$C_t[r^{(k)}_{i,t+k}, r^{(k)}_{j,t+k}] = \sum_{l=1}^{k} C_t[r_{i,t+l}, r_{j,t+l}] + 2\sum_{l=1}^{k-1}\sum_{p=1}^{k-l} C_t[r_{i,t+l}, r_{j,t+l+p}]$$

$$= \sum_{l=1}^{k}\left([\frac{\beta^2}{\lambda^2}\frac{(1-(\phi^2)^{l-1})}{1-\phi^2}+1]\sigma^{xc}_{DR,DR} + \sigma^{xc}_{CF,CF} - 2\sigma^{xc}_{CF,DR}\right) + 2\sum_{l=1}^{k-1}\sum_{p=1}^{k-l}\left(\frac{\beta\phi^{p-1}}{\lambda}(\sigma^{xc}_{CF,DR} - \sigma^{xc}_{DR,DR}) + \frac{\beta^2\phi^p}{\lambda^2}\frac{1-(\phi^2)^{l-1}}{1-\phi^2}\sigma^{xc}_{DR,DR}\right)$$

$$= \left([\frac{\beta^2}{\lambda^2}\frac{(k-\frac{1-(\phi^2)^k}{1-\phi^2})}{1-\phi^2}+k]\sigma^{xc}_{DR,DR} + k\sigma^{xc}_{CF,CF} - 2k\sigma^{xc}_{CF,DR}\right)$$

$$+2\sum_{l=1}^{k-1}\left(\frac{\beta}{\lambda(1-\phi)}(1-\phi^{k-l})(\sigma^{xc}_{CF,DR} - \sigma^{xc}_{DR,DR}) + \frac{\beta^2}{\lambda^2}\frac{1-(\phi^2)^{l-1}}{1-\phi^2}\frac{\phi(1-\phi^{k-l})}{1-\phi}\sigma^{xc}_{DR,DR}\right)$$

$$= \left([\frac{\beta^2}{\lambda^2}\frac{(k-\frac{1-(\phi^2)^k}{1-\phi^2})}{1-\phi^2}+k]\sigma^{xc}_{DR,DR} + k\sigma^{xc}_{CF,CF} - 2k\sigma^{xc}_{CF,DR}\right)$$

$$+2\left(\frac{\beta}{\lambda(1-\phi)}(k-1-\phi\frac{1-\phi^{k-1}}{1-\phi})(\sigma^{xc}_{CF,DR} - \sigma^{xc}_{DR,DR}) + \frac{\beta^2\phi}{\lambda^2(1-\phi^2)(1-\phi)}(k-1+\frac{(\phi^{k-1}-1)(\phi-\phi^{k-1})}{1-\phi} - \frac{1-(\phi^2)^{k-1}}{1-\phi^2})\sigma^{xc}_{DR,DR}\right)$$

$$= k\sigma^{xc}_{CF,CF} + 2k\left(\frac{\beta}{\lambda(1-\phi)}(\frac{k-1}{k} - \frac{\phi}{k}\frac{1-\phi^{k-1}}{1-\phi}) - 1\right)\sigma^{xc}_{CF,DR}$$

$$+\left(\frac{\beta^2}{\lambda^2}\frac{(k-\frac{1-(\phi^2)^k}{1-\phi^2})}{1-\phi^2} + 2\frac{\beta^2\phi}{\lambda^2(1-\phi^2)(1-\phi)}(k-1+\frac{(\phi^{k-1}-1)(\phi-\phi^{k-1})}{1-\phi} - \frac{1-(\phi^2)^{k-1}}{1-\phi^2}) - 2\frac{\beta}{\lambda(1-\phi)}(k-1-\phi\frac{1-\phi^{k-1}}{1-\phi}) + k\right)\sigma^{xc}_{DR,} \tag{16}$$

We further simplify the coefficient on $\sigma^{xc}_{DR,DR}$ as

$$\frac{\beta^2}{\lambda^2}\frac{(k-\frac{1-(\phi^2)^k}{1-\phi^2})}{1-\phi^2} + 2\frac{\beta^2\phi}{\lambda^2(1-\phi^2)(1-\phi)}(k-1+\frac{(\phi^{k-1}-1)(\phi-\phi^{k-1})}{1-\phi} - \frac{1-(\phi^2)^{k-1}}{1-\phi^2}) - 2\frac{\beta}{\lambda(1-\phi)}(k-1-\phi\frac{1-\phi^{k-1}}{1-\phi}) + k$$

$$= k\left(\frac{\beta^2}{\lambda^2}\frac{(1-\frac{1-(\phi^2)^k}{k(1-\phi^2)})}{1-\phi^2} + 2\frac{\beta^2\phi}{\lambda^2(1-\phi^2)(1-\phi)}(\frac{k-1}{k} + \frac{(\phi^{k-1}-1)(\phi-\phi^{k-1})}{k(1-\phi)} - \frac{1-(\phi^2)^{k-1}}{k(1-\phi^2)}) - 2\frac{\beta}{\lambda(1-\phi)}(\frac{k-1}{k} - \phi\frac{1-\phi^{k-1}}{k(1-\phi)}) + 1\right)$$

$$= k\left\{\frac{\beta^2}{\lambda^2(1-\phi)(1+\phi)}\left(1 - \frac{1-(\phi^2)^k}{k(1-\phi)(1+\phi)} + 2\frac{\phi}{(1-\phi)}(\frac{k-1}{k} + \frac{(\phi^{k-1}-1)(\phi-\phi^{k-1})}{k(1-\phi)} - \frac{1-(\phi^2)^{k-1}}{k(1-\phi)(1+\phi)})\right) - 2\frac{\beta}{\lambda(1-\phi)}(\frac{k-1}{k} - \phi\frac{1-\phi^{k-1}}{k(1-\phi)}) + 1\right\}$$

$$= k\left\{\left(\frac{\beta}{\lambda(1-\phi)}\right)^2\left(\frac{1-\phi}{1+\phi} - \frac{1-(\phi^2)^k}{k(1+\phi)(1+\phi)} + 2\frac{\phi}{(1+\phi)}\left(\frac{k-1}{k} + \frac{(\phi^{k-1}-1)(\phi-\phi^{k-1})}{k(1-\phi)} - \frac{1-(\phi^2)^{k-1}}{k(1-\phi)(1+\phi)}\right) - \left(\frac{k-1}{k} - \phi\frac{1-\phi^{k-1}}{k(1-\phi)}\right)^2\right) + \left(\frac{\beta}{\lambda(1-\phi)}\right)^2\right.$$

$$= k\left\{\left(\frac{\beta}{\lambda(1-\phi)}\right)^2\left(\frac{1-\phi}{1+\phi} - \frac{1-(\phi^2)^k}{k(1+\phi)(1+\phi)} + 2\frac{\phi}{(1+\phi)}\left(\frac{k-1}{k} + \frac{(\phi^{k-1}-1)(\phi-\phi^{k-1})}{k(1-\phi)} - \frac{1-(\phi^2)^{k-1}}{k(1-\phi)(1+\phi)}\right) - \left(\frac{k-1}{k} - \phi\frac{1-\phi^{k-1}}{k(1-\phi)}\right)^2\right) + \left(\left(\frac{\beta}{\lambda(1-\phi)}\right)\right.\right.$$

If we define $a(k;\beta,\phi,\lambda) \equiv 1 - \left(\frac{\beta}{\lambda(1-\phi)}\right)\left(\frac{k-1}{k} - \phi\frac{1-\phi^{k-1}}{k(1-\phi)}\right)$ then equation (11) could be written as

$$\frac{1}{k}C_t[r^{(k)}_{i,t+k}, r^{(k)}_{j,t+k}] = \sigma^{xc}_{CF,CF} + \left[a(k;\beta,\phi,\lambda)^2 + b(k;\beta,\phi,\lambda)\right]\sigma^{xc}_{DR,DR} - 2a(k;\beta,\phi,\lambda)\sigma^{xc}_{CF,DR} \tag{17}$$

where

$$b(k;\beta,\phi,\lambda) \equiv \left(\frac{\beta}{\lambda(1-\phi)}\right)^2\left(\frac{1-\phi}{1+\phi} - \frac{1-(\phi^2)^k}{k(1+\phi)(1+\phi)} + 2\frac{\phi}{(1+\phi)}\left(\frac{k-1}{k} + \frac{(\phi^{k-1}-1)(\phi-\phi^{k-1})}{k(1-\phi)} - \frac{1-(\phi^2)^{k-1}}{k(1-\phi)(1+\phi)}\right) - \left(\frac{k-1}{k} - \phi\frac{1-\phi^{k-1}}{k(1-\phi)}\right)^2\right) \tag{18}$$

we could show that $lim_{k\to+\infty}b(k;\beta,\phi,\lambda) = 0$.

