Currency Risk Factors in a Recursive Multicountry Economy

R. Colacito **M. M. Croce** **F. Gavazzoni** **R. Ready***

Abstract

Focusing on the ten most traded currencies, we provide empirical evidence regarding a significant heterogeneous exposure to global growth news shocks. We incorporate this empirical fact in a frictionless risk-sharing model with recursive preferences, multiple countries, and multiple consumption goods whose supply features both global and local short- and long-run shocks. Since news shocks are priced, heterogenous exposure to long-lasting global growth shocks results in a relevant reallocation of international resources and currency adjustments. Our unified framework replicates the properties of the HML-FX and HML-NFA carry-trade strategies studied by Lustig et al. (2011) and Della Corte et al. (2013).

JEL classification: C62; F31; G12.
This draft: May 14, 2017.

*Colacito and Croce are affiliated with the Kenan-Flagler Business School at the University of North Carolina–Chapel Hill. Gavazzoni is affiliated with INSEAD. Ready is affiliated with the Simon School at the University of Rochester. We thank Bernard Dumas, Xavier Gabaix, Tarek Hassan, and Adrien Verdelhan for early feedback. We also thank our discussants Philippe Bacchetta, Jie Cao, Emmanuel Farhi, Pierre-Olivier Gourinchas, Christian Heyerdahl-Larsen, Matteo Maggiori, Thomas Maurer, and Pablo Ottonello, along with seminar participants at INSEAD, the NBER SI (IAP group), the meetings of the Econometric Society, the annual meeting of the European Finance Association, the Macro-Finance Society Conference, BI Norwegian Business School, the annual meetings of the American Economic Association, the annual meeting of the American Finance Association, the annual meeting of the Midwest Finance Association, Bocconi University, the annual meeting of the Society for Financial Econometrics, and the Fourth International Finance Conference in Hong Kong for useful comments and insights. All errors remain our own.

1 Introduction

A growing empirical literature in international finance examines the structure of risk in the cross section of currency returns (see, among others, Lustig and Verdelhan 2007, Lustig et al. 2014, and Della Corte et al. 2013). These studies sort currencies on various criteria and highlight the empirical relevance of several economic and financial factors. In this paper we provide novel empirical evidence in support of a structural dynamic equilibrium model that can account for these factor structures in the context of a news shocks–based asset-pricing model. According to our findings, heterogenous exposure to global growth news shocks is a key driver of currency riskiness, interest rates, and international lending positions in the cross section of countries.

Specifically, we analyze an economy populated by multiple countries engaged in a frictionless recursive risk-sharing scheme, in the spirit of Colacito and Croce (2013). This model features long-run growth news shocks, which are directly priced by Epstein and Zin (1989) recursive preferences. This setting is of particular interest given its documented ability to account for several empirical regularities of the joint dynamics of international asset prices and quantities in a two-country setting (see Colacito 2008, Colacito and Croce 2011, and Bansal and Shaliastovich 2013). We expand and generalize this setting in at least two relevant directions.

In a first step, we show that the ability of the Colacito and Croce (2013) model to replicate the failure of the uncovered interest parity is not sufficient to produce a risk premium in the cross section of interest rate–sorted currencies. That is, the mean of the Lustig et al. (2011) HML-FX factor is close to zero in a model in which countries have the same endowment exposure to global news shocks.

In a second step, we introduce heterogenous exposure to growth news shocks in the cross section of countries in a way that is consistent with our novel empirical evidence on the ten countries with the most traded currencies (henceforth G-10 countries). Specifically, we model a very persistent stochastic heterogeneity in the exposure of country-level endowments to long-run global growth news. The long-run wealth distribution in this economy is well defined, since we are still adopting a symmetric calibration (see Colacito et al. 2017). In finite samples, however, our countries feature substantial heterogeneity, consistent with the empirical investigation by Lustig et al. (2011) and Hassan and Mano (2014). These heterogenous loadings are a reduced-form way of capturing a mix of fundamental differences across countries, such as size (Hassan 2013), commodity intensity (Ready et al. 2012), monetary policy rules (Backus et al. 2010), and financial development (Maggiori 2011).

Under our benchmark calibration, we are able to produce an average HML annual spread of about 3%, as large as the unconditional HML-FX in the data. This currency risk premium originates from a positive correlation between the returns to carry trade and global long-run consumption news. When a negative long-run shock hits, the carry trade yields a negative return, due to the appreciation of the funding currencies (i.e., high-exposure countries). In good times, the carry trade earns a positive return due to the appreciation of the investment currencies (i.e., low-exposure countries).

We also show that this setting can replicate the empirical distribution of currency-portfolio betas on the HML factor. In addition, we document that, in our model, sorting countries on interest rates is equivalent to sorting on net foreign asset (henceforth NFA) positions and exposure to long-run global growth news. This suggests that the factors proposed by Della Corte et al. (2013) and Lustig et al. (2011) may be the risk-sharing outcome of a single fundamental source of heterogeneity, that is, their different exposure to global long-run growth news.

To discipline our calibration, we use macroeconomic data from the ten most traded currencies in the world. In our cross section of countries, the price-dividend (p/d) ratio is a statistically significant predictor of future growth rates of output. For this reason, in each country we use the projection of the GDP growth rate onto lagged values of the p/d ratio as our measure of long-run growth. We denote the innovations to this estimated component as long-run growth news shocks and show that they have a sizeable impact on countries' future growth prospects. We use the residual of our predictive regressions as a measure of short-run growth shocks. In a second step, we regress each country's long-run risk on the cross-sectional average long-run risk in our cross section of countries and view the estimated coefficients as a measure of each country's exposure to global long-run risk. We find a substantial degree of heterogeneity in the estimated exposures.

In particular, countries like Australia and New Zealand, which are commonly featured in the long leg of carry-trade strategies, have very low exposures to global long-run risk. In contrast, countries like Switzerland, which represents a typical funding currency in the carry trade, feature a substantially higher degree of exposure to global long-run risk. Interestingly, we cannot identify any heterogeneity in terms of the exposure to global short-run risk, implying that abstracting away from growth news shocks would prevent us from identifying a key driver of the international heterogeneity.

Our theoretical explanation of the cross section of currency risk premia produces a rich set of novel testable predictions. We document that sorting the average risk-free rates, the volatility of exchange rates, and the first two moments of NFA positions with respect to the exposure to global long-run risks produces the same patterns in the model and in the data.

In addition, we explore the time-series relation between the cross section of exposures to global long-run risks and international capital flows. According to our model, when a negative global shock materializes, the countries that are less exposed to global shocks should provide insurance to the countries with a large degree of exposure to global shocks. That is, countries with lower exposure to global long-run news should experience a deterioration of their NFA positions, whereas highly exposed countries should retain a larger balance of net foreign assets.

We provide empirical support for this prediction by showing that countries like Australia and New Zealand experience a large drop (increase) in their NFA positions in times of negative (positive) long-run global growth prospects. Countries on the other end of the spectrum, such as Switzerland, experience a large inflow (outflow) of assets from abroad whenever global long-run growth prospects are weak (strong). Furthermore, the currencies of high-exposure countries have the tendency to appreciate in bad times relative to the currencies of low-exposure countries, a finding which is also consistent with our model.

We push our analysis one step further and estimate the preference parameters using Euler equation restrictions on currency returns, equity returns, and risk-free rates; each country's average budget constraint (which depends on its NFA positions); and the average volatility of each country's exchange rate as target moments in a GMM estimation exercise. We obtain estimates for the preference parameters which are in line with our benchmark calibration, and we cannot reject the model in any of the cases that we consider. We also find that we can strongly reject the model with constant relative risk aversion (CRRA) preferences.

Our analysis helps to shed light on the connection between currency risk and country-level characteristics related to international trade. On the one hand, we provide equilibrium foundations for the reduced-form analysis of Lustig et al. (2011), Zviadadze (2013) and Farhi et al. (2015). On the other hand, we reconcile currency risk factors with macroeconomic fundamentals by directly analyzing the role of international asset positions (Gourinchas and Rey 2007, Caballero et al. 2008, and Gourio et al. 2014). Furthermore, our benchmark model with heterogenous exposure to growth news is consistent with Verdelhan (2015), as it enables global long-run shocks to contribute to bilateral exchange rate variance.

Our study is also related to the growing body of literature that has investigated the macroeconomic foundations of international financial market fluctuations (see, inter alia, Farhi and Gabaix 2008, Hassan 2013, Stathopoulos 2012, Heyerdahl-Larsen 2015, and Verdelhan 2010). Our empirical evidence on the 2007 Great Recession is related to the work on rare disasters (Barro 2006, Gabaix 2012, and Gourio 2012) and its applications to international finance (see, for example, Gourio et al. 2014).

Several articles have documented limitations of the long-run risks model in a one-country setting (see, for example, Le and Singleton 2010 and Beeler and Campbell 2012). In our analysis, we document that while our complete-markets framework goes a long way in accounting for the international dynamics of asset prices and quantities, it does not fully replicate the cross section. The introduction of frictions (see, for example, Gabaix and Maggiori 2015, Froot and Stein 1991, and Farhi and Werning 2014) may resolve these limitations, and we regard this as an important direction for future research.

In the next section we report our empirical evidence concerning the heterogenous exposure to global long-run risk in G-10 countries. In section 3 we present our risk-sharing model and our equilibrium conditions. Section 4 presents our main results, and section 5 concludes.

2 Empirical Analysis

Data sources. For our empirical work we focus on the set of G-10 currency countries: Australia, Canada, Germany, Japan, Norway, New Zealand, Sweden, Switzerland, the United Kingdom, and the United States. These countries have highly developed economies with high levels of global financial integration and highly liquid currencies, yet they still provide a rich set of cross-sectional empirical differences. We therefore view them as an appropriate setting in which to test the predictions of our model, which features heterogenous exposure to global risk in the presence of complete financial markets.

Data on consumption and net exports are from the Organisation for Economic Cooperation and Development (OECD). We use the volume index of the private final consumption expenditure series as our consumption series, and the difference between the volume indices of exports of goods and services and imports of goods and services as our net export series. In what follows, the variable ΔGDP_t proxies for total endowment growth and is computed as the growth in consumption plus net exports for the entire cross section of countries that we use in our empirical investigation. We exclude investment and government expenditure from our empirical measure to be consistent with our endowment economy in which we abstract away from both physical investment and demand for public goods. Data on net foreign assets come from the updated and extended version of the data set constructed by Lane and Milesi-Ferretti (2007).[1]

Data on interest rates and inflation are also from the OECD. Real interest rates are constructed using 3-month interbank rates from the OECD, adjusting for realized inflation using the annual changes in CPI. Norwegian interest rates prior to 1979 are

[1]Data graciously provided on Philip Lane's website: http://www.philiplane.org/EWN.html.

obtained from Stats Norway. Swedish 3-month rates prior to 1982 are extrapolated from overnight rates reported by the OECD and long-term rates reported by Statistics Sweden. International p/d ratios are calculated using Kenneth French's cum- and ex-dividend country value-weighted dollar index returns.[2]

All exchange rate data are collected relative to the US dollar. Exchange rate data for Australia, Canada, Japan, New Zealand, Norway, Sweden, Switzerland, and the UK are obtained from the Board of Governors of the Federal Reserve. Exchange rate data for Germany are obtained from the OECD. The sample period is 1971–2014. Real exchange rates are obtained by dividing by the relative CPI index of the corresponding country.

Estimation procedure. We follow Colacito and Croce (2013) and Bansal et al. (2010) in identifying short- and long-run innovations to GDP growth rates by means of predictive regressions. Specifically, we interpret the projection of ΔGDP_t in each country onto lagged values of that country's p/d ratio as our measure of long-run risk. A balanced panel of p/d ratios starts in 1987, so we use this sample to estimate global risk exposure. The estimation of all the parameters of interest is obtained from the following specification:

$$\begin{aligned} \Delta GDP_t^i &= \phi \cdot pd_{t-1}^i + \sigma \cdot \underbrace{\varepsilon_t^i}_{\text{Short-Run Shock}} \\ z_t^i &= \rho_z \cdot z_{t-1}^i + \varphi_e \cdot \sigma \cdot \underbrace{\varepsilon_{z,t}^i}_{\text{Long-Run Shock}}, \end{aligned} \tag{1}$$

$\forall i \in$ {G-10 countries} and where $z_t^i = \phi \cdot pd_t^i$ is our measure of the long-run risk in country i. We omit the intercepts because all variables are demeaned. Furthermore, we pool the estimation of the parameters $(\phi, \sigma, \rho_z, \varphi_e)$ for parsimony.[3] The system

[2] http://mba.tuck.dartmouth.edu/pages/faculty/ken.french/data_library.html

[3] We have adopted this pooled specification because we cannot reject the null hypothesis that all the country-specific ϕ coefficients are equal to each other.

in (1) is estimated using GMM, and standard errors take into account both serial and cross-sectional correlation.

We report the results of our estimation in panel A of table 1. Several interesting findings emerge. First, all estimated parameters are statistically significant, suggesting the relevance of lagged p/d ratios as a source of predictability for future growth rates of macroeconomic variables. Second, the projection of the GDP growth rate on the lagged p/d ratio is highly persistent, with an annual autocorrelation on the order of 0.77. This implies that shocks to the p/d ratio are likely to have long-lasting effects on the future dynamics of the growth rate of GDP. Third, the shocks to the predictive component of the GDP growth rate are small compared to the magnitude of the short-run shocks, that is, the parameter φ_e is on the order of 0.06. All of these features of the data discipline the calibration of the model that we present in the next sections.

Exposure to global risk. In panel B of table 1 we document the exposure of each country's endowment to a measure of global endowment risk. Specifically, we regress each country's GDP growth rate, ΔGDP_t^i, on the average of all countries' GDP growth rates:

$$\Delta GDP_t^i \quad = \quad \left(1+\beta_{\Delta y}^i\right) \cdot \left(\frac{1}{n}\sum_{i=1}^{n} \Delta GDP_t^i\right) + \xi_t^i, \quad \forall i \in \{\text{G-10 countries}\}. \tag{2}$$

With the sole exception of Norway and Canada, we cannot reject the null hypothesis that all the estimated coefficients in panel B of table 1 are equal to zero. Equivalently, looking at global endowment growth does not generate any significant form of heterogeneity in the cross section.

This conclusion is completely reversed when we focus on the exposure to global long-run news. Specifically, we measure the loading of each country's long-run risk on a global measure of long-run risk by regressing each z_t^i estimated in (1) on the average

Table 1: Dynamics of Endowments and Predictive Components

Panel A: Estimation of Predictive Components

	ϕ	ρ_x	σ	φ_e
Parameters	0.005^{***}	0.773^{***}	0.020^{***}	0.058^{***}
(S.E.)	(0.000)	(0.006)	(0.000)	(0.001)

Panel B: Exposure to Global Endowment Risk

	NZ	**AUS**	**UK**	**GER**	**CAN**	**NOR**	**JPN**	**SUI**	**US**	**SWE**
$\beta^i_{\Delta y}$	-0.28	-0.18	0.05	-0.12	0.14^{*}	0.61^{**}	0.15	-0.11	-0.11	-0.16
(S.E.)	(0.299)	(0.234)	(0.164)	(0.218)	(0.085)	(0.269)	(0.269)	(0.177)	(0.104)	(0.199)

Panel C: Exposure to Global Long-Run Risk

	NZ	**AUS**	**UK**	**GER**	**CAN**	**NOR**	**JPN**	**SUI**	**US**	**SWE**
β^i	-0.51^{***}	-0.44^{***}	-0.08	-0.02	0.00	0.08	0.12	0.26^{**}	0.27^{*}	0.33^{**}
(S.E.)	(0.154)	(0.064)	(0.098)	(0.094)	(0.131)	(0.173)	(0.165)	(0.130)	(0.166)	(0.148)
Chow	[0.109]	[0.245]	[0.299]	[0.841]	[0.729]	[0.506]	[0.802]	[0.667]	[0.596]	[0.385]

Notes - Panel A reports the estimation of the parameters of the endowment process reported in equation (1). The parameters are estimated using the longest available sample for each country, as described in section 2. Panel B reports the estimation of the exposure of each country's GDP growth rate to the global GDP growth rate (see equation (2)). The sample is 1970–2013. Panel C reports the estimation of the exposure of each country's predictive component of GDP to the global predictive component (see equation (3)). The sample is 1987–2013. The numbers in square brackets are the p-values associated with the null hypthesis that the estimated exposure in the first half of the sample (1987–2000) is different from the estimated exposure in the second half of the sample (2001–2013). In all panels, the numbers in parentheses are heteroskedasticity-adjusted standard errors. One, two, and three stars denote statistical significance at the 10%, 5%, and 1% levels, respectively.

of all z^i_t's:

$$z^i_t \quad = \quad \left(1+\beta^i\right) \cdot \left(\frac{1}{n}\sum_{i=1}^{n} z^i_t\right) + \zeta^i_t, \quad \forall i \in \{\text{G-10 countries}\}. \tag{3}$$

We estimate these coefficients using the longest common sample across all countries, 1987–2013, and report our results in panel C of table 1. We note that there is a sizeable degree of dispersion in the estimated β^i's, which range from -0.51 for New Zealand to 0.33 for Sweden. Furthermore, the estimated coefficients are statistically different from zero for the most extreme countries. In panel C, we also report the p-values associated with a Chow test for the null hypothesis that the β^i's are statistically different in the first and second halves of our sample. We always reject the null at conventional confidence levels.

According to our analysis, the US is a country with relatively high exposure to global long-run risk. This finding does not contradict the view of Lustig and Verdelhan (2007) that the US is an average country in the cross-sectional distribution of real interest rates. Specifically, the US had a relatively high real rate in the first part of the post-Bretton-Woods era (see panel (a) of figure G2 in the appendix), but not in the more recent sample period that we consider.

Equivalently, an econometrician focusing on the entire sample would conclude that the US has been on average a median interest-rate country, whereas someone looking at the post-1987 sample would conclude that the US is a below-median country.

Similar considerations apply to Japan. Table 1 indicates that Japan is not the country with the highest exposure to global long-run risk, a finding which may seem at odds with the idea that the Japanese yen should be a funding currency in the carry trade. We note that the real interest rate in Japan was extremely low until the early 2000s, but it has since sharply increased in the cross section of interest rates observed among our G-10 countries (see panel (b) of figure G2).

Our empirical findings could, in principle, be generated by different models. In the next sections, we show that modeling the presence of persistent and stable heterogeneity in the exposure to global long-run risk across countries combined with recursive preferences reconciles several features of the cross section of real interest rates, equity premia, and currency risks. We formally test this crucial interaction in the context of our GMM exercise in section 4.4 and document the ability of this model to match several key moments in the cross section of international asset prices and macroeconomic quantities.

Additional results. The p/d ratios that we use in our empirical exercises do not take into account equity buy-backs and new issuances, which are common policies in the US and other major industrialized countries. Given the difficulty of identifying a reliable dataset on total nonfinancial corporate distributions for our large cross section of countries, we provide additional evidence using Tobin's Q, a forward-looking variable which does not require the use of dividends. In appendix A, we document in detail that using Tobin's Q to forecast long-run risk confirms our results about heterogenous exposure to global growth news shocks across countries. Furthermore, we show that our results are robust to: (i) starting our sample in 1975, (ii) using short- and long-run shocks, and (iii) using the portion of the p/d ratios that is orthogonal to time-varying risk premia.

