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# Currency Risk Factors in a Recursive Multicountry Economy

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## ABSTRACT

Focusing on the 10 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, heterogeneous 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, Roussanov, and Verdelhan and Della Corte, Riddiough, and Sarno.

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, Roussanov, and Verdelhan (2014), Della Corte, Riddiough, and Sarno (2016)). 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, heterogeneous exposure to global growth news shocks is a key driver of

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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), Bansal and Shaliastovich (2013)). We expand and generalize this setting in at least two relevant directions.

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

Second, we introduce heterogeneous exposure to growth news shocks in the cross section of countries in a way that is consistent with our novel empirical evidence on the 10 countries with the most traded currencies (henceforth, G-10 countries). Specifically, we model 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, Croce, and Liu (2018)). In finite samples, however, our countries feature substantial heterogeneity, consistent with the empirical investigation of Lustig, Roussanov, and Verdelhan (2011) and Hassan and Mano (2014). These heterogeneous loadings are a reduced-form way of capturing a mix of fundamental differences across countries, such as size (Hassan (2013)), commodity intensity (Ready, Roussanov, and Ward (2017)), monetary policy rules (Backus et al. (2010)), and financial development (Maggiori (2017)).

Under our benchmark calibration, we are able to produce an average High Minus Low (HML) annual spread of about 3%, which is 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 (NFA) positions and exposure to long-run global growth news. This suggests that the factors proposed by Della Corte, Riddiough, and Sarno (2016) and Lustig, Roussanov, and Verdelhan (2011) may be the risk-sharing outcome of a single fundamental source of heterogeneity, namely, their different exposure to global long-run growth news.

To discipline our calibration, we use macroeconomic data from the 10 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 Gross Domestic Product (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 take the estimated coefficients to be a measure of each country's exposure to global long-run risk. We find a substantial degree of heterogeneity in the estimated exposures.[1]

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, while 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 do not identify any heterogeneity in terms of the exposure to global short-run risk, which implies that abstracting 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 in their NFA positions, whereas highly exposed countries should retain a larger balance of NFA.

We provide empirical support for this prediction by showing that countries like Australia and New Zealand experience a large decrease (increase) in their NFA positions in times of negative (positive) long-run global growth prospects, while countries on the other end of the spectrum, such as Switzerland, experience a large inflow (outflow) of assets from abroad when global long-run growth prospects are weak (strong). Furthermore, the currencies of high-exposure

[1] We note that the estimation of these exposures is identical to regressing each country's p/d ratio onto the cross-sectional average p/d ratio. However, our first-step regressions are key to documenting that p/d ratios have predictive power for macroeconomic fundamentals, and thus to supporting the long-run risk model interpretation that we offer in this paper.

countries have a tendency to appreciate in bad times relative to the currencies of low-exposure countries, a finding that is also consistent with our model.

To push our analysis one step further, we estimate the preference parameters using Euler equation restrictions on currency returns, equity returns, and risk-free rate, as well as 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 Generalized Method of Moments (GMM) estimation exercise. The estimates for the preference parameters are in line with our benchmark calibration, and we find that we cannot reject the model in any of the cases that we consider. In contrast, we can strongly reject the model with constant relative risk aversion (CRRA) preferences.

Our analysis helps 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, Roussanov, and Verdelhan (2011), Zviadadze (2013), and Farhi et al. (2009). 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, Farhi, and Gourinchas (2008) and Gourio, Siemer, and Verdelhan (2014)). Furthermore, our benchmark model with heterogeneous exposure to growth news is consistent with Verdelhan (2018), as it enables global long-run shocks to contribute to bilateral exchange rate variance.

Our study is also related to the growing literature that investigates the macroeconomic foundations of international financial market fluctuations (see, e.g., Verdelhan (2010), Stathopoulos (2012), Hassan (2013), Heyerdahl-Larsen (2014), Farhi and Gabaix (2015)). Our empirical evidence on the 2007 Great Recession is related to work on rare disasters (Barro (2006), Gabaix (2012), Gourio (2012)) and its applications to international finance (see, e.g., Gourio, Siemer, and Verdelhan (2014)).