Finally we have the asymptotic result

$$lim_{k\to+\infty}\ \frac{C_t[r^{(k)}_{i,t+k},r^{(k)}_{j,t+k}]}{k}=\sigma^{xc}_{CF,CF}+2(\frac{\beta}{\lambda(1-\phi)}-1)\sigma^{xc}_{CF,DR}+(\frac{\beta^2}{\lambda^2(1-\phi^2)}+\frac{2\beta^2\phi}{\lambda^2(1-\phi^2)(1-\phi)}-\frac{2\beta}{\lambda(1-\phi)}+1)\sigma^{xc}_{DR,DR} \tag{19}$$

Now we derive the range of the coefficients for variance-covariance terms in Eq (12), note that $\lambda=\frac{\rho\beta}{1-\rho\phi}$

$$\frac{\beta}{\lambda(1-\phi)}-1=\frac{1-\rho\phi}{\rho}\frac{1}{(1-\phi)}-1>\frac{1}{\rho}-1>0$$

and

$$\begin{aligned}&\frac{\beta^2}{\lambda^2(1-\phi^2)}+\frac{2\beta^2\phi}{\lambda^2(1-\phi^2)(1-\phi)}-\frac{2\beta}{\lambda(1-\phi)}+1\\
&=\left(\frac{\beta}{\lambda(1-\phi)}\right)^2-\frac{2\beta}{\lambda(1-\phi)}+1\\
&=\left(\frac{\beta}{\lambda(1-\phi)}-1\right)^2\\
&=\left(\frac{1-\rho\phi}{\rho-\rho\phi}-1\right)^2\end{aligned}$$

we know that ρ and ϕ are close to but smaller than 1, and if we assume that $\rho>\frac{1}{2-\phi}$, we have $\left(\frac{1-\rho\phi}{\rho-\rho\phi}-1\right)^2<1$. Thus we could have

$$0<\frac{\beta^2}{\lambda^2(1-\phi^2)}+\frac{2\beta^2\phi}{\lambda^2(1-\phi^2)(1-\phi)}-\frac{2\beta}{\lambda(1-\phi)}+1<1$$

under the assumption.

Numerical Calibration:

We try to use the formula to explain the positive gap between the portfolio variance of the benchmark case and the case in which integration is purely driven by increased DR news correlation. In our benchmark case, we set $\sigma^{xc}_{CF,CF}=\sigma^{xc}_{CF,DR}=\sigma^{xc}_{DR,DR}=0$, therefore

$$lim_{k\to+\infty}\ \sqrt{V_t[r^{(k)}_{p,t+k}]/k}=lim_{k\to+\infty}\ \sqrt{\frac{1}{N}V_t[r^{(k)}_{i,t+k}]/k} \tag{20}$$

. And for the integrated case purely driven by increased DR news correlation, we have

$$lim_{k\to+\infty}\ \sqrt{V_t[r^{(k)}_{p,t+k}]/k}=lim_{k\to+\infty}\ \sqrt{\frac{1}{N}V_t[r^{(k)}_{i,t+k}]/k+(1-\frac{1}{N})(\frac{\beta^2}{\lambda^2(1-\phi^2)}+\frac{2\beta^2\phi}{\lambda^2(1-\phi^2)(1-\phi)}-\frac{2\beta}{\lambda(1-\phi)}+1)\sigma^{xc}_{DR,DR}} \tag{21}$$

and we have

$$\frac{\beta^2}{\lambda^2(1-\phi^2)}+\frac{2\beta^2\phi}{\lambda^2(1-\phi^2)(1-\phi)}-\frac{2\beta}{\lambda(1-\phi)}+1=0.0175 \tag{22}$$

therefore explains the positive gap between the two variance plot in our 2 country symmetric experiment.

The coefficient of the term $\sigma^{xc}_{DR,DR}$ in Eq (11) standardized by k

$$\frac{1}{k}\left(\frac{\beta^2}{\lambda^2}\frac{(k-\frac{1-(\phi^2)^k}{1-\phi^2})}{1-\phi^2}+2\frac{\beta^2\phi}{\lambda^2(1-\phi^2)(1-\phi)}(k-1+\frac{(\phi^{k-1}-1)(\phi-\phi^{k-1})}{1-\phi}-\frac{1-(\phi^2)^{k-1}}{1-\phi^2})-2\frac{\beta}{\lambda(1-\phi)}(k-1-\phi\frac{1-\phi^{k-1}}{1-\phi})+k\right) \tag{23}$$

is a function of investment horizon k, and the coefficient annualized by k should converge to the value in Eq (15). The coefficient as a function of k is plotted in Figure 3.

In the next step, we calibrate the variance under the two cases (integration purely driven by increased cross country CF-CF/ DR-DR correlation). Under the limit case where $k\to+\infty$ we have

$$(\frac{\beta^2}{\lambda^2(1-\phi^2)}+\frac{2\beta^2\phi}{\lambda^2(1-\phi)^2}-\frac{2\beta}{\lambda(1-\phi)}+1)\sigma^{xc}_{DR,DR}=0.000010$$

where $\sigma^{xc}_{DR,DR} = \rho^{xc}_{DR,DR}\sigma_{DR}\sigma_{DR}$ and cross country DR correlation $\rho^{xc}_{DR,DR} = 0.25$. Similarly we get

$$\sigma^{xc}_{CF,CF} = \rho^{xc}_{CF,CF}\sigma_{CF}\sigma_{CF} = 0.0012$$

where $\rho^{xc}_{CF,CF} = 0.335$. In the calibration, we see that when integration purely driven by increased cross country CF-CF correlation, the impact on portfolio variance is permanent. When the integration is purely driven by increased cross country DR-DR correlation, the impact on portfolio variance is temporary, and dies out at long horizons. This matches with our intuition perfectly, and we see from the calibration that $(\frac{\beta^2}{\lambda^2(1-\phi^2)} + \frac{2\beta^2\phi}{\lambda^2(1-\phi)^2} - \frac{2\beta}{\lambda(1-\phi)} + 1)\sigma^{xc}_{DR,DR} \ll \sigma^{xc}_{CF,CF}$.

Lemma: Assuming
(1) $0.5 < \rho < 1$ and $0.5 < \phi < 1$ (trivially satisfied for time preference factor ρ and persistence of state variable ϕ).
(2) $\rho > \frac{2\phi^2+3\phi+1}{\phi^2+3\phi+2}$
We can conclude that the coefficient in Equation (16) is positive and decreasing in k (these are sufficient but not necessary conditions).

Proof: $f(k) \equiv \frac{1}{k}\left(\frac{\beta^2}{\lambda^2}\frac{(k-\frac{1-(\phi^2)^k}{1-\phi^2})}{1-\phi^2} + 2\frac{\beta^2\phi}{\lambda^2(1-\phi^2)(1-\phi)}(k-1+\frac{(\phi^{k-1}-1)(\phi-\phi^{k-1})}{1-\phi} - \frac{1-(\phi^2)^{k-1}}{1-\phi^2}) - 2\frac{\beta}{\lambda(1-\phi)}(k-1-\phi\frac{1-\phi^{k-1}}{1-\phi}) + k\right)$
$= \left(\frac{\beta^2}{\lambda^2}\frac{1}{1-\phi^2}(1-\frac{1-(\phi^2)^k}{k(1-\phi^2)}) + 2\frac{\beta^2\phi}{\lambda^2(1-\phi^2)(1-\phi)}(1-\frac{1}{k}+\frac{(\phi^{k-1}-1)(\phi-\phi^{k-1})}{k(1-\phi)} - \frac{1-(\phi^2)^{k-1}}{k(1-\phi^2)}) - 2\frac{\beta}{\lambda(1-\phi)}(1-\frac{1}{k}-\frac{\phi}{k}\frac{1-\phi^{k-1}}{1-\phi}) + 1\right)$
$= Const + \frac{1}{k}\left(-\frac{\beta^2}{\lambda^2}\frac{(1-\phi^k)(1+\phi^k)}{(1-\phi^2)^2} + 2\frac{\beta^2\phi}{\lambda^2(1-\phi^2)(1-\phi)}\frac{-1+\phi^2+(\phi^{k-1}-1)(\phi-\phi^{k-1})(1+\phi)-1+\phi^{2(k-1)}}{(1-\phi^2)} + 2\frac{\beta}{\lambda(1-\phi)}\frac{1-\phi^k}{1-\phi}\right)$
$= Const + \frac{1}{k}\left(-\frac{\beta^2}{\lambda^2}\frac{(1-\phi^k)(1+\phi^k)}{(1-\phi^2)^2} + 2\frac{\beta^2\phi}{\lambda^2(1-\phi)}\frac{(2+\phi-\phi^{k+1})(\phi^k-1)}{(1-\phi^2)^2} + 2\frac{\beta}{\lambda(1-\phi)}\frac{1-\phi^k}{1-\phi}\right)$
$= Const + \frac{1}{k}\frac{\beta}{\lambda}\frac{1-\phi^k}{(1-\phi)^2}\left(\frac{\beta}{\lambda}\frac{\phi^k(2\phi^2+\phi-1)-2\phi^2-3\phi-1}{(1+\phi)^2(1-\phi)} + 2\right)$
where