3 The Economy

The economy consists of N countries, and N goods, $\{X_i\}_{i=1}^N$. Agents' preferences are defined over consumption aggregates of the N goods as follows.

Consumption aggregate. Let $x_{i,t}^j$ denote the consumption of good j in country i at date t. The consumption aggregates in the N countries are

$$C_t^i = (x_{i,t}^i)^\alpha \prod_{j \neq i} (x_{j,t}^i)^{\frac{1-\alpha}{N-1}}. \quad (4)$$

The parameter $\alpha \in (0,1)$ captures the degree of bias of the consumption of each representative agent. In what follows we assume that each country i receives a stochastic endowment of good $X_{i,t}$ at each point in time. Following some of the international macrofinance articles surveyed by Lewis (2011), we assume that α is larger than 0.5. This allows us to build consumption home bias into the model.

Preferences. As in Epstein and Zin (1989), agents' preferences are recursive but non–time separable:

$$U_{i,t} = \left[(1-\delta) \cdot (C_{i,t})^{1-1/\psi} + \delta E_t \left[(U_{i,t+1})^{1-\gamma} \right]^{\frac{1-1/\psi}{1-\gamma}} \right]^{\frac{1}{1-1/\psi}}, \quad \forall i \in \{1, \ldots, N\}. \tag{5}$$

The coefficients γ and ψ measure the relative risk aversion (RRA) and the intertemporal elasticity of substitution (IES), respectively.

The main departure from the constant RRA case often analyzed in the literature lies in the fact that our model's preferences allow agents to be risk averse in future utility as well as future consumption. The extent of such utility risk aversion depends on the preference for early resolution of uncertainty measured by $\gamma - 1/\psi > 0$. To better highlight this feature of the preferences, we note that the ordinally equivalent transformation

$$V_t = \frac{U_t^{1-1/\psi}}{1-1/\psi}$$

can be approximated as follows:

$$\begin{aligned} V_t &= (1-\delta)\frac{C_t^{1-1/\psi}}{1-1/\psi} + \delta E_t \left[V_{t+1}^{1-\theta} \right]^{\frac{1}{1-\theta}} \\ &\approx (1-\delta)\frac{C_t^{1-1/\psi}}{1-1/\psi} + \delta E_t \left[V_{t+1} \right] - \frac{\delta}{2} \frac{\theta}{E_t \left[V_{t+1} \right]} Var_t \left[V_{t+1} \right], \end{aligned} \tag{6}$$

where $\theta \equiv \frac{\gamma - 1/\psi}{1-1/\psi}$. Note that the sign of $\left(\frac{\theta}{E_t[V_{t+1}]} \right)$ depends on the sign of $(\gamma - 1/\psi)$. When $\gamma = 1/\psi$, the agent is utility-risk neutral and preferences collapse to the standard time-additive case. When the agent prefers early resolution of uncertainty, that is, when $\gamma > 1/\psi$, the coefficient θ is positive: uncertainty about continuation utility reduces welfare and generates an incentive to trade off future expected utility, $E_t \left[V_{t+1} \right]$, for future utility risk, $Var_t \left[V_{t+1} \right]$. This mean-variance trade-off is an appealing feature of these preferences, and one that is absent when agents have standard time-additive preferences. This trade-off drives international allocations and exchange rate adjustments in our economy, and it represents the most important element of our analysis. Our study is the first to fully characterize trade with Epstein and Zin (1989) preferences in an economy composed of an arbitrary number of countries.

Since there is a one-to-one mapping between utility, $U_{i,t}$, and lifetime wealth, the optimal risk-sharing scheme can also be interpreted in terms of mean-variance tradeoff of wealth. For this reason, in what follows we will use the terms "wealth" and "continuation utility" interchangeably.

Endowments. We choose to endow each country with a stochastic supply of its most-preferred good. Endowments are cointegrated processes, and they feature predictive variables as follows:

$$\log X_t^i = \mu_x + \log X_{t-1}^i + z_{i,t-1} - \tau \left[\log X_{t-1}^i - \frac{1}{N} \log \left(\sum_{j=1}^{N} X_{i,t} \right) \right] + \varepsilon_{i,t}^X, \tag{7}$$

where $\tau \in (0,1)$ determines the extent of cointegration, and the processes z_i are modeled as highly persistent AR(1) processes,

$$z_{i,t} = \rho_i z_{i,t-1} + \varepsilon_{i,t}^z, \forall i \in \{1, 2, \ldots, N\}. \tag{8}$$

Throughout the paper, we refer to $\varepsilon_{i,t}^z$ as the long-run shocks, due to their long-lasting impact on the growth rates of the endowments. Similarly, we call $\varepsilon_{i,t}^X$ short-run shocks. Shocks are jointly log-normal. We abstract from exogenous time-varying volatility in endowments to better quantify the amount of endogenous consumption and asset-price volatility generated by our recursive risk-sharing mechanism with complete markets.

Heterogenous exposure. Consistent with our empirical investigation, we introduce cross-country variation in the exposure to global long-run endowment shocks, $\beta^z_{i,t}$. Since long-run shocks are the primary driver of our risk-sharing mechanism, this modeling choice produces no loss of generality. Specifically, in each country we decompose our long-run shocks ($\varepsilon^z_{i,t}$) into a common global component and a country-specific component as follows:

$$\varepsilon^z_{i,t} \quad = \quad (1 + \beta^z_{i,t-1})\epsilon^z_{global,t} + \epsilon^{\tilde{z}}_{i,t}, \tag{9}$$

with the shocks to the two components being orthogonal to each other:

$$\textbf{corr}(\epsilon^z_{global,t}, \epsilon^{\tilde{z}}_{i,t}) = \textbf{corr}(\epsilon^{\tilde{z}}_{i,t}, \epsilon^{\tilde{z}}_{f,t}) = 0.$$

The volatilities of $\epsilon^z_{global,t}$ and $\epsilon^{\tilde{z}}_{i,t}$ are set to replicate both the unconditional standard deviation and correlation of the long-run shocks, $\epsilon^z_{i,t}$, described in the previous section. Country-specific sensitivity coefficients are modeled as a slowly moving AR(1) process,

$$\beta^z_{i,t} = \rho^\beta_z \beta^z_{i,t-1} + \epsilon^{\beta,z}_{i,t}$$

with $\epsilon^{\beta,z}_{i,t} \sim i.i.d.N(0, \sigma_{\beta,z})$. These shocks are both long-lived ($\rho^\beta_z \approx 1$) and uncorrelated to other shocks, as they are meant to approximate nearly unconditional differences in the exposure of countries to global news. Countries with high $\beta^z_{i,t}$ have relatively riskier endowments, in the sense that their local growth processes are more exposed to shocks to global long-run growth.

Our way to model country-specific exposure to shocks produces a twofold benefit. First of all, it enables us to study an economy with ex ante symmetrically calibrated countries for which a well-defined equilibrium exists (Colacito et al. 2017). Under the assumption of a permanent degree of heterogeneous exposure, there would be a degenerate ergodic distribution of the equilibrium variables.[4] Second, after simulating a history of heterogenous exposure shocks ($\epsilon_{i,t}^{\beta,z}$), we are able to study the characteristics of a cross section of countries that remain substantially heterogenous in finite samples. We think of the $\beta_{i,t}^{z}$ coefficients as devices to capture the heterogeneity documented by Lustig et al. (2011), Backus et al. (2010), Ready et al. (2012), and Hassan and Mano (2014) in a parsimonious, reduced-form manner.

Market structure. At each date, agents trade a complete set of state-contingent one-period-ahead claims to the numeraire good. Without loss of generality, we choose good 1 as our numeraire. Since both financial and goods markets are assumed to be frictionless, the budget constraint of our agents can be written as

$$\sum_{j=1}^{N} p_{j,t} x_{i,t}^{j} + \int_{\zeta^{t+1}} A_{i,t+1}\left(\zeta^{t+1}\right) Q_{t+1}(\zeta^{t+1}) \quad = \quad A_{i,t} + p_{i,t} X_{i,t}, \tag{10}$$

where $p_{i,t}$ denotes the price of good i relative to that of good 1, $A_{i,t}\left(\zeta^{t}\right)$ denotes country i's claims to time t consumption of good X_1, and $Q_{t+1}(\zeta^{t+1})$ gives the price of one unit of time $t+1$ consumption of good X_1 contingent on the realization of ζ^{t+1} at time $t+1$. In equilibrium, the market for international state-contingent claims clears, implying that $\sum_i A_{i,t} = 0, \forall t$.

Allocations. Since markets are complete, we can compute efficient allocations by solving the associated Pareto problem. The planner attaches date 0 nonnegative Pareto weights $\{\mu_i\}_{i=1}^{N}$ to the consumers and chooses the sequence of allocations $\{x_{i,t}^{j}\}_{t=0}^{+\infty}$, $\forall i$ and $j \in \{1, ..., N\}$ to maximize

$$\Lambda = \sum_{i=1}^{N} \mu_i \cdot U_{i,0},$$

[4] This outcome is very common in multiple-agent economies with fixed degrees of heterogeneity. For the case of survivorship with heterogeneous risk aversion, see Anderson (2005).

subject to the following sequence of economy-wide feasibility constraints:

$$\sum_{j=1}^{N} x_{i,t}^{j} = X_{i,t}, \quad \forall t \geq 0 \quad and \quad \forall i \in \{1, ..., N\},$$

where the state-dependent notation is omitted for the sake of clarity. In characterizing the equilibrium, we follow Anderson (2005) and Colacito and Croce (2013) and formulate the problem using the ratio of time-varying pseudo-Pareto weights, $S_{j,t} = \mu_{j,t}/\mu_{1,t}$, as an additional state variable. This technique enables us to take into account the nonseparability of the utility functions. We show in the appendix that the first-order necessary conditions imply the following allocations:

$$\begin{aligned} x_{i,t}^{i} &= \left(1 + \frac{1-\alpha}{\alpha(N-1)} \sum_{j \neq i} \frac{S_{j,t}}{S_{i,t}}\right)^{-1} X_{i,t}, \quad \forall i \in \{1, 2, ..., N\} \\ x_{i,t}^{j} &= \frac{1-\alpha}{\alpha} \frac{1}{N-1} \frac{S_{j,t}}{S_{i,t}} x_{i,t}^{i}, \quad \forall i \neq j \in \{1, 2, ..., N\} \end{aligned} \tag{11}$$

where

$$S_{j,t} = S_{j,t-1} \cdot \frac{M_{j,t}}{M_{1,t}} \cdot \left(\frac{C_{j,t}/C_{j,t-1}}{C_{1,t}/C_{1,t-1}}\right), \quad \forall t \geq 1 \tag{12}$$

and $S_{j,0} = 1$, as we start the economy from an identical allocation of wealth and endowments. This is consistent with the ergodic distribution of the model, which implies that on average all countries consume an identical share of world resources because of symmetry.

Prices. The stochastic discount factor (SDF) that is used to discount future uncertain payoffs is

$$M_{i,t+1} = \delta \left(\frac{C_{i,t+1}}{C_{i,t}}\right)^{-\frac{1}{\psi}} \left(\frac{U_{i,t+1}^{1-\gamma}}{E_t\left[U_{i,t+1}^{1-\gamma}\right]}\right)^{\frac{1/\psi-\gamma}{1-\gamma}}. \tag{13}$$

Since markets are assumed to be complete, the log growth rate of the real exchange rate between the consumption bundles of countries i and j is

$$\Delta e_{i,t}^{j} = \log M_{j,t} - \log M_{i,t}, \tag{14}$$

and the relative price of good j and good 1 is $p_{j,t} = \frac{(1-\alpha)}{\alpha(N-1)} \frac{x_{1,t}^{1}}{x_{j,t}^{1}}$.

Bilateral imports and exports. At each point in time, the exports of country 1 toward country j are equal to $EXP_{1,t}^{j} = x_{j,t}^{1}$, and the imports of country 1 from country j are equal to $IMP_{1,t}^{j} = p_{j,t} x_{1,t}^{j}$, where $x_{j,t}^{1}$, $p_{j,t}$, and $x_{1,t}^{j}$ are defined above. It follows that the bilateral volume of trade and the bilateral net exports rescaled by total output are equal to

$$\frac{Vol_{1,t}^{j}}{X_{1,t}} = \frac{(1-\alpha)\cdot(1+S_t^j)}{\alpha(N-1)+(1-\alpha)\sum_{j\neq 1} S_t^j} \tag{15}$$

and

$$\frac{NX_{1,t}^{j}}{X_{1,t}} = \frac{(1-\alpha)\cdot(S_t^j-1)}{\alpha(N-1)+(1-\alpha)\sum_{j\neq 1} S_t^j}, \tag{16}$$

respectively. Detailed derivations are reported in the appendix.

3.1 Calibration and solution method

We detail our baseline monthly calibration in table 2. We choose a monthly frequency to be consistent with the empirical methods adopted in the carry-trade literature. When possible, our parameters are chosen to be the monthly counterpart of our annual estimates.

All preference parameters are set in the spirit of the long-run risk literature (see Bansal and Yaron 2004, Bansal et al. 2010, and Colacito and Croce 2013). We set the home bias parameter (α) to 0.98. This number is within the empirical range of the import share for the countries that we consider in our analysis (between 2% and 28%). In appendix E we report a sensitivity analysis with respect to this parameter. The parameters governing the dynamics of the growth rates of the endowments are chosen to reflect an average annual growth rate of 2%, an unconditional volatility of 2%, and a modest degree of autocorrelation. These moments are set to be consistent with the median values in our data set and are discussed further in the next section.

We choose to calibrate the cross-country correlation of the short-run shocks to a low level, so that the model is consistent with the moderate cross-country correlation of consumption growth rates across major industrialized countries. Under the benchmark calibration with heterogenous exposure, this correlation is set to 0.40, a number that falls in the middle of the correlation range estimated among our ten countries. When we consider the special setting without heterogenous exposure, we set this parameter to a slightly smaller number, 0.20, to prevent the consumption profiles from being excessively correlated. Both parameter values are consistent with those of other studies in the international macroeconomics literature.

We set the cross-country correlation of the long-run components to 0.93, as the growth rates of major countries are highly correlated over the long horizon (see, among others, Colacito and Croce 2011). This finding also holds in our data, as our results show that the correlation of our long-run components can be as high as 0.95.[5] Both the amount of long-run risk (σ_z/σ) and the persistence of the long-run components (ρ) are consistent with the empirical estimates reported in table 1.

We set $\rho_z^\beta = .999$ to create nearly permanent heterogeneity in exposure to world output shocks. We choose $\sigma_{\beta,z} = 0.05\%$ to obtain a very moderate conditional volatility of

[5]The median correlation across all of the 45 pairwise correlations of p/d ratios for G-10 countries is about 0.8, and the 95% range of all the pairwise correlations is from 0.4 to 0.95.

our exposure parameters. Consistent with our empirical results, in the short sample the time variation of our exposure coefficients is statistically undetectable. In the same spirit, we set τ so as to have an almost undetectable extent of cointegration, consistent with Rabanal et al. (2011).[6] We present a sensitivity analysis with respect to this parameter in appendix E.

Given these parameters, we use perturbation methods to solve our system of equations. We compute an approximation of the third order of our policy functions using the dynare++ package. As documented in Colacito et al. (2017); Colacito and Croce (2013), a third-order approximation is required to capture endogenous time-varying volatility due to the adjustments of the pseudo-Pareto weights. We document the accuracy of our solution method in appendix F and show that it features very small approximation errors. All variables included in our dynare++ code are expressed in log-units.

4 The Cross Section of Currency Risk Premia

In this section, we study the main moments produced by a recursive risk-sharing scheme with multiple countries. We start from a setting with homogeneous exposure to news shocks and show that the resulting allocation dynamics are broadly consistent with the data. Most importantly, we show that a recursive risk-sharing scheme cannot produce the observed returns of a carry-trade strategy, even though it endogenously accounts for the forward-premium anomaly. This limitation is resolved in section 4.2 by introducing heterogenous exposure to global long-run shocks. In section 4.3, we show novel empirical evidence that supports the distinctive channels at work in our model.

[6] Cointegration of the endowments is needed to ensure the model's stationarity or, equivalently, balanced growth across countries.

TABLE 2: Calibration

Description	Values
Preferences:	
Relative Risk Aversion [γ]	6.50
Intertemporal Elasticity of Substitution [ψ]	1.60
Subjective Discount Factor [δ^{12}]	0.98
Degree of Home Bias [α]	0.98
Endowments:	
Mean of Endowment Growth (%) [12μ]	2.00
Short-Run Risk Volatility (%) [$\sigma\sqrt{12}$]	1.87
Long-Run Risk Volatility (%) [σ_z/σ]	6.00
Long-Run Risk Autocorrelation (%) [ρ^{12}]	0.78
Cross-Correlations of Long-Run Shocks [ρ_z]	0.93
Cointegration Speed [$(1-\tau)^{12}$]	0.99

	Exposure	
	Homogeneous	Heterogenous
Cross-Correlations of Short-Run Shocks [ρ_X]	0.20	0.40
Orthogonalization:		
Volatility of Global Long-Run Shocks (%) [σ_z^{global}/σ]	–	5.80
Volatility of Local Long-Run Shocks (%) [σ_z^{global}/σ]	–	1.60
Time-Varying Exposure:		
Autocorrelation of $\beta_{i,t}^z$ [$(\rho_z^\beta)^{12}$]	–	0.99
Volatility of Shocks to $\beta_{i,t}^z$ (%) [$\sigma_{\beta,z}$]	–	0.05

Notes - This table reports our benchmark monthly calibration. Under the homogeneous exposure case, $\beta_{i,t}^z = 1 \quad \forall i, \forall t$.

4.1 Homogeneous exposure

We set the number of countries in our model to be five, a figure large enough to form a proper cross section of currency portfolios and small enough to keep our computations feasible. We report the resulting main moments usually studied in international macrofinance in table 3. In appendix D, we show that most of these moments are not sensitive to the number of countries, implying that our results are quite general. We also characterize the behavior of common measures of risk-sharing as we increase the number of countries. Since most of our findings replicate those in the two-country economy of Colacito and Croce (2013), we defer the inspection of the mechanism of the model with homogeneous exposure to appendix D. In what follows we briefly highlight the successes and limitations of this setting.

Successes. With homogeneous exposure to global news shocks, our model produces consumption dynamics very close to the data. This is true for both within-country and cross-country moments. Thanks to the presence of highly cross-country-correlated long-run growth news, stochastic discount factors are volatile and highly correlated across countries even though consumption growth rates are not. As a result, the exchange rate growth volatility is not subject to the Brandt et al. (2006) puzzle.

This setting produces low and smooth risk-free rates. Across countries, the risk-free rates are as highly correlated as the stochastic discount factors. Furthermore, the volatility of the NFA position in each country is consistent with the data, since recursive preferences and long-run growth news make the valuation channel as strong as in the data (see Gourinchas and Rey 2013 and Colacito and Croce 2013). In appendix D, we show that the model can also account for the almost complete lack of correlation between consumption growth differentials and exchange rate fluctuations (cf. the Backus and Smith 1993 puzzle). This is the result of the opposite response of consumption growth differentials to short- and long-run shocks already documented in Colacito and Croce (2013).

As in Colacito and Croce (2013), the model is able to produce a negative slope of the uncovered interest rate parity regressions (β_{UIP}). In the presence of local news shocks, agents with recursive preferences are willing to swap current consumption for smoother future consumption profiles. As a result, agents implement a trade of securities that produces sizable endogenous time-varying volatility.