Several articles have documented limitations of the long-run risks model in a one-country setting (see, e.g., Le and Singleton (2010), 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. Specifically, our equilibrium exchange rates are too volatile, and our interest rates too smooth, both in the time series and in the cross section. As a result, we have an excessive manifestation of the forward premium anomaly. For many variables of interest, we can replicate their coefficient of variation in the cross section but we cannot fully replicate their cross-country spread. Furthermore, our benchmark calibration relies on a high cross-country correlation of long-run growth. The introduction of frictions (see, e.g., Froot and Stein (1991), Farhi and Werning (2014), Gabaix and Maggiori (2015)) may resolve these limitations, and we regard this as an important direction for future research.

The paper is organized as follows. Section I reports our empirical evidence concerning the heterogeneous exposure to global long-run risk in G-10 countries. In Section II, we present our risk-sharing model and its equilibrium conditions. Section III presents our main results, and Section IV concludes.

## I. Empirical Analysis

### *A. 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 demonstrate 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 heterogeneous exposure to global risk in the presence of complete financial markets.

Data on consumption and net exports come from the Organisation for Economic Co-operation 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, we proxy for total endowment growth using $\Delta GDP_t$ , which we compute 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 expenditures from our empirical measure to be consistent with our endowment economy, in which we abstract from both physical investment and demand for public goods. Data on NFA come from the updated and extended version of the data set constructed by Lane and Milesi-Ferretti (2007).[2]

Data on interest rates and inflation also come from the OECD. Real interest rates are constructed using three-month interbank rates from the OECD, adjusting for realized inflation using the annual changes in Consumer Price Index (CPI). Norwegian interest rates prior to 1979 are obtained from Stats Norway. Swedish three-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.[3]

All exchange rate data are collected relative to the U.S. dollar (USD). Exchange rate data for Australia, Canada, Japan, New Zealand, Norway, Sweden, Switzerland, and the United Kingdom 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 to 2014. Real exchange rates are obtained by dividing by the relative CPI index of the corresponding country.

[2] These data are graciously provided on Philip Lane's website: http://www.philiplane.org/EWN.html.

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

### B. Estimation Procedure

We follow Colacito and Croce (2013) and Bansal, Kiku, and Yaron (2008) and identify short- and long-run innovations to GDP growth rates using predictive regressions. Specifically, we interpret the projection of $\Delta GDP_t$ in each country onto lagged values of that country's p/d ratio as our measure of long-run risk. We have a balanced panel of p/d ratios from 1987 to 2013, so we use data from this period to estimate global risk exposure. The estimation of all the parameters of interest is obtained from the following specification:

$$\Delta GDP_t^i = \phi \cdot pd_{t-1}^i + \sigma \cdot \underbrace{\varepsilon_t^i}_{\text{Short-Run Shock}} \tag{1}$$
$$z_t^i = \rho_z \cdot z_{t-1}^i + \varphi_e \cdot \sigma \cdot \underbrace{\varepsilon_{z,t}^i}_{\text{Long-Run Shock}}, \forall i \in \{\text{G-10 countries}\},$$

where $z_t^i = \phi \cdot pd_t^i$ is our measure of long-run risk in country $i$. Note that we omit the intercepts because all variables are demeaned, and we pool the estimation of the parameters $(\phi, \sigma, \rho_z, \varphi_e)$ for parsimony.[4] The system in (1) is estimated using GMM. Standard errors take into account both serial and cross-sectional correlation.

Table I, Panel A, reports the estimation results. Several interesting findings emerge. First, all estimated parameters are statistically significant, which supports the relevance of lagged p/d ratios as a source of predictability for macroeconomic variables' future growth rates. 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 GDP growth. 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 $\varphi_e$ 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.

### C. Exposure to Global Risk

In Panel B of Table I, we report the exposure of each country's endowment to a measure of global endowment risk. Specifically, we regress each country's GDP growth rate, $\Delta GDP_t^i$, on the average of all countries' GDP growth rates:

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

With the sole exception of Norway and Canada, we cannot reject the null hypothesis that the estimated coefficients in Panel B of Table I are equal to

[4] We adopt this pooled specification because we cannot reject the null hypothesis that all the country-specific $\phi$ coefficients are equal to each other.