$$Const = \frac{\beta^2}{\lambda^2}\frac{1}{1-\phi^2} + 2\frac{\beta^2\phi}{\lambda^2(1-\phi^2)(1-\phi)} - 2\frac{\beta}{\lambda(1-\phi)} + 1$$
$$= \frac{\beta^2(1-\phi) + 2\beta^2\phi - 2\beta\lambda(1-\phi^2) + \lambda^2(1-\phi^2)(1-\phi)}{\lambda^2(1-\phi^2)(1-\phi)}$$
$$= \frac{(\beta-\lambda(1-\phi))^2}{\lambda^2(1-\phi)^2} > 0$$

Note that ρ and ϕ are close to but smaller than 1, and $\frac{\beta}{\lambda} = \frac{1-\rho\phi}{\rho}$. We want to find sufficient conditions so that $f(k)$ is decreasing in k. Since $f(k) = g(k)h(k)$ and $f'(k) = g\prime(k)h(k) + g(k)h'(k)$, $f'(k) < 0 \iff g(k)h'(k) < -g'(k)h(k)$. Since $g(k) > 0$, it will be sufficient if we could show that $g'(k) < 0$, $h'(k) < 0$ and $h(k) > 0$.

We first show that $g(k) \equiv \frac{1}{k}\frac{\beta}{\lambda}\frac{1-\phi^k}{(1-\phi)^2}$ decrease in k for $\phi \in (0,1)$. Take the first order derivative we get $g'(k) = \frac{\beta}{\lambda}\frac{1}{(1-\phi)^2}\frac{\phi^k(1-k\ ln\phi)-1}{k^2}$. To show $g'(k) < 0$, we need to show that $m(\phi) = \phi^k(1-k\ ln\phi) - 1 < 0$ for $\phi \in (0,1)$ and $\forall k$. This could be easily proved since $m'(\phi) = -k^2\phi^{k-1}ln(\phi) > 0$ for $\phi \in (0,1)$ and $m(1) = 0$. Thus $g(k)$ is positive and decrease in k. Then we want to know the property of $h(k) = \frac{\beta}{\lambda}\frac{\phi^k(2\phi^2+\phi-1)-2\phi^2-3\phi-1}{(1+\phi)^2(1-\phi)} + 2$. We also notice given that $2\phi^2+\phi-1 > 0$ (which hold as long as $\phi > 0.5$), $h(k)$ is decreasing in k. Thus it would be sufficient to prove the lemma if we know $h(k) > 0$ for $\forall k$. Since $h(k)$ is decreasing in k, we only need $\lim_{k\to\infty} h(k) = -\frac{\beta}{\lambda}\frac{2\phi^2+3\phi+1}{(1+\phi)^2(1-\phi)} + 2 = -\frac{1-\rho\phi}{\rho(1-\phi)}\frac{2\phi^2+3\phi+1}{(1+\phi)^2} + 2 > 0$ to hold. This is equivalent to $\rho > \frac{2\phi^2+3\phi+1}{\phi^2+3\phi+2}$. Under this condition, we know both $g(k)$ and $h(k)$ are positive and decreasing, therefore $f(k) = g(k)h(k)$ is positive and decreasing in k.

Appendix C. Symmatric Model for Asset Returns

We introduce a two-state-variable symmetric toy model for stocks, which includes excess stock return and dividend price ratio as state variables. In particular, the dynamics of the variables are given by:

$$xr_{s,t+1} = \lambda_{xr,0} + \lambda_{xr,1}(d_t - p_t) + v_{xr,t+1} \tag{24}$$

$$d_{t+1} - p_{t+1} = \lambda_{dp,0} + \lambda_{dp,1}(d_t - p_t) + v_{dp,t+1} \tag{25}$$

We denote $v_t = [v_{xr,t}, v_{dp,t}]'$ and assume the VAR shocks are covariance stationary $E(v_t) = \mathbf{0}$, $E(v_t v_s) = \begin{cases} \Sigma^{wc} & (t = s) \\ \mathbf{0} & (t \neq s) \end{cases}$.The superscript wc stands for within-country, and we use xc to represent cross-country in later part of the paper.

C.1 Connect VAR shocks to structural shocks

We decompose stock excess returns into two structural shocks: cash flow news and discount rate news. In the toy model VAR with two state variables, there's actually a one-to-one mapping from the structural shocks to VAR shocks. Recall from the decomposition

$$N_{RR,t+1} \equiv (E_{t+1} - E_t) \left[\sum_{j=0}^{\infty} \rho_s^j r_{f,t+1+j}^N \right] = (E_{t+1} - E_t) \left[\sum_{j=0}^{\infty} \rho_s^j \left(y_{1,t+j}^{N,\$} - \pi_{t+1+j} \right) \right] = 0$$

This is because the short nominal rate and inflation are assume to be zero in our toy model.

$$N_{RP,t+1} \equiv (E_{t+1} - E_t) \left[\sum_{j=0}^{\infty} \rho_s^j xr_{s,t+1+j} \right] = \frac{\rho_s \lambda_{xr,1}}{1 - \rho_s \lambda_{dp,1}} v_{dp,t+1}$$

Therefore we have the discount rate news

$$N_{DR,t+1} = N_{RR,t+1} + N_{RP,t+1} = \frac{\rho_s \lambda_{xr,1}}{1 - \rho_s \lambda_{dp,1}} v_{dp,t+1}$$

and the cash flow news is calculated from the identity

$$N_{CF,t+1} = (E_{t+1} - E_t) \left[xr_{s,t+1} \right] + N_{DR,t+1} = v_{xr,t+1} + \frac{\rho_s \lambda_{xr,1}}{1 - \rho_s \lambda_{dp,1}} v_{dp,t+1}$$

To summarize, we have

$$\begin{bmatrix} N_{CF,t+1} \\ N_{DR,t+1} \end{bmatrix} = \begin{bmatrix} 1 & \frac{\rho_s \lambda_{xr,1}}{1-\rho_s \lambda_{dp,1}} \\ 0 & \frac{\rho_s \lambda_{xr,1}}{1-\rho_s \lambda_{dp,1}} \end{bmatrix} \begin{bmatrix} v_{xr,t+1} \\ v_{dp,t+1} \end{bmatrix} \tag{26}$$

which connects the VAR shocks to structural shocks. Or in matrix notation $\varepsilon_{t+1} = P v_{t+1}$, where ε_{t+1} is the structural shock, v_{t+1} the VAR shocks and P the transformation matrix.

C.2 From single country to a world with N identical countries

To further explore the benefit of international diversification, we design an experiment in a world with N clones (N-replica world composed of N identical countries, and we use the US data to get empirical results). To explain the experiment in detail, we first introduce some notations. Let $\Sigma^{wc} \equiv Var(v_{t+1})$ be the within country VAR variance-covariance matrix, and $\Sigma^{xc} \equiv Cov(v_{i,t+1}, v_{j,t+1})$ $(i \neq j)$ is defined as the cross-country VAR variance-covariance matrix (between country i and j). Since all variance-covariance matrix Σ could be decomposed into volatility component $G \equiv diag(\Sigma)^{1/2}$ and correlation component $(\Gamma \equiv diag(\Sigma)^{-1/2} \Sigma diag(\Sigma)^{-1/2})$, we have the following decomposition for within-country and cross-country VAR variance-covariance matrix

$$\Sigma^{wc} \equiv G_\Sigma \Gamma_\Sigma^{wc} G_\Sigma' \tag{27}$$

$$\Sigma^{xc} \equiv G_\Sigma \Gamma_\Sigma^{xc} G_\Sigma' \tag{28}$$

By using this notation we have implicitly assumed all countries are identical, i.e. $\Sigma_i^{wc} = \Sigma_j^{wc}$ and $\Sigma_{ij}^{xc} = \Sigma_{lm}^{xc}$ $(i \neq j, l \neq m)$, which also implies $G_{\Sigma,i} = G_{\Sigma,j}$, $\Gamma_{\Sigma,i}^{wc} = \Gamma_{\Sigma,j}^{wc}$, $\Gamma_{\Sigma,ij}^{xc} = \Gamma_{\Sigma,lm}^{xc}$.