Limitations. We conclude our quantitative analysis by noting the inability of the model to produce a sizeable risk premium in the cross section of risk-free rate–sorted currencies ($E[HML]$). In a model with recursive preferences, the distribution of wealth or, equivalently, pseudo-Pareto weights ($S_{i,t}$), is an endogenous state variable that drives allocations and prices. Because of risk aversion and home bias, the optimal policy is nonlinear and the response to news shocks changes with the distribution of wealth.

In figure 1, we show the response to global shocks of countries that have the same endowment but different wealth. At the equilibrium, this is possible because of the arrival of different local news shocks. Specifically, country 1 is assumed to have received better local long-run news than country 5. Because of higher expected growth, country 1 has a higher interest rate. Because of risk sharing, country 1 must transfer resources to country 5, and thus country 1 has a negative NFA position. The strategy that goes long in the currency of country 1 (high interest rate country) and shorts the currency of country 5 (low interest rate country) produces a negligible carry-trade discount of -0.02%, a number well below the empirical premium of 3%.

The model also produces a very modest and smooth spread of the risk-free rates across countries. Since our exchange rate is more volatile than in the data, our model-implied β_{UIP} is excessively negative.

4.2 Heterogeneous exposure

In this section, we show that accounting for persistent heterogeneity in the exposure to world news shocks can produce sizeable cross-sectional currency premia. Furthermore, our model produces equivalent results when sorting countries on (i) nearly permanent heterogenous exposure to endowment shocks (Lustig et al. 2011, Hassan

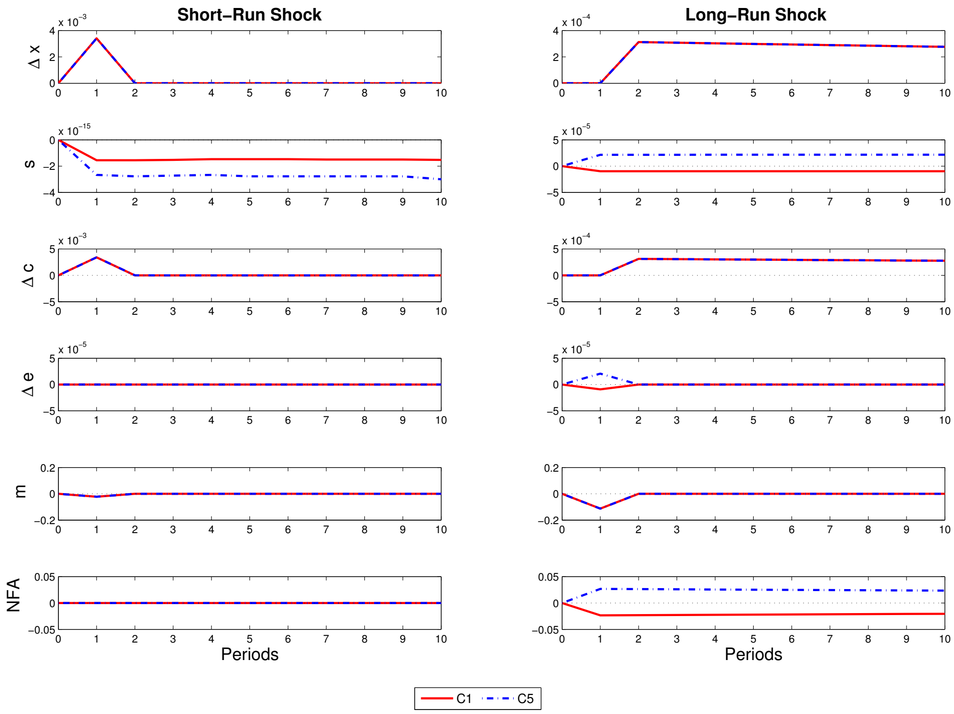


FIG. 1 - Impulse response functions under heterogenous wealth. The left (right) panels report the response of endowment growth ($\Delta \log X^i$), relative Pareto weights with respect to country 3 ($\log S_j/S_3$), consumption growth (Δc_i), exchange rate growth (Δe_3^i), stochastic discount factors (m_i), and NFA (A_i/X_i) to a one-standard-deviation short-run (long-run) global shock. All panels refer to the case in which the economy consists of 5 countries ($i = 1, ..., 5$). The exchange rate is measured with respect to country 3, implying that $\Delta e < 0$ for country 1 denotes a depreciation of its real exchange rate with respect to country 3. Country 1 (5) is initialized with positive (negative) NFA. Exposure to global news shocks is assumed to be homogeneous.

and Mano 2014), (ii) NFA positions (Della Corte et al. 2013), or (iii) the level of their risk-free rate (Lustig et al. 2011).

Conditional heterogeneity. Under the null of the model, the coefficients β^i in equation (3) are equivalent to those featured in equation (9). We introduce time variation in the exposure of our five countries and initialize our economy so that country 1 has an exposure of 0.65 and country 5 has an exposure of -0.65, a plausible range given our results from panel B of table 1. All other countries have an exposure coefficient equally spaced in this range.[7]

In order to reach this cross section of exposure values, we create a burn-in simulation period in which we give a sequence of positive (negative) exposure shocks to countries 1 and 2 (4 and 5), and no exposure shock to country 3. All results presented in this section are obtained from repetitions of small samples with a number of monthly periods consistent with the empirical data set. All repetitions start from the same initial cross section of exposure coefficients.

In our setting, all countries are risk averse and buy insurance against shocks that increase their exposure to long-run world growth news. Along our burn-in simulation path, country 1 is by construction the country that experiences the most adverse sequence of exposure shocks. As a result of its financial portfolio allocation strategy, this country accumulates wealth against the rest of the world, that is, it acquires a positive NFA position. Furthermore, high risk exposure induces stronger precautionary saving motives, and hence it lowers the interest rate. As a result, in our simulated samples country 1 is a net lender with a low risk-free rate, whereas country 5 is a net borrower with a high interest rate (Della Corte et al. 2013). From a qualitative point of view, no additional financial frictions are required to obtain this sorting.

The positive link between average NFA positions and exposure to the long-run news shocks can also be explained by solving forward the budget constraint in (10):

$$\begin{aligned} A_{i,t} &= \left[\sum_{j=1}^{N} p_{j,t} x_{i,t}^{j} - p_{i,t} X_{i,t}\right] + \int_{\zeta^{t+1}} Q_{t+1}\left(\zeta^{t+1}\right) \cdot A_{i,t+1}\left(\zeta^{t+1}\right) \\ &= -\sum_{k=0}^{\infty} \mathcal{M}_{t+k} \cdot NX_{i,t+k} \end{aligned}$$

[7] In appendix A, we show that estimating the β^i coefficients using long-run risk innovations, as opposed to levels, yields similar results. In appendix G, we show that when we apply our estimation procedure to simulated data, we can reproduce the cross section of the calibrated β^i coefficients.

where $\mathcal{M}_{t+k} = \prod_{l=0}^{k} Q_{t+l}$ and $NX_{i,t+k} = \sum_{j=1}^{N} p_{j,t} x_{i,t}^{j} - p_{i,t} X_{i,t}$, $\forall i \in \{1, ..., N\}$. The NFA level of a country tracks the present value of future promised transfers. In computing this expected present value, agents with a preference for early resolution of uncertainty assign stronger weight to states of the world in which negative global news shocks are realized. That is, negative global news shocks are the main determinant of the NFA positions. The trade balance of a low-exposure country improves in response to a negative global long-run shock: because of risk sharing, it must transfer resources to countries with high exposure. This means that countries with low exposure are expected to have predominantly positive net exports in the future. As a result, their average NFA position is negative. By the same argument, high-exposure countries must have positive NFA on average.

In what follows, we think of countries 1 and 5 as Switzerland and Australia, respectively. These countries are representative of the two legs of the carry trade. By construction, country 3 represents the median country.

Response to global shocks. In figure 2 we show the impulse response of our variables of interest for countries 1 and 5 with respect to global shocks. These impulse responses are created after our burn-in simulation sample, that is, at time zero country 1 has an exposure of +0.65, whereas country 5 has an exposure of -0.65.

With respect to global short-run shocks, both the consumption growth rate and stochastic discount factor adjust within each country as they would in a one-country economy. Since the exposure to global short-run shocks is the same for all countries, these shocks produce no international reallocation of resources. As a result, exchange rates do not move, and hence they feature zero exposure to global short-run risk. On this dimension the model is similar to many canonical models of exchange rates featuring symmetric countries and local shocks (e.g., Obstfeld and Rogoff 1995). As in the pre-

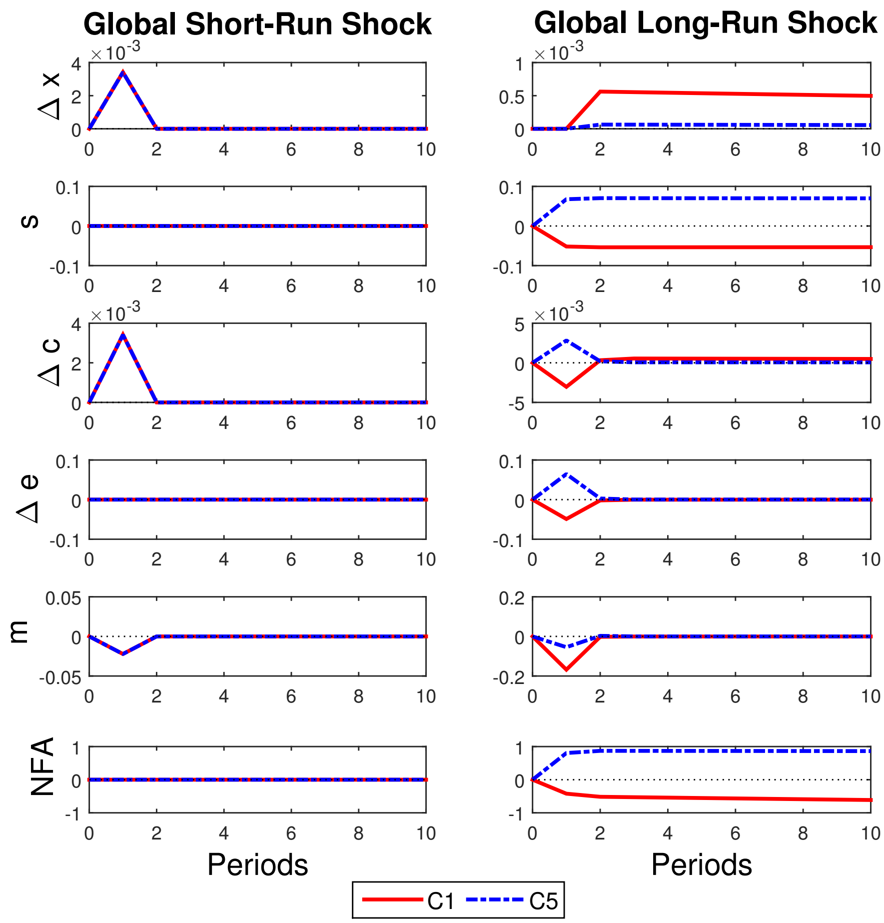


FIG. 2 - Impulse response functions under heterogenous exposure. The left (right) panels report the response of endowment growth ($\Delta \log X^i$), relative Pareto weights with respect to country 3 ($\log S_j/S_3$), consumption growth (Δc_i), exchange rate growth (Δe_3^i), stochastic discount factors (m_i), and NFA (A_i/X_i) to a one-standard-deviation short-run (long-run) global shock. All panels refer to the case in which the economy consists of 5 countries ($i = 1, ..., 5$). The exchange rate is measured with respect to country 3, implying that $\Delta e < 0$ for country 1 denotes a depreciation of its real exchange rate with respect to country 3. Country 1 (5) is initialized with an exposure to long-run shocks of 0.65 (-0.65).

vious section, short-run global risk prescribes a null average return to the carry-trade strategy.

In contrast to short-run shocks, global long-run news shocks promote relevant international reallocations and generate currency risk. Specifically, upon the arrival of good news, the most exposed country (country 1) gets better relative news and features the lowest marginal utility. This adjustment causes two effects. First, by risk sharing, the NFA position of country 1 deteriorates. In the next section, we show that the direction of these responses is consistent with that observed in our data set. Second, by no arbitrage, the real exchange rate of country 1 immediately depreciates, whereas the opposite is true for country 5. Since all countries experience a drop in their marginal utility, the HML-FX carry-trade strategy (going long in country 5 and short in country 1) is risky, as it pays high excess returns in good states of the world.

To better explain this result, in figure 3 we depict key characteristics of our countries. The top panel shows the country-specific exposure to global output shocks, as defined in equation (2). Consistent with our empirical findings, our countries have similar exposure to global output growth shocks, since by construction their exposure to short-run shocks is constant in the cross section. The second panel shows our distribution of long-run risk exposure coefficients $(1 + \beta_i^z)$, which is our main driver of cross-sectional heterogeneity.

The remaining panels refer to four currency portfolios created as follows. Without loss of generality, we select the median country (country 3) as numeraire and focus on the remaining four bilateral exchange rates to form four currency portfolios sorted on interest-rate differentials, $i_{j,t} - i_{3,t}$ with $j \neq 3$. The portfolio formation allows for frequent rebalancing, but due to the persistent differences in exposure to global long-run risk the transition of countries across different portfolios is infrequent. That is, portfolio 4 almost always corresponds to country 5, whereas portfolio 1 coincides with the riskiest country, that is, country 1.

As already pointed out, because of precautionary saving motives there is an inverse relationship between a country's exposure to global long-run news and the average level of its own risk-free rate (panel 3 of figure 3). The high endowment-beta countries have low currency betas, that is, their currencies depreciate in global good times, consistent with the impulse response of the country 1 exchange rate (panel 4). This

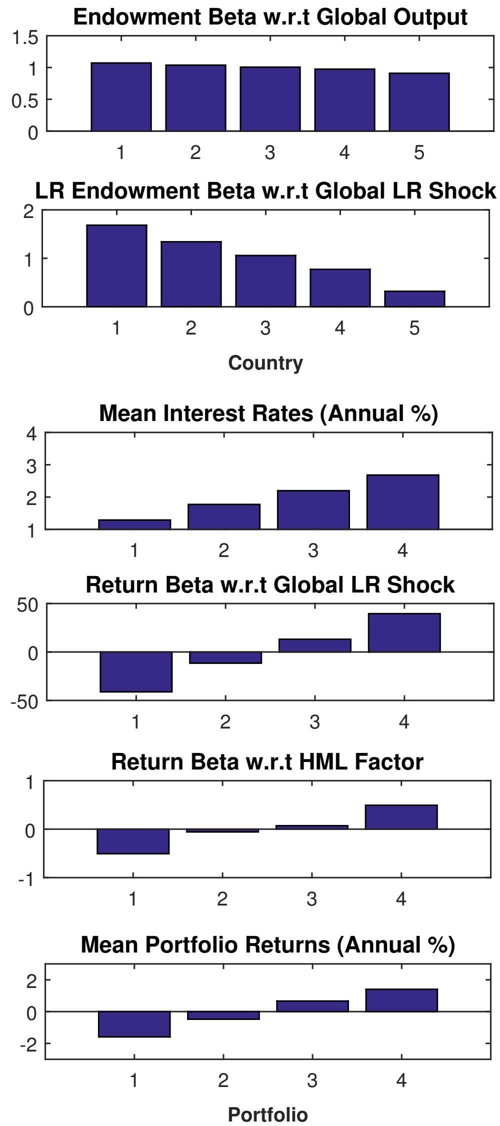


FIG. 3 - Cross-sectional risk (EZ case). The top panel shows the the exposure of our five countries to global output ($1+\beta^i_{\Delta y}$) as defined in equation (2). The second panel depicts their exposure to the global long-run component ($1+\beta^z_i$). The bottom four panels show the cross-sectional characteristics of four portfolios formed by sorting our countries each period on their lagged interest rate differential with respect to a fixed numeraire country (country 3). All the parameters are calibrated to the monthly values reported in table 2. Statistics are the averages across 300 simulations of 250 monthly observations.

mechanism is sufficient to generate a cross section of loadings on currency returns to the HML-FX factor that is consistent with the data (panel 5). The implied average currency returns have an annual spread of about 3% (bottom panel), consistent with the estimated unconditional HML-FX of Lustig et al. (2011). Given the absence of country turnover across portfolios, the unconditional HML-FX is the appropriate empirical counterpart of our currency risk premium.

The role of heterogeneity for median moments. In table 3, we report simulated moments from our model with and without heterogenous exposure and recursive preferences. In all cases, we report median values from our cross section of 5 countries. In the absence of heterogenous exposure (first column), the moments are identical to those obtained in the last column of table D7.

All moments remain basically unchanged across columns, implying that the introduction of heterogenous exposure does not undermine the basic successes of our model. We highlight just three differences. First, this configuration of the model is able to replicate both the forward premium anomaly and the unconditional HML observed in the data. As in the case of homogeneous exposure, however, our β_{UIP} is excessively negative.

Second, heterogeneity introduces more cross-country variation in the exposure to shocks and hence reduces the correlation of the stochastic discount factors. As a result, the volatility of the bilateral exchange rate of the median country is slightly higher than that obtained under homogeneous exposure and larger than in the data. We consider this limitation as an important challenge for future research.

Lastly, heterogeneity increases the risk-sharing motives and makes the adjustment of both the NFA positions and the net exports more severe. This explains why with recursive preferences the volatility of these variables more than doubles. We regard this result as a success, because in standard models with time-additive preferences these variables are excessively smooth, and especially so the net exports.

TABLE 3: Simulated Moments with Heterogenous Exposure

	Data	S.E.	Homog. EZ	Heterog. EZ	Heterog. CRRA
Std(Δx)	2.10	0.26	1.93	1.95	1.95
ACF$_1$(Δx)	0.21	0.13	0.29	0.30	0.35
corr(Δx_t^h,Δx_t^f)	0.23	0.06	0.43	0.40	0.40
Single-Country Moments					
Std(Δc)	1.91	0.25	1.78	1.96	1.74
ACF$_1$ (Δc)	0.46	0.11	0.31	0.28	0.30
Bilateral Moments					
corr(Δc_t^h,Δc_t^f)	0.24	0.05	0.55	0.38	0.59
Std (Δe)	9.10	0.91	14.65	17.01	10.07
corr(m, m^f)			0.94	0.85	0.59
Std(NX/X)	5.12	0.74	0.47	1.48	1.00
ACF$_1$(NX/X)	0.92	0.06	0.86	0.90	0.94
Financial Variables					
E(r_f)	2.16	0.74	2.26	2.13	11.79
Std[r_f]	2.88	0.41	1.04	1.14	11.74
corr(r_f^h, r_f^f)	0.57	0.05	0.92	0.71	0.89
Std(NFA/X)/Std(Δx)	18.58	2.95	11.34	25.76	10.29
ACF$_1$(NFA/X)	0.99	0.05	0.81	0.88	0.74
β_{UIP}	-0.94	0.48	-5.54	-4.62	0.78
E(HML)	3.20	1.10	0.11	3.01	0.13
Cross Sectional Standard Deviation					
Std(Δc)	0.45	0.12	0.06	0.21	0.09
E(r_f)	1.27	0.26	0.18	0.54	2.86
Std(r_f) (CoV)	0.42	0.08	0.03	0.46	0.34
Std(NFA/X)/Std(Δx) (CoV)	0.55	0.09	0.01	0.68	0.74
Std(NX/X) (CoV)	0.52	0.09	0.02	0.61	0.48
β_{UIP} (CoV)	0.87	0.29	1.30	1.16	0.58
Std (Δe) (CoV)	0.21	0.04	0.03	0.41	0.04

Notes - The table reports both empirical moments computed using the data set described in section 2 and simulated moments from the model with both heterogeneous and homogeneous exposure. All parameters are set to their benchmark values reported in table 2. For the CRRA case, we set $\gamma = 1/6.5$. The first panel reports the moments for the dynamics of exogenous endowment growth rates. The panel labeled “Single-Country Moments” reports the moments of the consumption growth rate within each country. The panel labeled “Bilateral Moments” reports the cross-country moments for each country pair. The panel labeled “Financial Variables” reports the median moments for the risk-free rates (r_f), stochastic discount factors (M), NFA-to-output (NFA), slope coefficient of the UIP regressions (β_{UIP}), and average currency risk premium (E[HML]). In the bottom panel, we report cross-sectional standard deviations for the listed moments. CoV denotes the cross-sectional coefficient of variation.