Then the variance-covariance matrix for the global VAR shock in the N-replica economy is

$$\Sigma_{glo} = \begin{bmatrix} \Sigma^{wc} & \Sigma^{xc} & \cdots & \Sigma^{xc} \\ \Sigma^{xc} & \Sigma^{wc} & \cdots & \Sigma^{xc} \\ \vdots & \vdots & \cdots & \vdots \\ \Sigma^{xc} & \Sigma^{xc} & \cdots & \Sigma^{wc} \end{bmatrix}$$

with Σ^{wc} as diagonal blocks and Σ^{xc} as off diagonal blocks . Later we use Σ_{glo} international portfolio allocation analysis.

C.3 Connect the VAR variance-covariance matrix to structural variance-covariance matrix in a world with N identical countries

Let $\Omega^{wc} \equiv Var(\varepsilon_{t+1})$ be the within country structural variance-covariance matrix, and $\Omega^{xc} \equiv Cov(\varepsilon_{i,t+1}, \varepsilon_{j,t+1})$ $(i \neq j)$ is defined as the cross-country structural variance-covariance matrix (between country i and j). Analogous to the decomposition above, we have

$$\Omega^{xc} \equiv G_\Omega \Gamma_\Omega^{xc} G'_\Omega \tag{29}$$

$$\Omega^{wc} \equiv G_\Omega \Gamma_\Omega^{wc} G'_\Omega \tag{30}$$

From the relation $\varepsilon_{t+1} = Pv_{t+1}$, we can take cross-country covariance $Cov(\varepsilon_{i,t+1}, \varepsilon_{j,t+1}) = PCov(v_{i,t+1}, v_{j,t+1})P'$ and get an identity $\Omega^{xc} = P\Sigma^{xc}P'$. Of course, $\Omega^{wc} = P\Sigma^{wc}P'$ also holds.

The identity could be rewritten as

$$G_\Omega \Gamma_\Omega^{xc} G'_\Omega = PG_\Sigma \Gamma_\Sigma^{xc} G'_\Sigma P' \tag{31}$$

Applying the vec operator to both sides and using the trick that $vec(ABC) = (C' \otimes A) \cdot vec(B)$ (see Hamilton 1994 Proposition 10.4) we have

$$(G_\Omega \otimes G_\Omega) \cdot vec\,(\Gamma_\Omega^{xc}) = ((PG_\Sigma) \otimes (PG_\Sigma)) \cdot vec\,(\Gamma_\Sigma^{xc}) \tag{32}$$

Now we've got a mapping from cross-country structural shock correlation matrix to cross-country VAR shock correlation matrix. If $((PG_\Sigma) \otimes (DG_\Sigma))$ is nonsingular, we could rewrite the relationship as

$$vec\,(\Gamma_\Sigma^{xc}) = ((PG_\Sigma) \otimes (PG_\Sigma))^{-1} (G_\Omega \otimes G_\Omega) \cdot vec\,(\Gamma_\Omega^{xc}) \tag{33}$$

And similarly, we have

$$(G_\Omega \otimes G_\Omega) \cdot vec\,(\Gamma_\Omega^{wc}) = ((PG_\Sigma) \otimes (PG_\Sigma)) \cdot vec\,(\Gamma_\Sigma^{wc}) \tag{34}$$

We could also analogously define the variance-covariance matrix for the global structural shock

$$\Omega_{glo} = \begin{bmatrix} \Omega^{wc} & \Omega^{xc} & \cdots & \Omega^{xc} \\ \Omega^{xc} & \Omega^{wc} & \cdots & \Omega^{xc} \\ \vdots & \vdots & \cdots & \vdots \\ \Omega^{xc} & \Omega^{xc} & \cdots & \Omega^{wc} \end{bmatrix}$$

And equations (33) and (34) give us the connection between Ω_{glo} and Σ_{glo}.

C.4 Illustrative example using the symmetric model

From the analysis above, we know there's a connection between the global structural shocks and global VAR shocks. And we could design some experiments using this connection to study the effect of international integration on portfolio allocation. Empirically, we follow the steps below:

1. Estimate a single country toy model using the US historical data. From this we could get a estimate for the variance-covariance matrix Σ^{wc} (or equivalently G_Σ and Γ_Σ^{wc}). P matrix could also be calculated from the reduced form VAR coefficients.
2. Using the identity $\Omega^{wc} = P\Sigma^{wc}P'$, we have an estimate of Ω^{wc}(or equivalently G_Ω and Γ_Ω^{wc}).
3. Manually set values for the cross-country structural shock correlation matrix Γ_Ω^{xc}. From equation (?) we will be able to get the implied cross-country VAR shock correlation matrix Γ_Σ^{xc}.
4. Construct the implied global VAR variance-covariance matrix Σ_{glo}, based on our input Γ_Ω^{xc} in step 3. Given Σ_{glo}, we could study the implications of international integration on global portfolio allocation.

Specifically, we assign 3 set of values toΓ_Ω^{xc} in step 3 above, each corresponds a scenario below :

1st Scenario). $\Gamma_{\Omega}^{xc} = \mathbf{0}$
This is a benchmark case without international integration, where all cross-country structural shocks are uncorrelated.
2nd Scenario). $\Gamma_{\Omega}^{xc} = \begin{bmatrix} \Gamma_{\Omega,11}^{xc} & 0 \\ 0 & 0 \end{bmatrix}$
where $\Gamma_{\Omega,11}^{xc}$ denote the cross-country CF news correlation.
This is a case with international integration, and the integration is purely driven by increased CF news correlation:
3rd Scenario). $\Gamma_{\Omega}^{xc} = \begin{bmatrix} 0 & 0 \\ 0 & \Gamma_{\Omega,22}^{xc} \end{bmatrix}$
where $\Gamma_{\Omega,22}^{xc}$ denote the cross-country DR news correlation.
This is a case with international integration, and the integration is purely driven by increased DR news correlation.

C.5 Implied Correlation Structure of VAR in Section 3.3

	First Scenario		Second Scenario		Third Scenario	
Corr	$u_{xr,s}$	u_{dp}	$u_{xr,s}$	u_{dp}	$u_{xr,s}$	u_{dp}
$u_{xr,s}$	0	0	0.070	0	0.070	-0.087
u_{dp}	0	0	0	0	-0.087	0.109

C.6 From 2 state variables (toy model) to 6 state variables (general model)

It's very easy to incorporate the toy model in a more general framework. Recall that our general model for a single country is a VAR with 6 state variables

$$\tilde{\mathbf{z}}_{t+1} = a + \mathbf{A}\tilde{\mathbf{z}}_t + \mathbf{u}_{t+1}$$

where $\tilde{\mathbf{z}}_{t+1} = [xr_{s,t+1}, xr_{n,t+1}, d_{t+1} - p_{t+1}, \pi_{t+1}, y_{1,t+1}, y_{10,t+1} - y_{1,t+1}]$. Toy model is a special case of the general model with

$$a = \begin{bmatrix} \lambda_{xr,0} \\ 0 \\ \lambda_{dp,0} \\ 0 \\ 0 \\ 0 \end{bmatrix}$$

$$\mathbf{A} = \begin{bmatrix} 0 & 0 & \lambda_{xr,1} & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & \lambda_{dp,1} & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 \end{bmatrix}$$

and

$$\mathbf{u}_{t+1} = \begin{bmatrix} v_{xr,t+1} \\ 0 \\ v_{dp,t+1} \\ 0 \\ 0 \\ 0 \end{bmatrix}$$

Appendix D. Data Description

We consider a number of time series from 7 major OECD countries, which accounts for 62% of total world market shares by end of 2014.The full sample period is 1986:01 to 2013:12, yielding 336 monthly observations. We split the full sample to two sub-periods, with the sub-period 1 from 1986:01 to 1999:12 and the sub-period 2 from 2000:01 to 2013:12. Returns are in U.S. dollar currency-hedged terms in excess of the three-month U.S. Treasury bill rate.