The role of heterogeneity for cross-sectional moments. In the bottom panel of table 3, we focus on key cross-sectional moments of both quantities and prices. First of all, our benchmark model is the only one producing a cross-sectional dispersion of consumption volatility sizable enough to be within our empirical confidence interval. This result supports our consumption-based approach.

As already mentioned, all models produce moderate time-series variation in the risk-free rates. This limitation carries over to the cross section, as the cross country volatility of the average risk-free rates is almost three times smaller than in the data. Equivalently, the spread in interest rates reported in figure 3 is not as large as in the data.

Given the discrepancies between the data and the median model-implied volatility for net exports, NFA position, exchange rates, interest rates, and β_{UIP}, we analyze whether the model can reproduce the coefficients of variation in our empirical cross section. Under homogeneous exposure, the result is negative. With heterogenous exposure, in contrast, the model produces figures well within our empirical confidence intervals. This outcome is reassuring, as it confirms that our multicountry model features a rich, although not perfect, characterization of key moments in a broad international cross section.

The role of preferences. Since in the next section we study in detail the cross-sectional properties of our model, we conclude this part of our analysis by showing the results for the special case in which we set IES = 1/RRA = 1/6.5, that is, the CRRA configuration. Figure 4 and the last column of table 3 document that a number of counterfactual results arise in this particular setup.

First, we compute measures of consumption exposure to both global output growth shocks and long-run shocks. Since long-run news shocks are not directly priced and

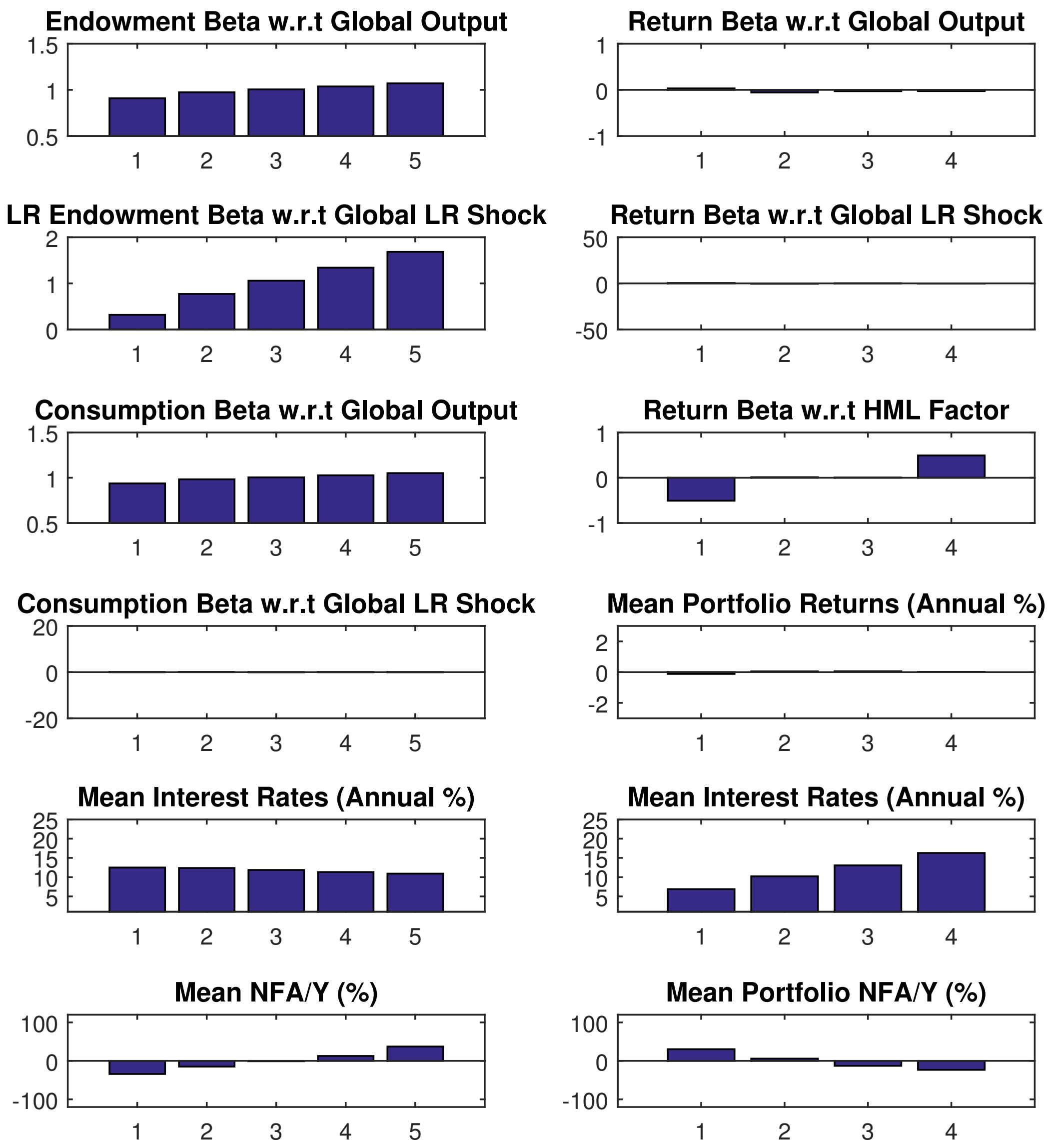


FIG. 4 - Cross-sectional risk (CRRA case). The panels on the left show the cross-sectional characteristics of five countries sorted on their exposure to the global long-run shock $(1+\beta_i^z)$. The panels on the right show the cross-sectional characteristics of four portfolios formed by sorting our countries each period on their lagged interest rate differential with respect to a fixed numeraire country. All the parameters are calibrated to the monthly values reported in table 2, except the IES, which is set to 1/6.5, the inverse of the risk aversion coefficient. Statistics are the averages across 100 simulations of short samples (360 monthly observations).

produce no immediate movements in the marginal utilities of our countries, there is no significant reallocation. As a result, the consumption growth betas with respect to long-run news are zero across all countries, even though their endowment growth exposures continue to be heterogenous.

Second, the amount of financial trade in the economy is much more limited than before, as documented by (i) the counterfactually moderate spread in the average NFA positions of our five countries (left column, bottom panel), and (ii) the reduced volatility of the NFA positions. Furthermore, with this particularly low value of the IES, the risk-free rates are too high (Weil 1989) and basically constant across countries (left column, second panel from the bottom). As a consequence, the average returns on currency portfolios sorted according to interest rate differentials have an irregular pattern. The same statements apply to the exposure of currency portfolio returns with respect to both global long-run shocks and the HML factor. While some of the cross-sectional coefficients of variation in table 3 are consistent with the data, the same cannot be stated for the corresponding median values.

To summarize, long-run global growth news shocks can be an important driver of multiple phenomena in the cross section of currency, provided that agents price them directly. Epstein and Zin (1989) preferences enable news shocks to be priced and generate a recursive risk-sharing scheme that can explain key features of trade and international asset prices both at a country-pair level and in the cross section of countries. In the next section, we provide evidence supporting a number of implications of our frictionless model. In several relevant dimensions, the quantitative performance of the model is surprisingly close to the data.

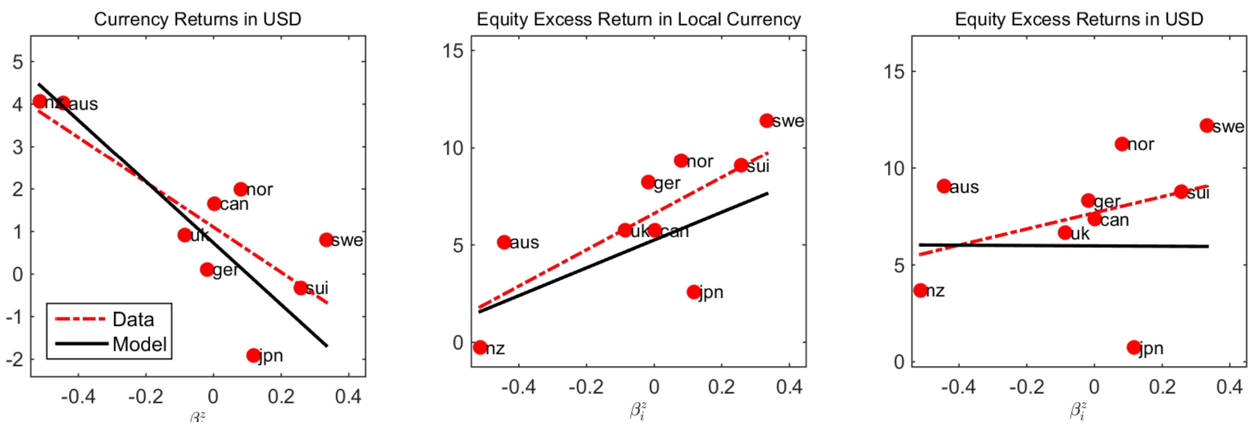


FIG. 5 - Equity and FX excess returns. In each panel, the horizontal axis refers to the estimated exposure coefficient β_i^z. For each country in our data set, we use the estimates provided in table 1. The left panel links exposure to global long-run risk to average currency returns expressed in USD. The middle (right) panel refers to average equity (excess) returns in each country in local (USD) units. In each panel, the dots correspond to the actual data, the dashed line provides a linear fit of the data, and the thick line represents the model's prediction. The model is calibrated as in table 2 and it features heterogenous exposure to global long-run growth shocks. The equity excess return in local units is defined as $r_{d,t}^{ex,i} = \lambda r_{c,t}^{ex,i} + \epsilon_t^i$, $i \in \{1, 2, ..., N\}$, where $\lambda = 6$, $r_{c,t}^{ex,i}$ is the excess return on the consumption claim, and $\epsilon_t^i \sim_{i.i.d.} N(0, 0.15^2)$ captures dividend-specific shocks.

4.3 Evidence of qualitative predictions of the model

In this section, we provide direct empirical evidence supporting the implications of our model for several aggregates of interest. In figure 5, we assess the ability of our model to replicate the cross section of currency and equity returns. In each panel, we report the data values for the cross section of countries that we employ in our empirical investigation (dots), a linear fit of the data (dashed lines), and the model's predictions (thick lines). Consistent with our previous simulation results, a carry-trade strategy based on differences in exposure to long-run growth news can produce a premium comparable to that observed in our data (left panel).

Furthermore, our model performs well in replicating the joint distribution of currency and equity risk across countries. As in the data, countries with higher exposure to global growth news have higher local excess returns (middle panel), whereas the cross section of equity returns from the perspective of a US-based investor is close to flat

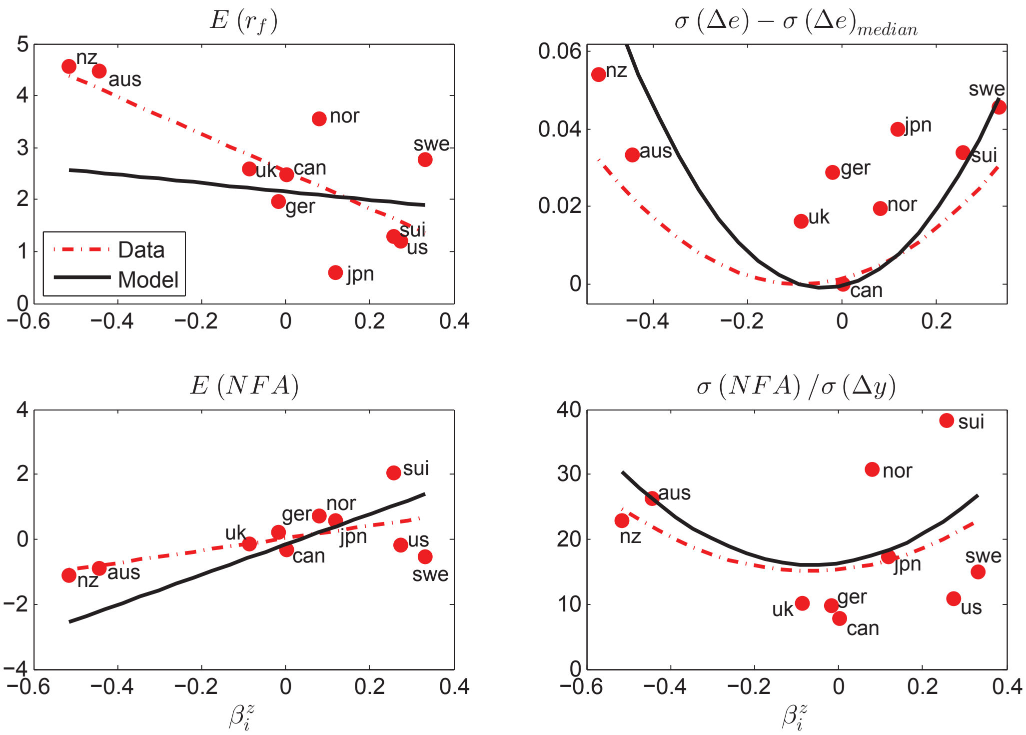


FIG. 6 - Cross-sectional patterns. In each panel, the horizontal axis refers to the estimated exposure coefficient β_i^z. For each country in our data set, we use the estimates provided in table 1. The top-left panel links exposure to global long-run risk to the average risk-free rate in the cross section of countries. The top-right panel refers to the volatility of the growth rate of the real exchange rate of each country against the US dollar. Volatilities are expressed in terms of deviation from the median country. The bottom-left (bottom-right) panel focuses on the average level (volatility) of the NFA to output ratio. In each panel, the dots correspond to the actual data, the dashed line provides a linear-quadratic fit of the data, and the thick line represents the model's prediction. The model is calibrated as in table 2 and features heterogenous exposure to global long-run growth shocks.

both in the data and in our model (right panel). Consistent with our risk-sharing mechanism, the exchange rate of countries with high exposure to global growth news provides a powerful hedge against equity risk.

The top left panel of figure 6 documents that average risk-free rates are decreasing in the degree of exposure to global long-run risk, consistent with the precautionary saving channel discussed in the previous section. As in our model, countries typically featured in the long legs of carry-trade strategies (e.g., Australia and New Zealand) feature both low exposure to global risk and a high average risk-free rate. In contrast, the countries commonly used as the funding currency (e.g., Switzerland) have a large

β and a low average interest rate.

By no arbitrage, the model suggests that the volatility of real exchange rate fluctuations is a function of the spread of the βs across countries. Thus, countries with an exposure close to the median beta should have smooth exchange rates, whereas countries with extreme betas should have highly volatile currencies (the "volatility smile"). The top right panel of figure 6 shows that this prediction finds strong support in the data, as the volatilities of the growth rate of the real exchange rate (vis-$\grave{a}$-vis the US dollar) are well approximated by a concave quadratic function of the exposure coefficient.

Our risk sharing–based model predicts that countries with low exposure to global risk provide insurance to countries with high exposure to global risk in international capital markets. The bottom left panel of figure 6 documents that our model's prediction lines up with the data, as countries like Australia and New Zealand (which have the lowest βs) have negative average NFA positions and countries like the US, Switzerland, and Sweden (which have the largest βs) have positive average NFA positions.

Additionally, the model predicts that countries with extreme βs should experience a larger extent of fluctuation in their NFA positions, as they engage in a substantial amount of trading of securities to provide and receive insurance against their exposure to global risk (the reallocation channel). In contrast, countries with βs close to the median should have relatively smoother NFA positions. As shown in the bottom right panel of figure 6, this prediction is confirmed in the data.

We further explore the mechanism of the model by analyzing the response of foreign assets to a global long-run shock at a country level in our cross section. For each country, we regress the time series of the NFA position on the level of the global long-run risk:

$$\frac{NFA_{i,t}}{GDP_{i,t}} = \alpha_i^{NFA} + \lambda_i^{NFA} \cdot z_{global,t} + \xi_{i,t}. \tag{17}$$

According to our model, countries that have low exposure to global long-run risk (i.e., low-β countries) should experience an outflow (inflow) of resources at the occurrence of a negative (positive) global long-run shock. Equivalently, the λ^{NFA} coefficient should be positive for countries with low β and negative for countries with high β. We test this negative link in our cross section of β-sorted countries by jointly estimating the following system of equations via GMM:

$$\frac{NFA_{i,t}}{GDP_{i,t}} = \alpha_i^{NFA} + \left(\vartheta_0^{NFA} + \vartheta_1^{NFA} \cdot \beta_i^z\right) \cdot z_{global,t} + \xi_{i,t} \quad i = 1, 2, ..., 10. \tag{18}$$

Under the null hypothesis of the model, the estimated ϑ_1^{NFA} in equation (18) should be negative.

We illustrate our results in figure 7. Countries with low β, such as Australia and New Zealand, have positive estimated λ^{NFA} coefficients in equation (17), whereas countries with high β, such as Switzerland and Sweden, have negative estimated λ^{NFA} coefficients. For the countries with the most extreme exposure to global shocks, we can typically reject the null hypothesis that their estimated λ^{NFA} coefficient is equal to zero. Furthermore, the negative cross-sectional link between β and λ^{NFA} predicted by our model cannot be rejected in the data.