D.1 Currency-hedged Return

Before further explaining our data in details, we first introduce the concept of currency hedged excess return. Consider a home investor from US buying assets in a foreign country (for example in Japan), we are interested in his excess returns from this investment denominated in home currency. We use a superscript $*$ to denote a foreign variable. S_t denotes the spot foreign exchange rate, and an increase in S_t means home currency is weakening relative to foreign currency. To conduct this trade, the investor at time t has to exchange 1 US dollar into $\frac{1}{S_t}$ Japanese yen and invest in Japanese capital market, then converts the money back to USD at time $t+1$ when the investment is liquidated. Thus the (unhedged) 1-period return in Japanese market (measured in dollars) is

$$1 + R_{JPN,t+1} \equiv (1 + R^*_{JPN,t+1})\frac{S_{t+1}}{S_t}$$

where $R^*_{JPN,t+1}$ is return in Japanese asset denominated in Japanese yen (local return).

However, due to the uncertainty in future exchange rate S_{t+1}, the investor will want to lock down the future exchange rate using a currency forward at forward rate F_t. So the currency hedged return of a US investor investing in Japan is defined as

$$1 + R^h_{JPN,t+1} \equiv (1 + R^*_{JPN,t+1})\frac{F_t}{S_t}$$

Recall from the covered interest rate parity (CIP), we also have

$$1 + i_{US,t+1} = (1 + i^*_{JPN,t+1})\frac{F_t}{S_t}$$

where $i_{US,t+1}$ is the nominal interest rate for the US, while $i_{JPN,t+1}$ is the nominal interest rate for Japan. The intuition for this equation is that the investor should not have arbitrage opportunities, or alternatively, should be indifferent to invest locally or abroad if the currency risk of investing in foreign country is hedged. This equation holds pretty well unless there's counter-party risk or barriers to financial integration (transaction costs, taxes, capital controls, et cetera).

Combining equations (?) and (?), we know that the excess currency hedged return of a US investor investing in Japan is

$$\frac{1 + R^h_{JPN,t+1}}{1 + i_{US,t+1}} = \frac{1 + R^*_{JPN,t+1}}{1 + i^*_{JPN,t+1}}$$

or in log terms

$$r^h_{JPN,t+1} - r_{f,US,t+1} = r^*_{JPN,t+1} - r^*_{f,JPN,t+1}$$

where $r_{f,US,t+1} = \ln(1 + i_{US,t+1})$ and $r_{f,JPN,t+1} = \ln(1 + i^*_{JPN,t+1})$ are the risk free rates in US and Japan. Thus, we have shown that the excess currency-hedged return of US investors investing in Japan is the same as the excess return of Japanese investors investing in home country (local excess return).

D.2 Main Variables

Now we introduce our main variables briefly.

Returns, Dividend Yield and Inflation

The international portfolio we consider are constructed from country level index in equity and bonds. The country level stock returns are measured as dollar returns on MSCI net total return indices, which reinvest dividends after the deduction of withholding taxes. We use Merill Lynch total return indices (7yr-10yr) to get bond returns. The dividend yield is measured as the log of MSCI dividend yield (MSDY), which is calculated using the trailing 12-month cash earnings per share figure. All the data on stock and bond returns as well as dividend yields are from Datastream. Table 2.A reports sample correlations of monthly bond and stock returns for the period January 1986 to December 2013. Returns are in U.S. Dollar currency-hedged terms in excess of the three-month U.S. Treasury bill rate. (???). Table 2.B and 2.C further look at the correlations in the two sub-samples we are studying.

For the inflation, we get data from both Datastream and Global Financial Data (GFD). We first get annualized inflation rates from Datastream. But for France and UK, the data does not go back far enough because data comes from newer HICP that started in 1990's; thus, we compute inflation manually using CPI for France and RPI for UK from GSD.

Foreign Exchange Rates

We get spot currency levels and one-month forward currency levels from Datastream. The currency levels are all in terms of 1 US dollar except for British Pound (GBP), so we invert GBP to get correct reference frame. The (unhedged) currency returns are calculated as $\ln(\frac{S_{t+1}}{S_t})$ for spot currency levels for 1 USD, and the currency-hedged returns are calculated as $\ln \frac{F_t}{S_t}$ for forward and spot currency levels for 1 USD. Note that French and German data switch to Euros at the beginning of 1999.

Short Term and Long Term Nominal Interest Rate

We use 1 month T-bill rate for US short term nominal interest rate, and for other countries we use different rates on short term financial instruments including 1 month Euribor rates, bank loan rates or overnight money market interest rates. The data are from GFD and central bank websites. Long term nominal interest rate are represented using 10 year yields. The US series is from CRSP Fixed Term Indices and other countries from GFD.

Data Source

Variable	Source	Description	Download Information
Equity Index	Datastream	MSCI net returns in USD using MSNR (net dividends reinvested); sheet also contains MSCI price indices in USD using MSPI (no dividends reinvested) and MSCI return indices in USD using MSRI (gross dividends reinvested); get returns with simple division of levels; can also get local returns as opposed to USD returns. Take simple USD returns from MSNR and takes LN of gross returns.	MSAUSTL, MSCNDAL, MSFRNCL, MSGERML, MSJPANL, MSUTDKL, MSUSAML with fields MSNR, MSPI, or MSRI
Dividend yields	Datastream	Dividend yields; take LN	MSAUSTL, MSCNDAL, MSFRNCL, MSGERML, MSJPANL, MSUTDKL, MSUSAML with field MSDY
Bond Index	Datastream	Merrill Lynch total return indices; get simple returns with simple division of levels; numbers are already in USD. We take only 7y-10y sector TR and takes LN of gross returns	Datastream tickers: MLAD1T3, MLAD3T5, MLAD5T7, MLAD710, MLCD1T3, MLCD3T5, MLCD5T7, MLCD710, MLFF1T3, MLFF3T5, MLFF5T7, MLFF710, MLDM1T3, MLDM3T5, MLDM5T7, MLDM710, MLJP1T3, MLJP3T5, MLJP5T7, MLJP710, MLUK1T3, MLUK3T5, MLUK5T7, MLUK710, MLUS1T3, MLUS3T5, MLUS5T7, MLUS710
Inflation	Datastream and Global Financial Data(GFD)	Get annualized inflation rates from Datastream and take monthly differences to account for seasonality; for France and UK, data does not go back far enough because data comes from newer HICP that started in 1990's; thus, use GFD to get older CPI for France and RPI for UK and manually compute inflation. We take LN of 1 + monthly difference.	Datastream tickers: AUCPANNL, BDCPANNL, CNCPANNL, FRCPANNL, JPCPANNL, UKCPANNL, USCPANNL; GFD tickers: CPAUSM, CPCANM, CPFRAM (this is French CPI), CPHFRAM (this is French HICP), CPDEUM, CPJPNM, CPGBRM (this is UK RPI), CPHGBRM (this is UK HICP), CPUSAM

FX Data (spot and forward currency level)	Datastream	Currency returns calculated as LN(SPOT(t+1)/SPOT(t)) for SPOT currency levels for 1 USD; hedged currency returns calculated as LN(FWD(t)/SPOT(t)) for FWD and SPOT currency levels for 1 USD; note that French and German data switch to Euros at the beginning of 1999	Get spot currency levels with BBAUDSP, BBCADSP, BBFRFSP, BBDEMSP, BBJPYSP, BBGBPSP, BBEURSP - these are all in terms of 1 USD except for GBP, so need to invert GBP to get correct reference frame; get 1m forward currency levels with BBAUD1F, BBCAD1F, BBFRF1F, BBDEM1F, BBJPY1F, BBGBP1F, BBEUR1F - these are all in terms of 1 USD except for GBP, so need to invert GBP to get correct reference frame
Short Term Interest Rate	GFD and websites	Short nominal rates; Australia: target FF rates; Canada: bank rates, which are discount rates or +25bp over target FF rates; France/Germany: 1 month Euribor rates; Japan: basic discount rates/basic loan rates; UK: bank rates, which are discount rates; US: 12*RF where RF is the 1 month T-bill rate; take LN of (1+SR) as defined above and divides by 12 to get monthly figure	Australia: GFD (from Global Currency Hedging paper) until 200605, then from `http://www.rba.gov.au/statistics/cash-rate.html`; Canada: `http://www.bankofcanada.ca/rates/interest-rates/canadian-interest-rates/`; France: GFD (from Global Currency Hedging paper) until 200412, then from `http://www.global-rates.com/interest-rates/euribor/2010.aspx`; Germany: GFD (from Global Currency Hedging paper) until 200412, then from `http://www.global-rates.com/interest-rates/euribor/2010.aspx`; Japan: `http://www.boj.or.jp/en/statistics/boj/other/discount/index.htm/`; UK: `http://www.bankofengland.co.uk/mfsd/iadb/Repo.asp?Travel=NIxRPx`; US: from Ken French's website
Long Term Interest Rate	GFD and CRSP	Long nominal rates; essentially CMT at 5y and 10y points; takes LN of 1 + LR using the 10y point and divides by 12 to get monthly figure	For non-US, use GFD and the following symbols: IGAUS5D, IGCANB5D, IGFRA5D, IGDEU5D, IGJPN5D, IGGBR5D; IGAUS10D, IGCAN10D, IGFRA10D, IGDEU10D, IGJPN10D, IGGBR10D for US, use CRSP Fixed Term Indices (Daily Series of Yield to Maturity) and the data for 2014 comes from, taking the yield at the end of each month `http://www.treasury.gov/resource-center/data-chart-center/interest-rates/Pages/TextView.aspx?data=yieldYear&year=2014`
Market Capitalization	World Bank	Market capitalization of each country	"Market capitalization of listed companies (current US$)" on world bank website `http://data.worldbank.org/indicator/CM.MKT.LCAP.CD/countries`

Appendix E. VAR Model Estimation

Table E1. Pooled VAR(1) Model Estimates

Panel A

Model estimates	Coefficients on lagged variables (1)	(2)	(3)	(4)	(5)	(6)	Rsq
(1) log stock excess returns	0.087	0.030	0.013	0.025	-0.985	1.074	0.014
	(2.288)	(0.283)	(2.086)	(0.057)	(-0.796)	(0.471)	
(2) log bond excess returns	-0.049	0.083	0.002	-0.268	0.724	2.590	0.047
	(-4.413)	(2.492)	(1.343)	(-1.930)	(1.934)	(3.613)	
(3) log dividend yield	-0.088	-0.093	0.977	0.123	-0.036	-3.645	0.961
	(-2.190)	(-0.822)	(149.138)	(0.257)	(-0.028)	(-1.512)	
(4) log inflation	0.004	-0.013	0.000	0.177	0.274	0.019	0.093
	(2.600)	(-2.731)	(-0.286)	(6.896)	(5.421)	(0.196)	
(5) log short rate	0.000	-0.002	0.000	0.004	1.007	0.075	0.978
	(1.294)	(-4.294)	(-1.857)	(2.164)	(194.594)	(6.734)	
(6) log yield spread	0.000	0.001	0.000	-0.002	-0.018	0.900	0.860
	(1.739)	(1.125)	(0.440)	(-0.748)	(-2.676)	(65.320)	

Panel B

Within-country Residual Correlation Matrix (1986-2013)
averaged over 7 countries
average annualized volatility*100 in diagonal

	(1)	(2)	(3)	(4)	(5)	(6)
(1) log stock excess returns	**18.150**	0.081	-0.896	0.014	-0.024	-0.036
(2) log bond excess returns	0.081	**5.853**	-0.066	-0.059	-0.172	-0.474
(3) log dividend yield	-0.896	-0.066	**20.211**	0.032	0.036	0.030
(4) log inflation	0.014	-0.059	0.032	**1.074**	0.051	-0.001
(5) log short rate	-0.024	-0.172	0.036	0.051	**0.106**	-0.719
(6) log yield spread	-0.036	-0.474	0.030	-0.001	-0.719	**0.123**

Cross-country Residual Correlation Matrix (1986-2013)
averaged over 7 countries
diagonal terms are average cross-country correlation of the same state variable

	(1)	(2)	(3)	(4)	(5)	(6)
(1) log stock excess returns	**18.150**	0.081	-0.896	0.014	-0.024	-0.036
(2) log bond excess returns	0.081	**5.853**	-0.066	-0.059	-0.172	-0.474
(3) log dividend yield	-0.896	-0.066	**20.211**	0.032	0.036	0.030
(4) log inflation	0.014	-0.059	0.032	**1.074**	0.051	-0.001
(5) log short rate	-0.024	-0.172	0.036	0.051	**0.106**	-0.719
(6) log yield spread	-0.036	-0.474	0.030	-0.001	-0.719	**0.123**

Panel C

Within-country Residual Correlation Matrix (1986.01-1999.12) averaged over 7 countries diagonal terms are annualized average volatility*100	(1)	(2)	(3)	(4)	(5)	(6)
(1) log stock excess returns	19.215	0.298	-0.927	-0.030	-0.093	-0.105
(2) log bond excess returns	0.298	6.430	-0.290	-0.066	-0.189	-0.445
(3) log dividend yield	-0.927	-0.290	20.859	0.061	0.086	0.113
(4) log inflation	-0.030	-0.066	0.061	1.028	0.039	0.020
(5) log short rate	-0.093	-0.189	0.086	0.039	0.136	-0.726
(6) log yield spread	-0.105	-0.445	0.113	0.020	-0.726	0.153

Cross-country Residual Correlation Matrix (1986.01-1999.12) averaged over 7 countries diagonal terms are average cross-country correlation of the same state variable	(1)	(2)	(3)	(4)	(5)	(6)
(1) log stock excess returns	0.540	0.185	-0.512	-0.022	-0.034	-0.075
(2) log bond excess returns	0.070	0.413	-0.073	-0.018	-0.059	-0.226
(3) log dividend yield	-0.528	-0.178	0.513	0.032	0.028	0.079
(4) log inflation	-0.054	-0.052	0.063	0.098	0.031	0.030
(5) log short rate	-0.046	-0.055	0.042	0.008	0.091	-0.050
(6) log yield spread	-0.012	-0.231	0.017	0.004	-0.030	0.190

Panel D

Within-country Residual Correlation Matrix (2000.01-2013.12) averaged over 7 countries diagonal terms are annualized average volatility*100	(1)	(2)	(3)	(4)	(5)	(6)
(1) log stock excess returns	16.914	-0.217	-0.864	0.055	0.112	0.099
(2) log bond excess returns	-0.217	5.158	0.233	-0.043	-0.131	-0.573
(3) log dividend yield	-0.864	0.233	19.312	0.003	-0.052	-0.129
(4) log inflation	0.055	-0.043	0.003	1.090	0.072	-0.017
(5) log short rate	0.112	-0.131	-0.052	0.072	0.060	-0.672
(6) log yield spread	0.099	-0.573	-0.129	-0.017	-0.672	0.079

Cross-country Residual Correlation Matrix (2000.01-2013.12) averaged over 7 countries diagonal terms are average cross-country correlation of the same state variable	(1)	(2)	(3)	(4)	(5)	(6)
(1) log stock excess returns	0.712	-0.208	-0.604	0.017	0.103	0.105
(2) log bond excess returns	-0.204	0.607	0.193	-0.025	-0.040	-0.408
(3) log dividend yield	-0.635	0.200	0.571	0.019	-0.034	-0.139
(4) log inflation	0.057	-0.062	-0.022	0.243	0.111	-0.038
(5) log short rate	0.097	-0.018	-0.050	0.065	0.244	-0.192
(6) log yield spread	0.079	-0.403	-0.105	-0.034	-0.163	0.457

Table E2. VAR(1) Model Estimates [Australia]

Panel A. Model estimates							
	Coefficients on lagged variables						
	(1)	(2)	(3)	(4)	(5)	(6)	Rsq
(1) log stock excess returns	0.038	-0.257	0.026	0.422	-1.228	0.589	0.019
	0.854	-1.425	1.042	0.374	-0.690	0.234	
(2) log bond excess returns	-0.044	0.096	0.007	-0.338	0.757	2.572	0.043
	-2.122	1.577	1.088	-0.728	1.457	2.492	
(3) log dividend yield	-0.065	0.267	0.945	0.731	0.507	-4.199	0.920
	-1.145	1.334	35.173	0.548	0.266	-1.416	
(4) log inflation	0.001	-0.003	0.000	0.735	0.127	0.013	0.701
	0.674	-1.158	-0.675	9.904	2.474	0.212	
(5) log short rate	0.000	0.002	0.000	0.046	0.982	0.179	0.948
	0.653	0.840	0.506	2.311	38.719	3.525	
(6) log yield spread	0.000	-0.003	0.000	-0.046	0.007	0.787	0.701
	0.235	-1.691	-0.972	-2.307	0.258	14.980	