We perform a similar analysis to study the response of each country's exchange rate to a global long-run shock. In particular, for each country i we estimate the coefficients

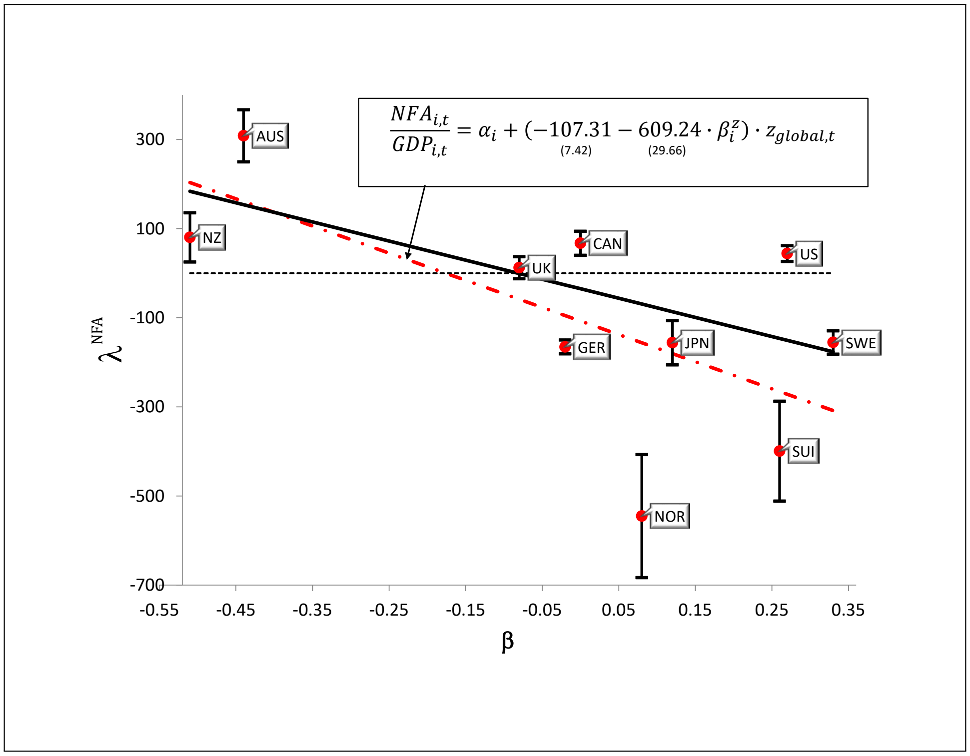


FIG. 7 - NFA exposure. Each dot represents the estimated sensitivity of the NFA over the GDP of a country with respect to global long-run risk (see equation (17), coefficient λ_i^{NFA}). For each dot, the vertical line represents the 90% confidence interval associated to the estimated coefficient. The dashed line corresponds to the point estimate of the line $\vartheta_0^{NFA} + \vartheta_1^{NFA} \cdot \beta_i^z$ in equation (18). The solid line represents the model prediction. The estimated βs are reported in table 1. Standard errors are adjusted for heteroskedasticity.

λ_i^{FX} in

$$\Delta e_{i,t} = \alpha_i^{FX} + \lambda_i^{FX} \cdot \Delta z_{global,t} + \xi_{i,t}, \tag{19}$$

and then estimate the slope coefficient ϑ_1^{FX} in the following system of equations:

$$\Delta e_{i,t} = \alpha_i^{FX} + \left(\vartheta_0^{FX} + \vartheta_1^{FX} \cdot \beta_i^z\right) \cdot \Delta z_{global,t} + \xi_{i,t} \quad i = 1, 2, ..., 9. \tag{20}$$

We depict the results of this part of the analysis in figure 8. Our model predicts that the currencies of countries with high exposure to global growth news (high β) experience stronger appreciations in response to negative global growth shocks. This is equivalent to a negative ϑ_1^{FX}. In our cross section, the estimate of this coefficient is

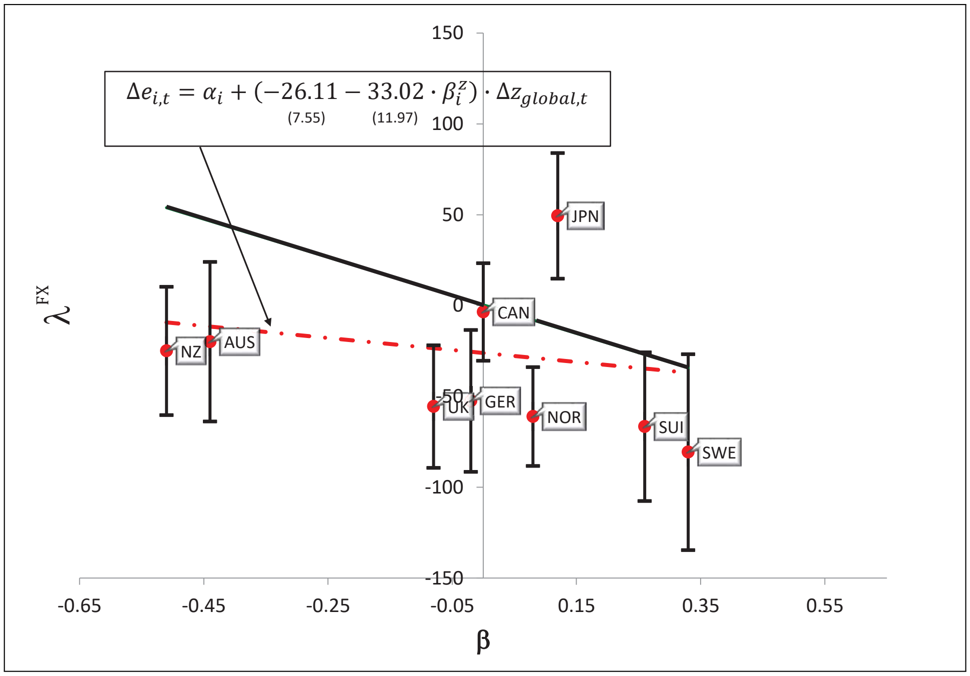


FIG. 8 - **Currency exposure.** Each dot represents the estimated sensitivity of the growth rate of the exchange rate of a country's currency relative to the US dollar with respect to a global long-run shock (see equation (19), coefficient λ_i^{FX}). For each dot, the vertical line represents the 90% confidence interval associated with the estimated coefficient. The dashed line corresponds to the point estimate of the line $\vartheta_0^{FX} + \vartheta_1^{FX} \cdot \beta_i^z$ in equation (20). The solid line represents the model prediction. The estimated βs are reported in table 1. Standard errors are adjusted for heterokedasticity.

negative and statistically significant, consistent with the model. Since in our economy we compute exchange rates with respect to the median country, the model-implied regression line is restricted to passing through the origin, which is in contrast to the data.[8]

We conclude this part of the analysis, by checking the robustness of our results to controlling for local growth news shocks and excluding specific countries. To streamline the presentation of the results, in table 4 we focus on portfolios constructed according to countries' βs, as opposed to focusing on each country separately.

[8]Japan is the only notable outlier in figure 8. This is due mostly to the strong depreciation of the Japanese yen during the first part of the 1990s (see Obstfeld 2010 for a detailed account). This period also coincides with the time in which the Japanese yen was less than prominently featured as the funding currency of carry-trade strategies. We document in the appendix that for the more recent 2007–2008 Great Recession, both the exchange rate and the NFA behaved exactly as predicted by the model (see Figure G3a and Figure G3b in the appendix.)

Specifically, we construct three portfolios based on the results of our analysis in table 1. The first portfolio (labeled "Low β_i^z") pools the estimates of λ_i^{FX} and λ_i^{NFA} for all the countries with βs that are negative and statistically different from zero (i.e., Australia and New Zealand). The second portfolio (labeled "Medium β_i^z") refers to the countries whose βs are not statistically different from zero (i.e., the UK, Germany, Canada, Norway, and Japan). The third portfolio (labeled "High β_i^z") consists of the countries with βs that are positive and statistically different from zero (i.e., the US, Switzerland, and Sweden).

Columns [1] and [5] of table 4 are the portfolio counterparts of the country-level estimated coefficients depicted in figure 7 and figure 8. Columns [2], [6], and [7] document that our results are robust to the exclusion of Japan from portfolios 2 and United States from portfolio 3.

In columns [3] and [4], we augment the specification in equation (19) by including the following two additional regressors: (i) the local long-run news for country i, and (ii) the local long-run news for the US, since the US is our base country. Our results are robust to the inclusion of these additional control variables. In columns [8], [9], and [10], we perform a similar exercise by including the local long-run risk component as an additional variable in equation (17). The results are virtually unchanged relative to columns [5], [6], and [7].

Overall, we regard the results presented in this section as strongly supporting our multicountry recursive risk-sharing mechanism. Our empirical findings confirm that the degree of heterogeneity in the exposure to global risk is well aligned with a number of quantities of interest in international financial markets. Furthermore, our general-equilibrium approach enables us to relate the no arbitrage–based hypothesis of Lustig et al. (2011) and Lustig et al. (2014) to macroeconomic fundamentals such as international consumption dynamics.

4.4 Risk-sharing and currency pricing using the model's SDF

In this section, we test whether the model's implied SDFs (i) can price the cross-section of currency returns, and (ii) are consistent with the cross sectional variation of exchange rates and NFA positions. We do this by using the time series of long- and short-run shocks estimated in section 2.

To carry out this empirical exercise, we first need a functional form of each country's SDF. In a single-country setup, this function is generally obtained via a log-linear approximation (e.g., Bansal et al. 2010). For our model, we do not have a closed-form approximation, so we proceed numerically via a second-order polynomial projection approach. Specifically, we first simulate the model over a grid of values of ψ and γ. We then use a projection to obtain an approximate functional form of each country's SDF in terms of ψ and γ and observable variables.

The observable variables are the risk-free rate ($r^i_{f,k,t}$), the exposure to the global shocks ($\beta^i_{k,t}$), the country-specific short-run shock ($\varepsilon^i_{k,t+1}$), the country-specific long run shock ($\varepsilon^i_{z,k,t+1}$), and the realizations of the global shocks ($\varepsilon^{global}_{k,t+1}$), where k indexes simulations, and t indexes time. Using these simulations, we obtain the fitted parameters that best approximate the equilibrium relationship between each country's SDF, preference parameters, and the model's shocks and observables. This regression yields a specification of each country's SDF:

$$m_{i,t+1} = F(\psi, \gamma, r^i_{f,t}, \beta_{i,t}, \varepsilon_{global,t+1}, \varepsilon^i_{t+1}, \varepsilon^i_{z,t+1}). \tag{21}$$

This approach provides a very accurate approximation of the equilibrium SDFs with R^2s in the order of 99%. We provide more details of this procedure in appendix F.

We feed the time series of the shocks estimated in section 2 into the stochastic dis-

TABLE 4: Conditional Analysis

	Exchange Rate				Net Foreign Assets					
Portfolio	[1]	[2]	[3]	[4]	[5]	[6]	[7]	[8]	[9]	[10]
Low β_i^z	-26.52	-26.52	-27.51	-27.51	384.03	384.03	384.03	437.92	437.92	437.92
	(16.05)	(16.05)	(13.91)	(13.91)	(55.67)	(55.67)	(55.67)	(129.36)	(129.36)	(129.36)
Medium β_i^z	-23.85	-38.49	-27.31	-52.15	-46.44	384.03	-15.74	33.44	33.44	7.23
	(6.08)	(9.35)	(12.06)	(12.72)	(2.95)	(55.67)	(1.95)	(7.79)	(7.79)	(7.36)
High β_i^z	-67.16	-67.16	-84.88	-84.88	-154.46	-149.21	-154.46	-141.02	-144.54	-141.02
	(23.75)	(23.75)	(20.43)	(20.43)	(33.57)	(24.20)	(33.57)	(22.72)	(22.69)	(22.72)
Include All	-	-	-	-	Y	N	N	Y	N	N
Exclude JPN	N	Y	N	Y	N	N	Y	N	N	Y
Exclude US	Y	Y	Y	Y	N	Y	N	N	Y	N
Control local shocks	N	N	Y	Y	N	N	N	Y	Y	Y

Notes - Unless otherwise specified in the last four rows of the Table, the "Low β_i^z" portfolio contains Australia, and New Zealand; the "Medium β_i^z" portfolio contains UK, Germany, Canada, Norway, and Japan; the "High β_i^z" portfolio contains Switzerland, US, and Sweden. The United States are excluded from the analysis in columns [1]-[4], because the US Dollar is the base currency for this part of the analysis. The coefficients reported in columns [1] and [2] are obtained by pooling the coefficients λ_i^{FX} in equation (19) for the countries of the corresponding portfolios. The coefficients in columns [3] and [4] repeat the same analysis, by augmenting equation (19) with the component of the domestic long-run shock that is orthogonal to the global long-run shock, i.e. $\Delta\left(z_{i,t} - \beta_i^z \cdot z_{global,t}\right)$, and the component of the US long-run shock that is orthogonal to the global long-run shock, i.e. $\Delta\left(z_{US,t} - \beta_{US}^z \cdot z_{global,t}\right)$. The coefficients reported in columns [5]-[7] are obtained by pooling the coefficients λ_i^{NFA} in equation (17) for the countries of the corresponding portfolios. The coefficients in columns [8]-[10] repeat the same analysis, by augmenting equation (17) with the non-global component of the local long-run risk, i.e. $z_{i,t} - \beta_i^z \cdot z_{global,t}$, as an additional variable. For all configurations, the sample is 1987-2013. All standard errors are adjusted for heteroskedasticity. The average exposures to the global long-run risk (β_i^z) are reported in Table G10 of the Appendix.

count factors reported in (21). Using GMM, we then estimate the elasticity of intertemporal substitution (ψ) and the risk aversion coefficient (γ) using the following moment restrictions:

1. The second moment of each country's exchange rate growth rate (labeled as "FX vol" in table 5)

$$\frac{1}{T}\sum_{t=1}^{T}\left(m_t^i - m_t^{US}\right)^2 - \left(\Delta e_t^i\right)^2 \quad = \quad 0, \tag{22}$$

2. Each country's NFA position (labeled "NFA" in table 5)

$$\frac{1}{T}\sum_{t=1}^{T}\exp\{m_{t+1}^i\}A_{t+1}^i - NX_t^i - A_t^i \quad = \quad 0, \tag{23}$$

3. Each country's Euler equation for pricing the local stock market return, $R_{s,t}$, and the risk-free asset, $R_{f,t}$ (labeled as "Market returns" and "Risk-free rates," respectively, in table 5)

$$\frac{1}{T}\sum_{t=1}^{T}\exp\{m_t^i\}R_{k,t}^i - 1 \quad = \quad 0, \forall k = s, f \tag{24}$$

4. The Euler equations for the pricing of the currency excess returns of the 6 HML and IMX portfolios ("HML (6 portf)" and "IMX (6 portf)" in table 5), the excess return of the sixth over the first HML and IMX portfolios ("HML (6-1)" and "IMX (6-1)" in table 5), and the excess return of a strategy that is short in the US and long in each of the remaining 9 currencies in the G10 set of countries ("RFX (G-10)" in table 5).

As a result, we have a total of up to 62 moment restrictions to estimate the ψ and the γ parameters in a cross section of 10 countries.

TABLE 5: Testing Moment Restrictions

	[1]	[2]	[3]	[4]	[5]	[6]	[7]	[8]	[9]
ψ	3.81	1.37	1.21	0.82	0.86	0.74	1.01	1.00	1.11
	(2.09)	(0.39)	(0.32)	(0.19)	(0.31)	(0.11)	(0.20)	(0.24)	(0.22)
γ	2.16	2.89	3.52	3.95	3.86	4.00	3.43	3.89	3.71
	(0.21)	(0.64)	(0.25)	(0.33)	(0.68)	(0.32)	(0.35)	(0.27)	(0.27)
$H_0 : \gamma = 1/\psi$	2.45	7.77	8.87	9.82	5.06	23.78	14.23	11.51	16.06
	[0.12]	[0.01]	[0.00]	[0.00]	[0.02]	[0.00]	[0.00]	[0.00]	[0.00]
J-stat (Full)	13.55	14.76	16.29	20.39	20.59	20.09	17.71	20.77	18.34
	[0.06]	[0.06]	[0.50]	[0.88]	[0.42]	[0.39]	[0.89]	[0.99]	[0.99]
J-stat (Pricing)				11.34	5.93	0.44	7.68	22	21.97
				[0.50]	[0.92]	[0.80]	[0.57]	[0.46]	[0.34]
FX Vol.	✓		✓	✓		✓	✓	✓	✓
NFA		✓	✓	✓	✓	✓	✓	✓	✓
RFX (G10)							✓		
HML (6 portf)				✓	✓				
HML (6-1)						✓		✓	
IMX (6 portf)				✓	✓				
IMX (6-1)						✓		✓	
Market Returns								✓	✓
Risk-Free Rates								✓	✓

Notes - This table reports the estimated preference parameters ψ and γ associated with 9 different combinations of moment restrictions. The numbers in parentheses are the standard errors of each estimated coefficient. The third row reports the Wald tests for the null hypothesis that $\gamma = 1/\psi$ with the associated p-value displayed in brackets underneath. The J-statistics reported in the second panel of the table refer to the full set of moment restrictions (labeled "J-stat (Full)") and to the subset of moment restrictions that is limited to asset returns (labeled "J-stat (Pricing)"). The numbers in brackets underneath each J-stat denote the corresponding p-value. The sample is 1987-2013. The restrictions are reported in the bottom panel of the table and refer to the second moment of the growth rate of the exchange rate of each G-10 currency relative to the risk-free rate (eq. (22) in the main text, labeled as "FX Vol."), the average intertemporal condition of each G-10 country's NFA position (eq. (23) in the main text, labeled as "NFA"); the Euler equation for the currency excess returns of each G-10 currency relative to the US (labeled "RFX (G10)"), for the currency exces returns of the 6 HML and IMX portfolios (labeled "HML (6 portf)" and "IMX (6 portf)," respectively), and for the excess return of the top over the bottom HML and IMX portfolios (labeled "HML (6-1)" and "IMX (6-1)"); and the Euler equation restrictions for each G-10 country's stock market and risk-free rate returns (labeled "Market Returns" and "Risk-Free Rates," respectively).

Discussion. Table 5 shows that, when we use only the second moments of the cross section of currencies, the preference parameters are in line with our baseline calibration, although imprecisely estimated (column 1). A test for the null hypothesis that investors have time-additive CRRA preferences (i.e., $\gamma = 1/\psi$) cannot be rejected at conventional confidence levels. The results improve when we add the cross section of NFA positions to our estimation exercise (columns 2 and 3): the preference parameters are now sharply identified, and we reject the hypothesis that $\gamma = 1/\psi$. Additionally, when we combine both sets of moment restrictions, the test for overidentifying restrictions cannot be rejected (column 3).

The introduction of currency returns produces estimates of ψ and γ of approximately 1 and 4, respectively (columns 4, 5, and 6). The results in column 7 document that these statements extend also to the case in which we focus on the smaller cross section of G-10 currencies, which is the main focus of this paper. Furthermore, we cannot reject the null hypothesis that the subset of moment conditions associated with the Euler equations is equal to zero (row labeled "J-stat (Pricing)"). This confirms the model's ability to match a large cross section of both macroeconomic and pricing restrictions.

The last two columns of table 5 demonstrate the ability of our model to also produce a good fit of the Euler equation restrictions of domestic stocks and short-term interest rates. The resulting estimated preference parameters are in line with the ones estimated for the other specifications and are robust to the inclusion or exclusion of moment restrictions that pertain to the cross-section of currency returns.

More broadly, these results support the ability of our model's specification to explain several empirical facts in a large international cross section of both macroeconomic and financial variables.

5 Concluding Remarks

In this paper, we provide novel empirical evidence regarding cross-country heterogeneity in exposure to global long-run growth news. In particular, we show that heterogenous exposure to global long-run output growth risk simultaneously accounts for many currency risk-factor structures that have been proposed in the literature. We then develop a frictionless general-equilibrium model featuring long-run growth news shocks and multiple countries with recursive preferences. Our model yields an array of important economic implications that can be empirically assessed. We find a good alignment between our model and the data.

Future developments should focus on extending this setting to international real business cycle models in order to study the role of international investment flows and international frictions for the cross section of currency risk premia. The investigation of the roles of trading frictions, portfolio composition, and market incompleteness offers other promising directions for future research.

References

Anderson, E. 2005. The dynamics of risk-sensitive allocations. *Journal of Economic Theory* 125(2):93–150.

Backus, D., and G. Smith. 1993. Consumption and real exchange rates in dynamic exchange economies with nontraded goods. *Journal of International Economics* 35: 297–316.

Backus, D. K., F. Gavazzoni, C. Telmer, and S. E. Zin. July 2010. Monetary policy and the uncovered interest parity puzzle. (16218). Working Paper, NBER.

Bansal, R., and I. Shaliastovich. 2013. A long-run risks explanation of predictability puzzles in bond and currency markets. *Review of Financial Studies* 26(1):1–33.

Bansal, R., and A. Yaron. 2004. Risks for the long run: A potential resolution of asset pricing puzzles. *Journal of Finance* 59:1481–1509.

Bansal, R., D. Kiku, and A. Yaron. 2010. Risks for the long run: Estimation and inference. *Working Paper.*

Barro, R. 2006. Rare disasters and asset markets in the twentieth century. *Quarterly Journal of Economics* 121:823–866.

Beeler, J., and J. Y. Campbell. 2012. The long-run risks model and aggregate asset prices: An empirical assessment. *Critical Finance Review* 1(1):141–182.

Brandt, W., J. Cochrane, and P. Santa-Clara. 2006. International risk sharing is better than you think, or exchange rates are too smooth. *Journal of Monetary Economics* 53(4):671–698.

Caballero, R. J., E. Farhi, and P.-O. Gourinchas. 2008. An equilibrium model of ”global imbalances” and low interest rates. *American Economic Review* 98(1):358–93.

Colacito, R. 2008. Six anomalies looking for a model. a consumption based explanation of international finance puzzles. *Working Paper, Department of Finance, University of North Carolina, Chapel Hill NC*.

Colacito, R., and M. Croce. 2013. International asset pricing with recursive preferences. *Journal of Finance* 68(6):2651–2686.

Colacito, R., and M. M. Croce. 2011. Risks for the long run and the real exchange rate. *Journal of Political Economy* 119(1):153–182.

Colacito, R., M. M. Croce, S. Ho, and P. Howard. 2014. Bkk the ez way. Working Paper.

Colacito, R., M. M. Croce, and Z. Liu. 2017. Recursive allocations and wealth distribution with multiple goods. Working paper.

Croce, M. M. 2014. Long-Run Productivity Risk: A New Hope For Production-Based Asset Pricing. *Journal of Monetary Economics*.

Della Corte, P., S. J. Riddiough, and L. Sarno. 2013. Currency premia and global imbalances. Working paper.

Epstein, L. G., and S. E. Zin. July 1989. Substitution, risk aversion, and the temporal behavior of consumption and asset returns: A theoretical framework. *Econometrica* 57(4):937–69.

Farhi, E., and X. Gabaix. 2008. Rare disasters and exchange rates. *Working Paper, Harvard and NYU*.

Farhi, E., and I. Werning. 2014. Dilemma not trilemma? Capital controls and exchange rates with volatile capital flows. *IMF Economic Review* 62(4):569–605.

Farhi, E., S. Fraiberger, X. Gabaix, R. Ranciere, and A. Verdelhan. 2015. Crash risk in currency markets. *Working Paper, Harvard and NYU*.

Froot, K. A., and J. C. Stein. 1991. Exchange rates and foreign direct investment: An imperfect capital markets approach. *The Quarterly Journal of Economics* 106(4): 1191–1217.

Gabaix, X. 2012. Variable rare disasters: An exactly solved framework for ten puzzles in macro-finance. *Quarterly Journal of Economics* 127(2):645–700.

Gabaix, X., and M. Maggiori. 2015. International liquidity and exchange rate dynamics. *Quarterly Journal of Economics, forthcoming*.

Gourinchas, P.-O., and H. Rey. 2007. International financial adjustment. *Journal of Political Economy* 115(4):665–703.

Gourinchas, P.-O., and H. Rey. 2013. External adjustment, global imbalances and valuation effects. *Handbook of International Economics* 4.

Gourio, F. 2012. Disaster risk and business cycles. *American Economic Review* 102: 2734–2766.

Gourio, F., M. Siemer, and A. Verdelhan. 2014. Uncertainty betas and international capital flows. *MIT Working Paper*.

Hassan, T. 2013. Country size, currency unions, and international asset returns. *The Journal of Finance* 68(6):2269–2308.

Hassan, T., and R. Mano. 2014. Forward and spot exchange rates in a multi-currency world. *IMF and University of Chicago Working Paper*.

Heyerdahl-Larsen, C. 2015. Asset Prices and Real Exchange Rates with Deep Habits. *Review of Financial Studies*.

Lane, P. R., and G. M. Milesi-Ferretti. 2007. The external wealth of nations mark ii: Revised and extended estimates of foreign assets and liabilities, 19702004. *Journal of International Economics* 73:223–250.

Le, A., and K. J. Singleton. 2010. An equilibrium term structure model with recursive preferences. *The American Economic Review* 100(2):557–561.

Lewis, K. 2011. Global asset pricing. *Annual Reviews of Financial Economics*.

Lustig, H., and A. Verdelhan. 2007. The cross section of foreign currency risk premia and consumption growth risk. *American Economic Review* 97:89–117.

Lustig, H., N. Roussanov, and A. Verdelhan. 2011. Common risk factors in currency markets. *Review of Financial Studies* 24(11):3731–3777.

Lustig, H., N. Roussanov, and A. Verdelhan. 2014. Counter-cyclical currency risk premia. *Journal of Financial Economics* 111(3):527–553.

Maggiori, M. 2011. Financial intermediation, international risk sharing, and reserve currencies. Working paper.

Obstfeld, M. Time of troubles: The yen and Japan's economy, 1985–2008. In Hamada, K., A. Kashyap, and D. Weinstein, editors, *Japan's Bubble, Deflation, and Long–term Stagnation* chapter 3. MIT Press 2010.

Obstfeld, M., and K. Rogoff. 1995. Exchange rate dynamics redux. *The Journal of Political Economy* 103(3):624–660.

Rabanal, P., J. Rubio-Ramirez, and V. Tuesta. 2011. Cointegrated tfp processes and international business cycles. *Journal of Monetary Economics* 58.

Rabitsch, K., S. Stepanchuk, and V. Tsyrennikov. 2015. International portfolios: A comparison of solution methods. *Journal of International Economics*.

Ready, R., N. Roussanov, and C. Ward. 2012. Commodity Trade and the Carry Trade: A Tale of Two Countries. Working Paper, the Wharton School.

Stathopoulos, A. 2012. Asset prices and risk sharing in open economies. *Working Paper, University of Washington*.

Verdelhan, A. 2010. A habit-based explanation of the exchange rate risk premium. *The Journal of Finance* 65(1):123–145.

Verdelhan, A. 2015. The share of systematic risk in bilateral exchange rates. *Working Paper, MIT*.

Weil, P. January 1989. The equity premium puzzle and the riskfree rate puzzle. *Journal of Monetary Economics* 24(3):401–422.

Zviadadze, I. 2013. Term-structure of consumption risk premia in the cross-section of currency returns. *Working Paper Stockholm School of Economics*.

Appendix

A Empirical Evidence: Robustness

In this section we report additional empirical evidence in support of the results presented in section 2 of the paper.

Tobin's Q. As an alternative to using p/d ratios, we have collected data on aggregate Tobin's Q, defined as

$$Q = \frac{\text{Stock market capitalization + Total credit}}{\text{book value of capital}}.$$

This variable is frequently used in the empirical literature as a forward-looking economic indicator (for a recent DSGE-based approach, see Croce (2014)), and it does not require us to use dividends. The one caveat for this variable is that the common sample for the cross section of our 10 countries starts in 1995. Given the short time series, we pool the estimates of the exposure of each country's Tobin's Q to the global Tobin's Q across three portfolios denoted as Low-, Medium-, and High-β, respectively. The Low-β portfolio includes New Zealand and Australia, that is, the countries that in table 1 of our manuscript have significantly negative coefficients. The High-β portfolio includes Switzerland, Sweden and the US, as they have a significantly positive exposure to the global component. All other countries feature an insignificant coefficient are grouped in the Medium-β portfolio. In panel B of table A1, we show that our results on heterogeneous exposure to global long-run risks are also confirmed when using Tobin's Q.

We also estimate the parameters that govern the transition dynamics of the growth rate of the endowments using the lagged Tobin's Q as the predictor (see panel A of

TABLE A1: Dynamics of Endowments and Predictive Components Using Tobin's Q

Panel A: Estimation of Predictive Components

	ϕ	ρ_x	σ	φ_e
Parameters	0.002***	0.930***	0.021***	0.005***
(S.E.)	(0.000)	(0.012)	(0.001)	(0.000)

Panel B: Exposure to Global Long-Run Risk

	Low β	**Medium** β	**High** β
β^i	-0.35^{***}	-0.04^{**}	0.25^{***}
(S.E.)	(0.04)	(0.02)	(0.04)

Notes - Panel A reports the estimation of the parameters of the endowment process reported in equation (1). Panel B reports the estimation of the exposure of each country's predictive component of GDP to the global predictive component (see equation (3)). The sample is 1995–2013. In both panels, the numbers in parentheses are heteroskedasticity-adjusted standard errors. One, two, and three stars denote statistical significance at the 10%, 5%, and 1% levels, respectively.

table A1). Our results confirm the basic set of findings that we have in table 1: (1) Q is a predictor of the endowments' growth rate, (2) the predictive component is very persistent, and (3) the size of the long-run risk is smaller than that of the short-run risk.

Longer sample. In table A2, we show the results for the case in which the exposures to the global long-run risk are computed for the longest sample that we have (starting in 1975). The drawback is that we do not have data on New Zealand for the entire sample dating back to 1975, and thus we must exclude this country from our estimation exercise. Other than this shortcoming, the results are very much consistent with those reported in table 1.

In addition to the limited availability of data for a large cross section of countries, we note that the choice to start with the common sample from 1987 onward is consistent with the presence of a common exchange rate regime. According to our model, we should estimate exposures in a sample in which all exchange rates are floating and financial markets are integrated. Given our cross section of countries, we note that

TABLE A2: Exposure to Global Long-Run Risk, Starting in 1975

	NZ	AUS	UK	GER	CAN	NOR	JPN	SUI	US	SWE
β^i	-	-0.55^{***}	-0.13	-0.03	-0.08	0.09	0.19^{*}	0.01	0.28^{**}	0.21^{**}
(S.E.)	(-)	(0.06)	(0.06)	(0.09)	(0.09)	(0.13)	(0.11)	(0.08)	(0.09)	(0.10)

Notes - This table reports estimates of the exposure of each country's shock to the predictive component of GDP to the global shock to the predictive component. The numbers in parentheses are heteroskedasticity-adjusted standard errors. One, two, and three stars denote statistical significance at the 10%, 5%, and 1% levels, respectively.

the collapse of the Bretton-Woods fixed exchange rate regime happened in the early 1970s, the Australian dollar started being free floating only in 1983, and New Zealand adopted a flexible FX regime only in 1985. For these reasons, our baseline analysis in the main text continues to use the 1987–2013 sample.

Time-varying risk premia. We use weekly data on stock market returns for the G-10 countries to compute country-level annual integrated variance (henceforth RV), which we interpret as a measure of time-varying risk premia. For each country, we project its p/d ratio onto its contemporaneous RVs and then we repeat our baseline analysis using the residuals of this regression in an effort to remove the time-varying risk premium component of the p/d ratio. More precisely, this is the system of equations that we jointly estimate:

$$\begin{aligned} pd_t^i &= \phi^{RV} \cdot RV_t^i + \varepsilon_{pd,t}^i \\ \Delta y_t^i &= \underbrace{\phi^y \cdot \varepsilon_{pd,t}^i}_{z_t^i} + \varepsilon_t^i \\ z_t^i &= \rho z_{t-1}^i + \varepsilon_{z,t}^i \end{aligned} \tag{A1}$$

We report our estimation results in table A3 and note that they confirm our initial findings. First of all, the portion of the p/d ratio orthogonal to RV ($\varepsilon_{pd,t}^i$) forecasts future growth with a significant positive sign ($\phi^y > 0$), just as in our baseline regressions in table 1. Furthermore, the persistence of expected growth is similar to that in

TABLE A3: The Role of the Cash-Flow Components of P/D Ratios

Panel A: Estimation of Predictive Components

	$\phi^{\mathbf{RV}}$	$\phi^{\mathbf{y}}$	ρ
Parameters	-1.03^{***}	0.01^{***}	0.79^{***}
(S.E.)	(0.08)	(0.00)	(0.02)

Panel B: Exposure to Global Long-Run Risk

	NZ	**AUS**	**UK**	**GER**	**CAN**	**NOR**	**JPN**	**SUI**	**US**	**SWE**
β^i	-0.52^{***}	-0.48^{***}	-0.10	0.02	-0.04	0.05	0.09	0.27^{*}	0.32	0.32^{*}
(S.E.)	(0.20)	(0.06)	(0.14)	(0.11)	(0.14)	(0.19)	(0.15)	(0.16)	(0.23)	(0.18)

Notes - Panel A reports the estimation of the parameters in the first two equations of system (A1). Panel B reports the estimation of the exposure of each country's predictive component of the GDP growth rate, z_t^i, to the cross-sectional average of the predictive components. The numbers in parentheses are heteroskedasticity-adjusted standard errors. One, two, and three stars denote statistical significance at the 10%, 5%, and 1% levels, respectively.

our baseline specification. Additionally, we find it reassuring that there is a negative relationship between p/d ratios and RVs ($\phi^{RV} < 0$), consistent with the idea that stock market valuation is low when risk premia are high.

Second, when we look at the degree of heterogeneity of the exposure to the global LRR measure obtained after removing the realized variance components, we get virtually the same results that we currently have in table 1. We interpret this as strong evidence that our results concerning the heterogeneous exposure to global long-run risk are driven by the cash-flow components of the p/d ratios.

Heterogeneous exposure using shocks. We repeat our analysis in table 1 by using the shocks, as opposed to the levels, of the long-run risks. Specifically, adopting the same notation of equation (1) in main text, we estimate the following system of equations:

$$\begin{aligned} \underbrace{\varepsilon_t^i}_{\text{Short-Run Shock}} &= \left(1+\beta_{\Delta y}^i\right) \cdot \left(\frac{1}{n}\sum_{i=1}^{n} \varepsilon_t^i\right) + \xi_t^i \\ \underbrace{\varepsilon_{z,t}^i}_{\text{Long-Run Shock}} &= \left(1+\beta^i\right) \cdot \left(\frac{1}{n}\sum_{i=1}^{n} \varepsilon_{z,t}^i\right) + \zeta_t^i, \end{aligned} \quad \text{(A2)}$$

TABLE A4: Exposures Using Shocks

Panel A: Exposure to Global Endowment Risk

	NZ	AUS	UK	GER	CAN	NOR	JPN	SUI	US	SWE
$\beta^i_{\Delta y}$	-0.65	-0.12	-0.01	-0.02	0.16	0.38	0.12	-0.22	-0.32^{**}	-0.04
(S.E.)	(0.73)	(0.64)	(0.32)	(0.50)	(0.20)	(0.38)	(0.45)	(0.53)	(0.14)	(0.25)

Panel B: Exposure to Global Long-Run Risk

	NZ	AUS	UK	GER	CAN	NOR	JPN	SUI	US	SWE
β^i	-0.39^{***}	-0.38^{***}	-0.04	0.14	-0.41	0.18	0.11	0.172	0.25^{*}	0.28^{**}
(S.E.)	(0.11)	(0.05)	(0.06)	(0.21)	(0.33)	(0.48)	(0.12)	(0.14)	(0.14)	(0.09)

Notes - Panel A reports the estimation of the exposure of each country's short-run shock to the GDP growth rate to the global shock to the GDP growth rate. Panel B reports the estimation of the exposure of each country's shock to the predictive component of GDP to the global shock to the predictive component. The numbers in parentheses are heteroskedasticity-adjusted standard errors. One, two, and three stars denote statistical significance at the 10%, 5%, and 1% levels, respectively.

$\forall i \in \{$G-10 countries$\}$, and report the results in table A4. Consistent with the results in table 1, there is no cross-country heterogeneity in exposure to global short-run shocks (see table A4, Panel A), whereas the opposite is true for long-run growth risk (see table A4, Panel B).

Robustness of the estimates in figure 8. The inference reported in figure 8 does not take into account the first-step estimation of the β coefficients. In order to address this issue, in this section we add the orthogonality conditions associated with the estimation of the βs to our GMM estimator and perform a joint estimation.

Before presenting the results, we must note that we cannot include the orthogonality conditions for the estimation of all 10 βs simultaneously, because this gives rise to invertibility issues in the covariance matrix of the moment conditions. This can be appreciated by looking at the set of moment conditions associated with the estimation of the βs

$$\frac{1}{T}\sum_{t=1}^{T}\left(z_t^i - (1+\beta^i)\sum_{j=1}^{10}\frac{z_t^j}{10}\right)\left(\sum_{j=1}^{10}\frac{z_t^j}{10}\right) = 0, \forall j = \{1, ..., 10\}.$$

By taking the sum of all moment conditions and dividing by 10, we get

$$\frac{1}{10}\sum_{i=1}^{10}\left[\frac{1}{T}\sum_{t=1}^{T}\left(z_t^i-(1+\beta^i)\sum_{j=1}^{10}\frac{z_t^j}{10}\right)\right]\left(\sum_{j=1}^{10}\frac{z_t^j}{10}\right)=0,$$

from which it follows that

$$\frac{1}{T}\sum_{t=1}^{T}\left[\sum_{i=1}^{10}\left(\frac{z_t^i}{10}-\left(\sum_{i=1}^{10}\frac{1+\beta^i}{10}\right)\sum_{j=1}^{10}\frac{z_t^j}{10}\right)\right]\left(\sum_{j=1}^{10}\frac{z_t^j}{10}\right)=0.$$

Using the estimated betas reported in table 1, we obtain that

$$\left(\sum_{i=1}^{10}\frac{1+\beta^i}{10}\right) \quad = \quad 0.999\approx 1,$$

which implies

$$\frac{1}{T}\sum_{t=1}^{T}\left[\sum_{i=1}^{10}\left(\frac{z_t^i}{10}-\sum_{j=1}^{10}\frac{z_t^j}{10}\right)\right]\left(\sum_{j=1}^{10}\frac{z_t^j}{10}\right)=0,$$

thus explaining the invertibility issues associated with the optimal GMM weighting matrix. We overcome this difficulty by excluding the estimation of one beta from the system. Given that the estimation of the λs captures the response of exchange rates relative to the US dollar, we also exclude the US from the first part of the estimation exercise (the estimation of the betas) and take the beta of the US as given.

We report the results for this joint estimation exercise in table A5 (row labeled "Joint"). For comparison, we also report the estimates for the two-step estimation exercise (row labeled "Two Steps"). Both the point estimates and the degrees of statistical significance are virtually identical across the two estimation procedures. Hence the message conveyed by figure 8 of main text remains unchanged.