Panel B. Residual correlation matrix						
	(1)	(2)	(3)	(4)	(5)	(6)
(1) log stock excess returns	17.555	0.222	-0.918	-0.006	-0.044	-0.032
(2) log bond excess returns	0.222	6.606	-0.185	-0.054	-0.060	-0.280
(3) log dividend yield	-0.918	-0.185	19.470	0.008	0.030	0.040
(4) log inflation	-0.006	-0.054	0.008	0.455	0.092	-0.072
(5) log short rate	-0.044	-0.060	0.030	0.092	0.226	-0.935
(6) log yield spread	-0.032	-0.280	0.040	-0.072	-0.935	0.241

Table E3. VAR(1) Model Estimates [Canada]

Panel A. Model estimates							
	Coefficients on lagged variables						
	(1)	(2)	(3)	(4)	(5)	(6)	Rsq
(1) log stock excess returns	0.112	0.118	0.009	0.480	-1.297	1.496	0.026
	1.801	0.914	0.787	0.665	-0.846	0.509	
(2) log bond excess returns	-0.083	0.027	0.001	0.100	0.554	2.735	0.059
	-2.843	0.379	0.312	0.287	1.022	2.161	
(3) log dividend yield	-0.127	-0.203	0.978	-0.401	-0.424	-5.249	0.969
	-1.938	-1.500	65.919	-0.453	-0.249	-1.651	
(4) log inflation	0.008	-0.013	0.000	0.075	0.280	-0.104	0.076
	1.607	-1.277	-0.453	1.075	2.509	-0.485	
(5) log short rate	0.000	-0.004	0.000	-0.001	1.005	0.035	0.988
	-0.161	-3.094	-1.781	-0.115	115.610	1.620	
(6) log yield spread	0.001	0.002	0.000	-0.001	-0.013	0.943	0.930
	2.065	2.211	1.387	-0.332	-1.395	44.625	

Panel B. Residual correlation matrix						
	(1)	(2)	(3)	(4)	(5)	(6)
(1) log stock excess returns	15.645	0.138	-0.910	0.089	-0.019	-0.046
(2) log bond excess returns	0.138	6.027	-0.129	0.009	-0.314	-0.342
(3) log dividend yield	-0.910	-0.129	17.650	-0.044	0.038	0.042
(4) log inflation	0.089	0.009	-0.044	1.166	0.020	-0.009
(5) log short rate	-0.019	-0.314	0.038	0.020	0.099	-0.733
(6) log yield spread	-0.046	-0.342	0.042	-0.009	-0.733	0.103

Table E4. VAR(1) Model Estimates [France]

Panel A. Model estimates							
	Coefficients on lagged variables						
	(1)	(2)	(3)	(4)	(5)	(6)	Rsq
(1) log stock excess returns	0.106	0.414	0.007	2.071	-0.175	4.137	0.037
	1.460	1.808	0.571	1.677	-0.068	0.681	
(2) log bond excess returns	-0.031	0.041	0.008	-0.806	0.843	2.447	0.060
	-1.950	0.610	2.553	-2.483	1.399	1.764	
(3) log dividend yield	-0.098	-0.574	0.969	-1.545	-0.614	-6.157	0.939
	-1.315	-2.474	66.193	-1.090	-0.225	-0.960	
(4) log inflation	0.004	-0.014	0.000	0.039	0.206	0.127	0.048
	1.745	-1.572	0.075	0.698	2.306	0.620	
(5) log short rate	0.000	-0.003	0.000	0.001	1.017	0.071	0.991
	-0.357	-3.312	-3.160	0.170	97.311	1.879	
(6) log yield spread	0.000	0.002	0.000	0.009	-0.031	0.904	0.925
	1.581	1.615	1.030	1.482	-2.531	22.376	

Panel B. Residual correlation matrix						
	(1)	(2)	(3)	(4)	(5)	(6)
(1) log stock excess returns	19.612	0.118	-0.858	-0.041	-0.021	-0.068
(2) log bond excess returns	0.118	5.304	-0.038	-0.085	-0.161	-0.467
(3) log dividend yield	-0.858	-0.038	22.475	0.138	-0.011	0.059
(4) log inflation	-0.041	-0.085	0.138	0.845	0.080	-0.011
(5) log short rate	-0.021	-0.161	-0.011	0.080	0.084	-0.757
(6) log yield spread	-0.068	-0.467	0.059	-0.011	-0.757	0.100

Table E5. VAR(1) Model Estimates [Germany]

Panel A. Model estimates							
	Coefficients on lagged variables						
	(1)	(2)	(3)	(4)	(5)	(6)	Rsq
(1) log stock excess returns	0.090	-0.112	0.021	-0.330	-2.835	2.763	0.024
	1.348	-0.463	1.398	-0.310	-1.269	0.487	
(2) log bond excess returns	-0.043	0.015	0.001	-0.392	0.077	1.254	0.050
	-3.142	0.240	0.325	-1.516	0.136	1.076	
(3) log dividend yield	-0.114	-0.024	0.953	0.168	0.477	-8.062	0.927
	-1.638	-0.093	59.004	0.148	0.206	-1.398	
(4) log inflation	0.004	-0.019	0.000	-0.112	0.140	-0.525	0.050
	1.593	-1.991	0.675	-1.840	0.998	-2.111	
(5) log short rate	0.000	-0.003	0.000	0.003	1.008	0.035	0.992
	1.131	-4.338	-1.717	0.791	166.971	1.878	
(6) log yield spread	0.000	0.002	0.000	0.003	-0.012	0.957	0.940
	1.397	2.736	1.370	0.873	-1.519	47.945	

Panel B. Residual correlation matrix						
	(1)	(2)	(3)	(4)	(5)	(6)
(1) log stock excess returns	22.027	-0.077	-0.873	0.066	0.076	-0.045
(2) log bond excess returns	-0.077	5.025	0.080	-0.104	-0.324	-0.487
(3) log dividend yield	-0.873	0.080	23.996	-0.039	-0.036	0.023
(4) log inflation	0.066	-0.104	-0.039	1.120	0.027	0.060
(5) log short rate	0.076	-0.324	-0.036	0.027	0.057	-0.587
(6) log yield spread	-0.045	-0.487	0.023	0.060	-0.587	0.070

Table E6. VAR(1) Model Estimates [Japan]

Panel A. Model estimates							
	Coefficients on lagged variables						
	(1)	(2)	(3)	(4)	(5)	(6)	Rsq
(1) log stock excess returns	0.120	0.131	0.010	-1.240	0.001	4.789	0.024
	1.742	0.565	0.980	-1.462	0.000	0.645	
(2) log bond excess returns	-0.040	0.135	0.009	-0.105	0.963	9.177	0.101
	-2.877	2.062	3.258	-0.494	1.236	4.035	
(3) log dividend yield	-0.130	-0.149	0.976	1.109	-2.760	-13.301	0.978
	-1.557	-0.537	65.333	1.094	-0.746	-1.400	
(4) log inflation	0.003	0.003	0.000	0.187	0.387	0.113	0.070
	0.775	0.194	-0.161	4.494	1.963	0.232	
(5) log short rate	0.000	-0.001	0.000	0.002	0.984	0.003	0.992
	-0.056	-2.340	-2.425	0.749	120.299	0.216	
(6) log yield spread	0.000	-0.001	0.000	-0.002	0.003	0.905	0.899
	2.336	-1.172	-2.427	-0.824	0.363	34.864	

Panel B. Residual correlation matrix						
	(1)	(2)	(3)	(4)	(5)	(6)
(1) log stock excess returns	20.175	0.043	-0.855	0.045	-0.048	-0.016
(2) log bond excess returns	0.043	5.237	-0.045	0.036	-0.190	-0.720
(3) log dividend yield	-0.855	-0.045	23.529	-0.010	0.070	0.020
(4) log inflation	0.045	0.036	-0.010	1.252	0.058	-0.042
(5) log short rate	-0.048	-0.190	0.070	0.058	0.041	-0.408
(6) log yield spread	-0.016	-0.720	0.020	-0.042	-0.408	0.061

Table E7. VAR(1) Model Estimates [United Kingdom]