TABLE A5: Joint Estimate of Loadings and Slope

	β_{NZ}	β_{AUS}	β_{UK}	β_{GER}	β_{CAN}	β_{NOR}	β_{JPN}	β_{SUI}	β_{US}	β_{SWE}	$\zeta_{0,FX}$	$\zeta_{1,FX}$
Two Steps	-0.52^{***}	-0.44^{***}	-0.09	-0.02	0.00	0.08	0.12	0.26^{**}	0.27^{*}	0.33^{**}	-26.11^{***}	-33.02^{***}
	(0.15)	(0.06)	(0.10)	(0.09)	(0.13)	(0.17)	(0.17)	(0.13)	(0.17)	(0.15)	(7.55)	(11.97)
Joint	-0.54^{***}	-0.47^{***}	-0.07	0.00	0.01	0.09	0.16	0.22^{**}	0.27	0.27^{***}	-25.29^{***}	-33.15^{**}
	(0.10)	(0.04)	(0.05)	(0.09)	(0.06)	(0.10)	(0.09)	(0.07)	##	(0.08)	(9.41)	(15.17)

Notes - The first ten columns report the estimation of the exposure of each country's shock to the predictive component of GDP to the global shock to the predictive component. The last two columns report the estimates of the intercept and slope coefficients in equation (20). The row labeled "Two Steps" reports the estimated coefficients of the sequential estimation (the estimates of $\zeta_{,FX}$ and $\zeta_{1,FX}$ are conditional on the estimates of the βs). The row labeled "Joint" reports the joint estimation of all the coefficients conditional on β_{US}. The numbers in parentheses are heteroskedasticity-adjusted standard errors. One, two, and three stars denote statistical significance at the 10%, 5%, and 1% levels, respectively.

B Derivations for the General Equilibrium Model

B.1 Allocations as a function of Pareto weights

Let $W_t^i = W(C_t^i, U_{t+1}^i)$ be the right-hand side of equation (5). If we denote the partial derivatives of the aggregator W^i as follows,

$$W_{1,t}^i := \frac{\partial W_t^i}{\partial C_t^i}, \quad W_{2,t}^i := \frac{\partial W_t^i}{\partial U_{t+1}^i},$$

the stochastic discount factor is equal to

$$M_{t+1}^i = \frac{W_{2,t}^i W_{1,t+1}^i}{W_{1,t}^i} \quad \forall i = \{h, f\}. \tag{B3}$$

The optimality condition for the allocation of good $X_{j,t}$ for $t = 1, 2, ...$ in each possible state is

$$\mu_0^j \cdot \left(\prod_{k=0}^{t-1} W_{2,k}^j \right) \cdot W_{1,t}^j C_t^j \frac{\alpha}{x_{j,t}^j} = \frac{(1-\alpha)}{(N-1)} \cdot \frac{1}{x_{j,t}^i} C_t^i W_{1,t}^i \cdot \left(\prod_{k=0}^{t-1} W_{2,k}^i \right) \cdot \mu_0^j \tag{B4}$$

for all countries $i \neq j$. Define the time t Pareto weights as

$$
\begin{aligned}
\mu_t^i &= \mu_0^i \cdot \left(\prod_{j=0}^{t-1} W_{2,j}^i \right) \cdot W_{1,t}^i C_t^i \\
&= \mu_{t-1}^i \cdot W_{2,t-1}^i \cdot \frac{W_{1,t}^i}{W_{1,t-1}^i} \cdot \frac{C_t^i}{C_{t-1}^i} = \mu_{t-1}^i \cdot M_t^i \cdot \exp\left\{ \Delta c_t^i \right\}, \quad \forall i \in \{h, f\}
\end{aligned}
$$

It follows that equation (B4) can be rewritten as

$$
\mu_t^j \cdot \frac{\alpha}{x_{j,t}^j} = \frac{(1-\alpha)}{(N-1)} \cdot \frac{1}{x_{j,t}^i} \cdot \mu_t^i \tag{B5}
$$

Let $S_{j,t} := \mu_{j,t}/\mu_{1,t}$. Then the optimality condition in equation (B5) combined with the feasibility constraint can be represented by the following system of recursive equations:

$$
\begin{aligned}
x_{i,t}^i &= \left(1 + \frac{1-\alpha}{\alpha(N-1)} \sum_{j \neq i} \frac{S_{j,t}}{S_{i,t}} \right)^{-1} X_{i,t}, \quad \forall i \in \{1, 2, ..., N\} \qquad \text{(B6)} \\
x_{i,t}^j &= \frac{1-\alpha}{\alpha} \frac{1}{N-1} \frac{S_{j,t}}{S_{i,t}} x_{i,t}^i, \quad \forall i \neq j \in \{1, 2, ..., N\} \\
S_{j,t} &= S_{j,t-1} \cdot \frac{M_{j,t}}{M_{1,t}} \cdot \left(\frac{C_{j,t}/C_{j,t-1}}{C_{1,t}/C_{1,t-1}} \right), \quad \forall t \geq 1.
\end{aligned}
$$

We use perturbation methods to solve our system of equations (4)–(14) in main text. We compute our policy functions using the `dynare++4.2.1` package. All variables are expressed in log-units.

B.2 Terms of trade, imports, and exports

We normalize the price of good 1 to 1. The terms of trade can be obtained from the intratemporal condition:

$$-p_{j,t}\frac{C_t^1}{x_{1,t}^1}\alpha+\frac{1-\alpha}{N-1}\frac{C_t^1}{x_{j,t}^1}=0,\quad \forall j=\{2,3,...,N\},$$

which implies

$$p_{j,t}=\frac{1-\alpha}{\alpha(N-1)}\frac{x_{1,t}^1}{x_{j,t}^1},\quad \forall j=\{2,3,...,N\}.$$

Consider country 1 and country j. The exports of country 1 to country j are

$$Exp_{1,t}^j=x_{1,t}^j,$$

where $x_{1,t}^j$ is defined as in equation (11). The imports of country 1 from country j are

$$\begin{aligned}Imp_{j,t}^1 &= p_{j,t}\cdot x_{j,t}^1=\frac{1-\alpha}{\alpha}\frac{1}{N-1}\frac{x_{1,t}^1}{x_{j,t}^1}x_{j,t}^1\\ &= \frac{1-\alpha}{\alpha}\frac{1}{N-1}\left(1+\frac{1-\alpha}{\alpha(N-1)}\sum_{j\neq 1}S_{j,t}\right)^{-1}X_{1,t}.\end{aligned}$$

It follows that the volume of trade between countries 1 and j normalized by the endowment of country 1 is

$$\begin{aligned}\frac{Vol_{1,t}^j}{X_{1,t}} &= \frac{\frac{1-\alpha}{\alpha}\frac{1}{N-1}(1+S_t^j)}{1+\frac{1-\alpha}{\alpha}\frac{1}{N-1}\sum_{j\neq 1}S_t^j}\\ &= \frac{(1-\alpha)\cdot(1+S_t^j)}{\alpha(N-1)+(1-\alpha)\sum_{j\neq 1}S_t^j}.\end{aligned} \tag{B7}$$

Similarly, the net exports–output ratio between countries 1 and j is

$$\begin{aligned} \frac{NX_{1,t}^j}{X_{1,t}} &= \frac{\frac{1-\alpha}{\alpha}\frac{1}{N-1}(S_t^j - 1)}{1 + \frac{1-\alpha}{\alpha}\frac{1}{N-1}\sum_{j\neq 1} S_t^j} \\ &= \frac{(1-\alpha)\cdot(S_t^j - 1)}{\alpha(N-1) + (1-\alpha)\sum_{j\neq 1} S_t^j}. \end{aligned} \tag{B8}$$

C Numerical Accuracy

In table C6, we report statistics regarding the maximum cumulative approximation error for our relative pseudo-Pareto weights, that is, the key drivers of both the consumption shares and the NFA positions. The maximum is taken across countries, and we account for the fact that the relative-Pareto weights are persistent and errors can cumulate over longer simulations.

More specifically, we construct recursively the following processes:

$$\tilde{S}_{j,t} = \tilde{S}_{j,t-1} \cdot \frac{\hat{M}_{j,t}}{\hat{M}_{1,t}} \cdot \left(\frac{\hat{C}_{j,t}/\hat{C}_{j,t-1}}{\hat{C}_{1,t}/\hat{C}_{1,t-1}} \right), \quad \forall t \geq 1 \tag{C9}$$

where we adopt the "$\hat{\cdot}$" notation to indicate approximated variables. We initialize the recursion with $\tilde{S}_{j,0} = 0$, consistent with our steady state. We then define the error for country j as follows:

$$err_t^j = \sum_{\tau=1}^{t} \left| \frac{\tilde{S}_{j,\tau}}{\hat{S}_{j,\tau}} - 1 \right| \cdot 100.$$

$\widehat{S}_j$ is the approximated relative pseudo-Pareto weight that we obtain directly from our perturbation method. $\tilde{S}_j$ is constructed by dynamically updating our equilibrium recursion (i.e., an exact dynamic equation) for $S_{j,t}$ using our approximated forcing elements, $\hat{M}_{j,t}$, $\hat{M}_{1,t}$, $\hat{C}_{j,t}/\hat{C}_{j,t-1}$, and $\hat{C}_{1,t}/\hat{C}_{1,t-1}$. Examination of the discrepancy between $\tilde{S}_j$ and $\hat{S}_j$ constitutes a valid metric for assessing the quality of the approximation,

TABLE C6: Approximation Errors

CRRA, Homog.		EZ, Homog.	
Median	95%	Median	95%
7.98E-09	2.32E-07	5.54E-07	3.30E-05
CRRA, Heterog.		EZ, Heterog.	
Median	95%	Median	95%
4.00E-10	8.68E-09	1.74E-09	6.63E-10

In this table, we report the median and the 95^{th} percentile of the distribution of the maximum cumulative approximation errors for our relative pseudo-Pareto weights, $S_t^i = \mu_t^i/\mu_t^1$ for $i = 2, ..., 5$. EZ (CRRA) refers to the setting with time-additive (recursive) preferences. Homog. (Heterog.) refers to the case in which all countries have the same (heterogeneous) exposure to global growth news shocks. All numbers are multiplied by 100.

as this discrepancy accounts for the persistent impact that the errors on our forcing terms can have on the relative pseudo-Pareto weights.

Across all cases, the approximation errors are small. In the case of homogeneous exposure and EZ preferences, the pseudo-Pareto weights are particularly persistent and, as a result, they feature slightly higher cumulative errors. Under our benchmark specification, however, the approximation errors are comparable to those obtained under regular time-additive preferences and thus are negligible for all practical purposes.

Rabitsch et al. (2015) point out that a global approximation is required when countries are subject to asymmetric constraints, such as a borrowing limit, and when their wealth distribution is nonstationary. Since we have a frictionless model with complete markets and a well defined ergodic distribution of wealth, a perturbation approach provides a good approximation of the equilibrium.

D Homogeneous Exposure: Inspecting the Mechanism

In this section, we explore further the basic properties of a recursive risk-sharing scheme with multiple countries with homogeneous exposure to endowment shocks. We show that in terms of allocation dynamics, a multicountry economy preserves the basic success of a two-country setting. In terms of risk-sharing measurements, in contrast, we highlight several relevant differences.

Response to shocks. In figure D1 we report the response of some variables of interest to short- and long-run local shocks. For the sake of the exposition we consider the case in which country 1 receives a positive local shock. That is, country 1 receives a positive shock (panels labeled Δx), whereas the remaining countries experience no change in their endowment. Since all countries share the same calibration, both the endowments and the pseudo-Pareto weight dynamics of countries 2–5 are identical. Because of our symmetric calibration, in this section we denote country 1 as the home country and refer to the remaining countries as the foreign countries.

The risk-sharing arrangement in this economy prescribes that in response to both a positive short- and long-run shock to the home country, the foreign countries experience an equal increase of their share of world consumption (panels labeled s). Specifically, upon the arrival of a positive endowment shock, the marginal utility of country 1 declines substantially because of home bias. In order to re-establish the equality of marginal utilities across countries, goods are reallocated toward the foreign countries. In both cases the real exchange rate of country 1 depreciates because of the larger current (expected) supply of the domestic good associated with a positive short-run (long-run) shock (panels labeled Δe).

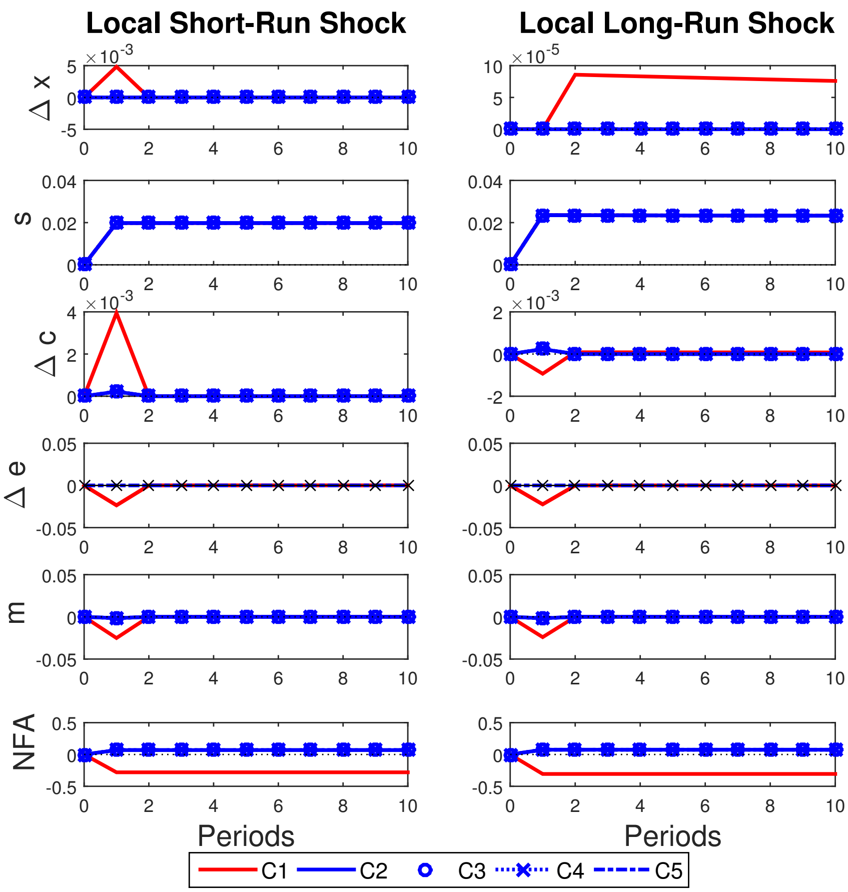


FIG. D1 - **Impulse response functions under homogeneous exposure.** The left (right) panels report the response of endowment growth ($\Delta \log X^i$), relative Pareto weights ($\log S_j/S_1$), consumption growth (Δc_i), exchange rate growth (Δe_3^i), stochastic discount factors (m_i), and NFA (A_i/X_i) to a one-standard-deviation short-run (long-run) shock to the endowment of country 1 only. All panels refer to the case in which the economy consists of 5 countries ($i = 1, ..., 5$). The exchange rate is measured with respect to country 3, implying that $\Delta e < 0$ for country 1 denotes a depreciation of its real exchange rate with respect to country 3.

The panels of figure D1 labeled Δc document a different response of consumption growth rates with respect to short- and long-run shocks. When a positive short-run shock materializes, consumption increases in the entire cross section of countries, as only part of country 1's additional resources are being redistributed abroad. In response to positive long-run news, in contrast, consumption drops in the home country as the risk-sharing redistribution effect dominates. The foreign countries experience a positive growth rate of consumption, which is required to equalize their marginal utilities to that of the home country.

This differential response of consumption to short- and long-run endowment shocks results in the spread between country 1's and the other countries' consumption being positive in one case and negative in the other. Since country 1's exchange rate depreciates in response to both types of shocks, the model produces a less-than-perfect correlation between currency movements and consumption differentials, as in the data (Backus and Smith 1993) and as already shown by Colacito and Croce (2013).

The last two panels in figure D1 report the response of the NFA to the two types of shock. The country that receives either a positive short- or long-run shock acts as an insurance provider to the other countries. As a consequence, it experiences a drop in its NFA. This negative link between NFA and positive exogenous shocks induced by our risk-sharing motive plays a crucial role in the next sections, once we account for heterogeneity.

Simulated moments. In table D7 we report the results of our analysis. Specifically, we analyze how several moments of interest change as the number of countries in the economy (NC) increases from 2 to 5. We obtain several relevant results as the number of countries increases.

First, country-level consumption volatility decreases. This is a direct reflection of the better risk-sharing opportunities that are available in a multicountry economy. A similar argument can be used to explain the decline in the volatility of both the pseudo-Pareto weights and the net-exports-to-output ratio. Second, the bilateral correlations of both consumption growth rates and stochastic discount factors decline. Under complete markets, less-correlated discount factors immediately imply more volatile bilateral exchange rates. At a country level, hence, better risk-sharing op-

TABLE D7: Simulated Moments with Homogeneous Exposure

	$NC = 2$	$NC = 3$	$NC = 4$	$NC = 5$
Std(Δy)	1.94	1.94	1.94	1.93
ACF$_1$(Δy)	0.07	0.07	0.07	0.06
corr(Δy_t^h,Δy_t^f)	0.25	0.25	0.25	0.24
Single-Country Moments				
Std(Δc)	1.78	1.75	1.74	1.72
ACF$_1$ (Δc)	0.07	0.07	0.07	0.06
Bilateral Moments				
corr(Δc_t^h,Δc_t^f)	0.49	0.43	0.4	0.38
Std (Δe)	14	15.14	15.47	15.72
corr(m, m^f)	0.94	0.93	0.93	0.93
corr($\Delta c^h - \Delta c^f$, Δe)	0.31	0.42	0.45	0.47
Std[NX/X]	2.04	1.80	1.80	1.80
Std[$\mu/\sum\mu$]	21.25	17.42	14.65	12.34
Trade-Weighted Moments				
Std (Δe^w)	14	13.80	13.52	13.37
corr(m, m^w)	0.94	0.95	0.95	0.95
Financial Variables				
E[r_f]	2.29	2.36	2.32	2.34
Std[r_f]	0.31	0.32	0.31	0.30
corr(r_f^h, r_f^f)	0.93	0.93	0.92	0.92
Std[M] /E[M]	40.41	40.46	41.35	42.01
Std[NFA/X]/Std(Δy)	13.08	12.15	12.09	12.47
β_{UIP}	-4.51	-5.75	-5.27	-5.35
E[HML]	n.a.	-0.25	-0.19	-0.02

Notes - The table reports annualized simulated moments from the model with homogeneous exposure as the number of countries in the economy (*NC*) ranges from 2 to 5. All parameters are set to their benchmark values reported in table 2. The first panel reports the moments for the dynamics of exogenous endowment growth rates. The panel labeled "Single-Country Moments" reports the moments of the consumption growth rate within each country. The panel labeled "Bilateral Moments" reports the cross-country moments for each country pair. The panel labeled "Trade-Weighted Moments" reports the moments for the case in which each foreign country is weighted in proportion to its volume of trade with country 1. The panel labeled "Financial Variables" reports the moments for the risk-free rates (r_f), stochastic discount factors (M), NFA-to-output ratio, excess returns (rx), slope coefficient of the UIP regressions (β_{UIP}), and average currency risk premium (E[HML]).

portunities are not necessarily accompanied by higher correlations of consumption profiles and smoother exchange rates (Brandt et al. 2006).

Third, international correlations and exchange rate volatility are very sensitive to the trade dynamics arising from our recursive risk sharing. We construct trade-weighted variables by weighting each country's stochastic discount factor and exchange rate in proportion to its volume of trade with country 1 (see equation (15)). As documented in the panel labeled "Trade-Weighted Moments," when the number of countries rises the trade-weighted correlation of stochastic discount factors increases, in contrast to what is obtained for bilateral correlations. The intuition for this result is that correlations tend to be larger between countries that are major trading partners. By no arbitrage, a higher correlation between the home country and the trade-weighted rest of the world results in a smoother exchange rate.

CRRA case. In table D8 we report some moments of interest for the special case of CRRA preferences. This case is obtained by imposing that the IES (ψ) equals the inverse of the RRA coefficient (γ). Specifically, we set $\psi = 1/6.5$, and keep all other values at their benchmark values.

This experiment highlights three relevant results. First, in a world with multiple goods and time-additive preferences, the volatility of the bilateral real exchange rate is still a misleading indicator of risk sharing, as in the case of recursive preferences.

Second, since growth news shocks are not priced, the risk-sharing motives are solely driven by short-run risk. As a result, the long-run risk-sharing opportunities that arise with an increasing number of countries play no role in the determination of S^j, and the consumption growth rates are predominantly driven by the properties of the exogenous short-run shocks. Since short-run shocks are poorly correlated internationally, as NC increases, the consumption bundle of each country features a larger number of goods with low correlations. By construction, a trade-weighted measure of the cross-country correlation of the discount factors reproduces the same pattern of the bilateral correlation, in contrast to what is seen with recursive preferences.

TABLE D8: Simulated Moments (CRRA Sase)

	$NC = 2$	$NC = 3$	$NC = 4$	$NC = 5$
Risk-Sharing Measures				
Std Δe	10.21	11.15	11.5	11.7
corr(m, m^f)	0.59	0.49	0.45	0.42
corr(m, m^w)	0.59	0.56	0.56	0.55
Anomalies				
corr($\Delta c^h - \Delta c^f$, Δe)	1	1	1	1
β_{UIP}	1.00	0.94	0.99	0.99
E[HML]	n.a.	0.00	0.00	0.00

Notes - This table reports simulated annualized moments as the number of countries in the economy (*NC*) ranges from 2 to 5. All parameters are set to their benchmark values reported in table 2, except for $\psi = 1/6.5$ (the CRRA case). The average correlation of the stochastic discount factors across country pairs is denoted by corr(m, m^f). The correlation of the country 1 stochastic discount factor with the volume of the trade-weighted stochastic discount factors of the remaining countries is denoted by corr(m, m^w).

To summarize, in an economy in which growth news shocks are not priced, the risk-sharing measures proposed by Brandt et al. (2006) may be misleading, regardless of whether they are bilateral or trade weighted. In an economy in which news shocks are priced, in contrast, cross-country correlations are directly related to trade volume, and hence trade-weighted measures of risk sharing may be more accurate.

Last but not least, we note that, absent trading motives associated with long-run news, well-known anomalies continue to be present even in our multicountry setting.

E Sensitivity Analysis

In this section, we report results for a sensitivity analysis of our result with respect to both the home bias parameter (α) and the speed of cointegration of the endowments (τ). We report our main results in table E9 and note that in our setting the persistence of the cointegration of the endowments across country pairs is $(1 - 2 \cdot \tau)$. For the

TABLE E9: Sensitivity Analysis

Panel A: The Role of Home Bias (α)					
Moments	Data	S.E.	Benchmark	$\alpha = 0.96$	$\alpha = 0.93$
Median					
Std(Δc)	1.91	0.25	1.96	2.08	2.24
corr($\Delta c_t^h, \Delta c_t^f$)	0.24	0.05	0.38	0.27	0.14
Std (Δe)	9.10	0.91	17.01	14.12	10.56
Std(NFA/X)/Std(Δx)	18.58	2.95	25.76	41.37	49.22
Std(NX/X)	5.12	0.74	1.48	2.13	2.61
β_{UIP}	-0.94	0.48	-4.62	-3.60	-2.50
E(HML)	3.20	1.10	3.01	2.01	1.02
Cross-sectional Dispersion					
Std(Δc)	0.45	0.12	0.21	0.36	0.50
Std (Δe) (CoV)	0.21	0.04	0.41	0.39	0.38
E(r_f)	1.27	0.26	0.54	0.45	0.42
Std(r_f) (CoV)	0.42	0.08	0.46	0.47	0.47
Std(NFA/X)/Std(Δx) (CoV)	0.55	0.09	0.68	0.55	0.56
Std(NX/X) (CoV)	0.52	0.09	0.61	0.64	0.62
Panel B: The Role of Cointegration (τ)					
Moments	Data	S.E.	Benchmark	$HL \cdot 50\%$	$HL \cdot 150\%$
Median					
Std(Δc)	1.91	0.25	1.96	1.90	2.00
Std (Δe)	9.10	0.91	17.01	16.26	18.75
Std(NFA/X)/Std(Δx)	18.58	2.95	25.76	23.27	33.55
Std(NX/X)	5.12	0.74	1.48	1.18	1.74
β_{UIP}	-0.94	0.48	-4.62	-4.04	-5.09
E(HML)	3.20	1.10	3.01	1.52	4.42
Cross-sectional Dispersion					
Std(Δc)	0.45	0.12	0.21	0.15	0.26
Std (Δe) (CoV)	0.21	0.04	0.41	0.34	0.42
Std(NX/X) (CoV)	0.52	0.09	0.61	0.60	0.67

Notes - This table reports both empirical moments computed using the data set described in section 2 and simulated moments from the model with heterogenuos exposure. All the parameters are calibrated as in table 2, unless otherwise specified. CoV denotes the cross-sectional coefficient of variation. In panel B, HL denotes the half-life of the cointegration of the endowment ($\ln(0.5)/\ln(1 - 2 \cdot \tau)$). Under the benchmark, $HL = 70$ years.

sake of brevity, we show only moments that change significantly from our benchmark calibration.

The role of home bias (α). A lower home bias alters the median moments in a very intuitive way. As the degree of home bias declines, consumption becomes more volatile, as exogenous foreign shocks matter more for the domestic consumption bundle. Simultaneously, goods are more substitutable, and hence the exchange rate volatility declines, whereas net exports and NFA positions become more volatile. Given the lower volatility of the exchange rate, our β_{UIP} mechanically declines as well, becoming closer to the data. The results also highlight the presence of a tension between moments, as the implied carry-trade risk premium declines as well.

In the cross section, lowering α helps with the cross-country spread of consumption volatility. All other moments are either unchanged or further away from the data. An exception worthy of notice is the cross-sectional dispersion of the volatility of the NFA. As consumption home bias declines, two channels work in opposite directions. On the one hand, the traded quantities become more volatile. On the other hand, the valuation channel is less relevant due to a less volatile exchange rate. As a result, the coefficient of variation for the NFA-to-output ratio declines and comes closer to the data.

The role of cointegration speed (τ). The inference on the extent of cointegration of output across countries is not very precise (see, among others, Rabanal et al. (2011) and Colacito et al. (2014)). As a result, we alter τ in order to consider a large range ($\pm 50\%$) of variation in the half-life of the cointegration residual from our benchmark.

Increasing the cointegration speed of our endowments (lower half-life) makes country-specific shocks less relevant over the long run. As a result, median volatilities tend to decline across all aggregates. In the cross section, coefficients of variation decline as well. Not surprisingly, the carry-trade risk premium declines as well.

F Numerical Approximation of the SDFs

Our GMM procedure estimates preference parameters from restrictions in which returns are taken as given, whereas SDFs are endogenous objects. In our economy with multiple countries and trade, SDF dynamics are nontrivial functions of preference parameters, fundamental shocks, and state variables. In what follows, we describe a two-step procedure to express equilibrium SDFs as a function of preference parameters and observable variables only. This procedure enables us to efficiently estimate our preference parameters.

1. **Simulation**. The target variables for the estimation are the IES, ψ, and the risk aversion, γ, so as a first step we generate simulated output from our model solved over a grid of values for these two parameters. Given spreads of β_t^i consistent with our empirical observations, we simulate our model at a monthly frequency and time aggregate all variables to an annual frequency to be consistent with our empirical data set.

2. **Projection**. After obtaining annual simulations of the equilibrium variables, we then perform a pooled projection across all simulations k, countries i, and years t of the log-SDF onto combinations of observable variables and preference parameters. Our set of observable variables comprises the risk-free rate ($r^i_{f,k,t}$); the exposure to the global news shock (β_t^i); the country-specific short-run shock ($\epsilon^i_{k,t+1}$); the country-specific long-run shock ($\epsilon^i_{z,k,t+1}$); and realizations of the global shocks ($\epsilon^{global}_{k,t+1}$). Including β_t^i and $R^i_{f,k,t}$ allows us to capture cross-sectional variation in exposure to global shocks, and cross-sectional differences in country-specific growth expectations. We also include the simulation-specific parameters γ_k and $\frac{1}{\psi_k}$ in order to capture their impact on the equilibrium policy functions.

In our interpolation, we include all first- and second-order terms of the above-mentioned variables and form a vector with 35 entries, $\Omega^j_{k,t}$, for each observation t and simulation k (7 linear terms, 7 square terms, and 21 cross-product terms). After we include an intercept, Γ^0, our equilibrium SDF is approximated by the following expression:

$$m^i_{k,t+1} = \Gamma^0 + \sum_{j=1}^{35} \Gamma^j \Omega^j_{k,t+1},$$

which gives us a mapping between observable variables, preference parameters to be estimated, and model-implied SDF dynamics. The approximation fit is very good, with an R^2 of 99%. After performing this regression, we discard 17 terms with very negligible fitting power. For numerical tractability, we use only the remaining 18 terms and retain an R^2 of 98.9%.

After these two steps, we plug our empirical real risk-free rates, our estimated fundamental endowment shocks, and our estimated endowment βs (see section 2 of the paper) into the expression above so that it becomes a function of (γ, ψ) only:[9]

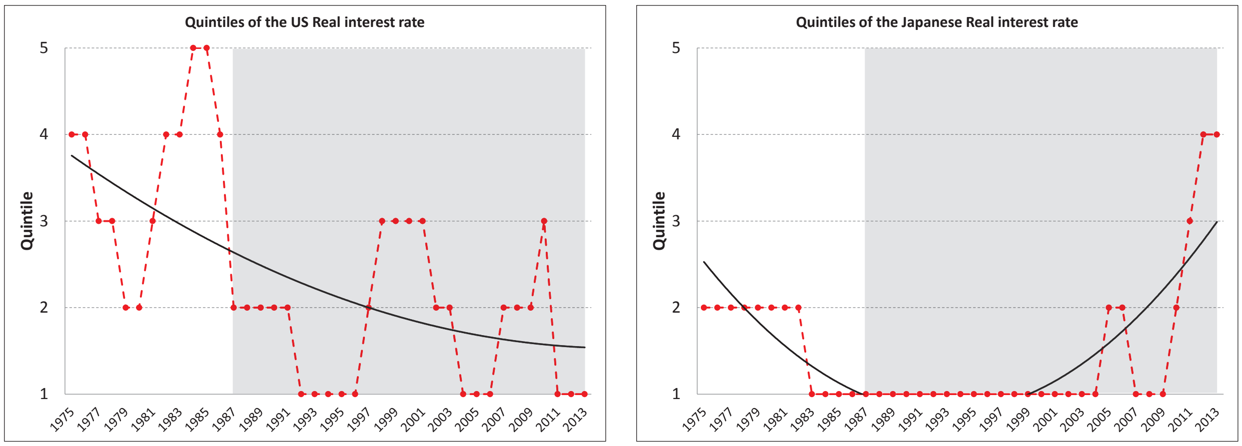
$$m^i_t = F(\psi, \gamma | r^i_{f,t}, \beta^i, \varepsilon_{global,t+1}, \varepsilon^i_{t+1}, \varepsilon^i_{z,t+1}).$$

The resulting GMM estimation is computationally standard and efficient.

G Additional Empirical and Model Results

Distribution of real risk-free rates. In figure G2, we report the quintiles of the cross-sectional distribution to which the US and the Japanese real interest rates belonged from 1975 to 2013. Specifically, for each G-10 country we compute its average

[9]Consistent with our highly persistent processes for β^i_t, we use our constant empirical estimates of β^i as the observable analog.

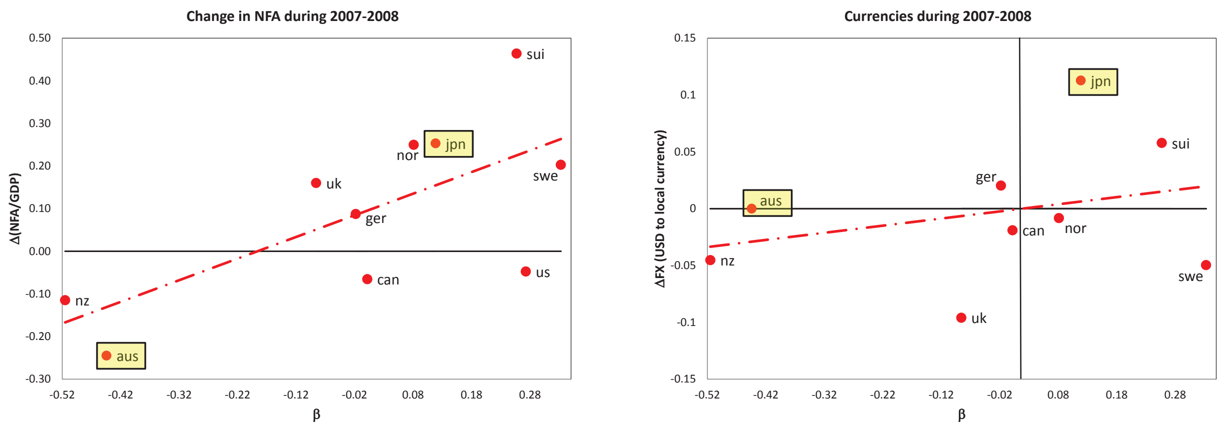


(a) (b)

FIG. G2 - Panel (a) (panel (b)) reports the quintiles to which the US (Japanese) real interest rate belongs during the period 1975–2013 (dashed-dot line). In each panel, the solid line represents a quadratic fit of the time series of quintiles, and a value of 1 (5) means that the interest rate is in the bottom (top) 20% of the cross-sectional distribution for G-10 countries. The grey area denotes the sample that we consider in our empirical analysis.

real interest rate using overlapping 5-year rolling windows. This procedure allows us to create a balanced panel of annual real rates for our G-10 countries. In each year, we compute the cross section of our 10 real rates and identify the quintile corresponding to the US and the Japanese real interest rates, respectively. A value of 1 (5) means that the interest rate is in the bottom (top) 20% of the cross-sectional distribution.

The US real rate was on the high side of the cross-sectional distribution in the first part of the post-Bretton-Woods era, but it has moved toward the low side over time (panel (a) of figure G2). An econometrician focusing on the entire sample would conclude that the US has, on average, a median interest rate, whereas someone looking at our post-1987 sample would conclude that the US is a below-median country. For Japan, the real interest rate was firmly in the bottom part of the cross-sectional distribution until the early 2000s, but sharply increased in the most recent part of the sample. In our post-1987 sample, the Japanese interest rate has on average been



(a) (b)

FIG. G3 - Panel (a) reports the change in the average NFA position of each country during 2008–2009 over the average of the previous two years against the β of the corresponding country. The estimated βs are reported in Table 1. Panel (b) reports the average change in exchange rate of each country vis à vis the US dollar during 2008–2009 against the β of the corresponding country. The estimated βs are reported in table 1. All changes are in excess of the appreciation rate of the Australian dollar (-1.5% in our sample). The reported value for AUS is therefore equal to zero by construction.

low, but it has not always been in the lowest quintile in the cross section of our G-10 countries.

Exposures and international capital flows. We explore the relation between the cross section of βs and international capital flows during the 2007–2008 Great Recession. This period is characterized by severe negative global news shocks that affected all countries in our cross section. The year 2008, in particular, features the lowest realization of our global long-run risk variable. Our model would predict that upon the occurrence of a negative shock, negative β countries (i.e., the countries that are less exposed to global shocks) should provide insurance to the countries with large positive βs (i.e., the countries with a large degree of exposure to global shocks). That is, countries with lower exposure should experience a deterioration of their NFA positions. Our results provide empirical support for this prediction by showing that in 2008-2009, Australia and New Zealand experienced a large drop in their net foreign asset positions, while countries on the other end of the spectrum, such as Japan, Swe-

TABLE G10: Portfolio βs of Conditional Analysis

	Exchange Rate				Net Foreign Assets					
Portfolio	[1]	[2]	[3]	[4]	[5]	[6]	[7]	[8]	[9]	[10]
Low β_i^z	-0.48	-0.48	-0.48	-0.48	-0.48	-0.48	-0.48	-0.48	-0.48	-0.48
Medium β_i^z	0.02	-0.01	0.02	-0.01	0.02	0.02	-0.01	0.02	0.02	-0.01
High β_i^z	0.30	0.30	0.30	0.30	0.29	0.30	0.29	0.29	0.30	0.29

Notes - This table reports the βs of the portfolios associated with the analysis in table 4 of the main text.

den, and Switzerland, experienced a large accumulation of foreign assets (see panel (a) in figure G3).

Furthermore, the currencies of countries with large positive βs appreciated relative to those of countries with negative βs during the Great Recession (see panel (b) in figure G3). Take, for example, Japan (a positive β country) and Australia (a negative β country): during 2008–2009, the Japanese yen appreciated by an average of about 10% per year relative to the Australian dollar. This is exactly what our model predicts, as shown in figure 2.

Portfolio exposures. Table G10 reports the βs of the portfolios used to carry out the analysis in table 4 of the main text.

Estimation of p/d ratio exposures in the model. In this section we document the ability of our estimation technique to recover the exposure to p/d ratios in the model. Specifically, we simulate 1,000 samples of 424 months for the series of interest in our analysis: Δy_t^i and pd_t^i, $\forall i = \{1, ..., 5\}$. We use the first 100 months as a burn-in period to obtain the required spread of β exposures across countries. Then we proceed with the time aggregation of the relevant variables to annual frequency. For GDP, we initialize each country with a value of $Y_1 = 1$ and then use the simulated path of log GDP growth, Δy_t, to obtain monthly levels of GDP, $Y_t, \forall t$. Lastly, for each year in our sample, we obtain the annual level of GDP by taking the sum of 12 consecutive

TABLE G11: Estimating Exposures from Model Simulations

Portfolio	1	2	3	4	5
Calibrated β^i	0.68	0.34	0.06	-0.23	-0.68
Estimated β^i (simulations)	0.54	0.27	0.03	-0.21	-0.62
[90% C.I.]	$[\ 0.02, 1.04\ ]$	$[\ 0.00, 0.05\ ]$	$[\ -0.11, 0.15\ ]$	$[\ -0.44, 0.05\ ]$	$[\ -1.09, -0.01\ ]$

Notes - The first row (labeled "Calibrated β^i") reports the initial loadings on the global long-run risks in our simulations. The second row (labeled "Estimated β^i") reports the loadings of each portfolio on the global long-run risk estimated as in table 1, obtained as the average across 1,000 simulated samples of size 27 years. The last row reports the 90% confidence intervals of the estimated point estimates.

monthly observations. For the p/d ratio we proceed as follows. We initialize each country with a value of 1 of the dividend at time 1, $D_1 = 1$, and then construct its annual time series in a way analogous to what we do for the GDP. Then we multiply each monthly observation of the p/d ratio by its monthly level of dividend, obtaining a monthly time series of the stock price. Last, we divide the end-of-year stock price by the sum of the dividends paid by the stock in the previous 12 months.

We report the results of this exercise in table G11. The range of estimated β exposures that we obtain from our simulations are in line with what we estimate from the data (see table 1). Furthermore, we always reject the null that the country with the most negative (positive) exposure has a positive (negative) estimated coefficient, thus confirming the ability of our approach to consistently identify the key countries in our long-short portfolios.