Panel A. Model estimates							
	Coefficients on lagged variables						
	(1)	(2)	(3)	(4)	(5)	(6)	Rsq
(1) log stock excess returns	0.058	0.035	0.031	0.051	-1.733	-0.927	0.023
	0.942	0.237	1.816	0.074	-0.815	-0.228	
(2) log bond excess returns	-0.032	0.127	0.002	-0.114	0.689	2.258	0.032
	-1.369	2.194	0.399	-0.367	1.008	1.750	
(3) log dividend yield	-0.051	-0.115	0.969	-0.152	1.335	-0.003	0.953
	-0.791	-0.691	51.994	-0.214	0.600	-0.001	
(4) log inflation	0.004	-0.026	0.001	0.156	0.210	-0.018	0.078
	0.960	-1.691	0.380	2.376	1.431	-0.058	
(5) log short rate	0.000	-0.004	0.000	0.005	1.015	0.048	0.994
	1.261	-3.927	-2.125	1.525	102.847	2.350	
(6) log yield spread	0.000	0.002	0.000	-0.006	-0.027	0.931	0.955
	-0.025	1.713	2.301	-1.150	-2.697	43.322	

Panel B. Residual correlation matrix						
	(1)	(2)	(3)	(4)	(5)	(6)
(1) log stock excess returns	15.937	0.150	-0.910	-0.030	-0.078	-0.092
(2) log bond excess returns	0.150	6.368	-0.133	-0.083	-0.257	-0.591
(3) log dividend yield	-0.910	-0.133	17.728	0.085	0.068	0.101
(4) log inflation	-0.030	-0.083	0.085	1.433	0.105	0.023
(5) log short rate	-0.078	-0.257	0.068	0.105	0.078	-0.585
(6) log yield spread	-0.092	-0.591	0.101	0.023	-0.585	0.099

Table E8. VAR(1) Model Estimates [United States]

Panel A. Model estimates							
	Coefficients on lagged variables						
	(1)	(2)	(3)	(4)	(5)	(6)	Rsq
(1) log stock excess returns	0.074	-0.021	0.025	0.064	-3.911	-7.121	0.026
	1.018	-0.151	2.442	0.081	-1.682	-1.683	
(2) log bond excess returns	-0.086	0.098	-0.004	-0.687	1.791	4.373	0.090
	-3.206	1.824	-1.175	-1.888	2.361	3.184	
(3) log dividend yield	-0.054	0.039	0.978	0.541	2.124	5.629	0.981
	-0.772	0.294	91.974	0.748	0.859	1.336	
(4) log inflation	0.009	-0.015	0.000	0.430	0.127	-0.098	0.246
	1.537	-1.698	0.446	5.186	0.929	-0.403	
(5) log short rate	0.001	-0.001	0.000	0.006	1.040	0.168	0.957
	0.746	-0.637	-3.003	0.978	80.709	5.540	
(6) log yield spread	0.000	0.000	0.000	0.002	-0.065	0.785	0.778
	0.380	0.028	2.944	0.260	-3.719	21.863	

Panel B. Residual correlation matrix						
	(1)	(2)	(3)	(4)	(5)	(6)
(1) log stock excess returns	15.531	-0.009	-0.955	-0.033	0.035	-0.036
(2) log bond excess returns	-0.009	6.148	-0.012	-0.133	0.056	-0.497
(3) log dividend yield	-0.955	-0.012	15.720	0.044	0.002	0.016
(4) log inflation	-0.033	-0.133	0.044	0.979	-0.026	0.076
(5) log short rate	0.035	0.056	0.002	-0.026	0.140	-0.889
(6) log yield spread	-0.036	-0.497	0.016	0.076	-0.889	0.166

Appendix F. Fisher Transformation and Correlation Contribution

F.1 Fisher Transformation

- Test hypothesis about population correlation coefficient ρ between X and Y using the sample correlation coefficient r.
- Define $z = \frac{1}{2}\ln\left(\frac{1+r}{1-r}\right)$. If (X, Y) is bivariate normal, and if (X_i, Y_i) used to form r are independent, then $z \sim \mathcal{N}\left(\frac{1}{2}\ln\left(\frac{1+\rho}{1-\rho}\right), \frac{1}{N-3}\right)$ where N is the sample size.
- For two samples of data, the early sample (1) and the late sample (2), define $z_1 = \frac{1}{2}\ln\left(\frac{1+r_1}{1-r_1}\right)$ and $z_2 = \frac{1}{2}\ln\left(\frac{1+r_2}{1-r_2}\right)$. The difference is $z_1 - z_2 \sim \mathcal{N}\left(\frac{1}{2}\ln\left(\frac{1+\rho_1}{1-\rho_1}\right) - \frac{1}{2}\ln\left(\frac{1+\rho_2}{1-\rho_2}\right), \frac{1}{N_1-3} + \frac{1}{N_2-3}\right)$. p-values can then be obtained in the normal way.

F.2 Correlation Contribution

- Stocks: $\tilde{xr}_{t+1} = N_{CF,t+1} - N_{RR,t+1} - N_{RP,t+1}$.
- Bonds: $\tilde{xr}_{t+1} = -N_{CF,t+1} - N_{RR,t+1} - N_{RP,t+1}$. An increase in $N_{CF,t+1}$ here is interpreted as inflation.
- The reported "Component Correlations" in Tables 4-6 look at correlations of the above components with no signs.
- The reported "Component Contributions" in Tables 4-6 look at how much of the average correlation in excess returns is being explained by covariances of news components. E.g., in Table 3, the bonds/bonds across countries pairwise correlation in the first sample period is $\sum_{i,j} \text{Corr}\left(xr_i, xr_j\right)$. Rewrite this as $\sum_{i,j} \frac{\mathbb{C}[xr_i, xr_j]}{\sigma(xr_i)\sigma(xr_j)}$. The numerator can be broken into 9 covariances of news components that contribute to the average correlation in excess returns. Note that because the news components come from innovations in excess returns as opposed to actual excess returns, the component contributions don't sum up exactly to the average correlation in excess returns.

Appendix G. Semidefinite Programming Method

We do a constrained minimization problem to estimate the covariance matrices which satisfy two constraints: A). volatility matrix and within-country correlation are the same across two sample period. B). covariance matrix is positive semi-definite. First we decompose a covariance matrix into volatility matrix and correlation matrix

$$\Sigma = D\Gamma D = \begin{pmatrix} \sigma_1 & \cdots & 0 \\ \vdots & \ddots & \vdots \\ 0 & \cdots & \sigma_m \end{pmatrix} \begin{pmatrix} 1 & \cdots & \rho_{1m} \\ \vdots & \ddots & \vdots \\ \rho_{1m} & \cdots & 1 \end{pmatrix} \begin{pmatrix} \sigma_1 & \cdots & 0 \\ \vdots & \ddots & \vdots \\ 0 & \cdots & \sigma_m \end{pmatrix}$$

Where the σ_i and ρ_{ij} $(i, j = 1, ..., m)$ are the coefficients to be estimated. Suppose $\widehat{\Sigma_1}$ and $\widehat{\Sigma_2}$ are the sample covariance matrices for early period and late period (known), then we need to estimate two covariance matrix $\Sigma_1 = D_1\Gamma_1 D_1$ and $\Sigma_2 = D_2\Gamma_2 D_2$ with the constraint $D_1 = D_2 = D$ and $\Gamma_1^{within} = \Gamma_2^{within}$. We use the minimum distance estimation, and this is a well defined constrained optimization problem

$$\min_{\Sigma_1,\Sigma_2} \left\{ \| \widehat{\Sigma_1} - \Sigma_1 \|_2 + \| \widehat{\Sigma_2} - \Sigma_2 \|_2 \right\}$$

$$\iff \min_{D,\Gamma_1,\Gamma_2} \left\{ \| \widehat{\Sigma_1} - D\Gamma_1 D \|_2 + \| \widehat{\Sigma_2} - D\Gamma_2 D \|_2 \right\}$$

$$s.t.\ \Gamma_i \succcurlyeq 0\ (i = 1, 2)$$

$$\Gamma_2^{within} = \Gamma_1^{within}$$

where $\| \,.\, \|_2$ represents the norm in L^2 space ($\| A - B \|_2 = \sum_{i,j}(a_{ij} - b_{ij})^2$), the notation $\Gamma \succcurlyeq 0$ means the matrix Γ is positive semi-definite, and Γ^{within} denotes the within-country correlation. To solve the Semidefinite programming (SDP) problem, we use the MATLAB package CVX by Stephen Boyd. `http://cvxr.com/cvx/doc/sdp.html`