**Table I**

## Dynamics of Endowments and Predictive Components

Panel A reports estimates for 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 I. Panel B reports estimates for the exposure of each country's GDP growth rate to the global GDP growth rate (see equation (2)). The sample is 1970 to 2013. Panel C reports estimates for the exposure of each country's predictive component of GDP to the global predictive component (see equation (3)). The sample is 1987 to 2013. The numbers in square brackets are the $p$-values associated with the null hypothesis that the estimated exposure in the first half of the sample (1987 to 2000) is different from the estimated exposure in the second half of the sample (2001 to 2013). In all panels, the numbers in parentheses are heteroskedasticity-adjusted standard errors. $^{*}$, $^{**}$, and $^{***}$ denote statistical significance at the 10%, 5%, and 1% levels, respectively.

Panel A: Estimation of Predictive Components

| | $\phi$ | $\rho_x$ | $\sigma$ | $\varphi_e$ |
|---|---|---|---|---|
| Parameters | 0.005*** | 0.773*** | 0.020*** | 0.058*** |
| (*SE*) | (0.000) | (0.006) | (0.000) | (0.001) |

Panel B: Exposure to Global Endowment Risk

| | NZ | AUS | UK | GER | CAN | NOR | JPN | SUI | USA | 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 |
| (*SE*) | (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 | USA | SWE |
|---|---|---|---|---|---|---|---|---|---|---|
| $\beta^i$ | −0.51*** | −0.44*** | −0.08 | −0.02 | 0.00 | 0.08 | 0.12 | 0.26** | 0.27* | 0.33** |
| (*SE*) | (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] |

0. 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. Here, 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 of all $z_t^i$'s:

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

We estimate these coefficients using the longest common sample across all countries, 1987 to 2013, and report our results in Panel C of Table I. We find a sizeable degree of dispersion in the estimated $\beta^i$'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. We also report the $p$-values associated with a Chow test for the null hypothesis that the $\beta^i$'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 United States 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 United States is an average country in the cross-sectional distribution of real interest rates. Specifically, the United States had a relatively high real rate in the first part of the post-Bretton Woods era (see Panel A of Figure IA.2 in the Internet Appendix), but not in the more recent sample period that we consider.[5] Equivalently, an econometrician focusing on the entire sample would conclude that the United States has been on average a median interest-rate country, whereas someone looking at the post-1987 sample would conclude that the United States is a below-median country.

Similar considerations apply to Japan. Table I indicates that Japan does not exhibit the highest exposure to global long-run risk, a finding that 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 IA.2 in the Internet Appendix).

Our empirical findings could, in principle, be generated by different models. In the following 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 III.D

[5] The Internet Appendix is available in the online version of the article on *The Journal of Finance* website.

and document the ability of this model to match several key moments in the cross section of international asset prices and macroeconomic quantities.

### D. 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 United States and other major industrialized countries. Given the difficulty of identifying a reliable data set on total nonfinancial corporate distributions for our large cross section of countries, we provide additional evidence using Tobin's $Q$, a forward-looking variable that does not require the use of dividends. In Section I of the Internet Appendix, we document in detail that using Tobin's $Q$ to forecast long-run risk confirms our results about heterogeneous 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.

## II. 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.

### A. 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

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

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 $\alpha$ is larger than 0.5. This allows us to build consumption home bias into the model.

### B. Preferences

As in Epstein and Zin (1989), agents' preferences are recursive but not time separable:

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

The coefficients $\gamma$ and $\psi$ measure the degree of relative risk aversion (RRA) and 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 preferences, we note that the ordinally equivalent transformation

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

can be approximated as

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

where $\theta \equiv \frac{\gamma-1/\psi}{1-1/\psi}$. Note that the sign of $(\frac{\theta}{E_t[V_{t+1}]})$ 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 $\theta$ is positive: uncertainty about continuation utility reduces welfare and generates an incentive to trade off future expected utility, $E_t[V_{t+1}]$, for future utility risk, $\mathrm{Var}_t[V_{t+1}]$. 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 with 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 trade-off of wealth. For this reason, in what follows we use the terms "wealth" and "continuation utility" interchangeably.

### C. 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,

$$z_{i,t} = \rho_i z_{i,t-1} + \varepsilon^z_{i,t}, \quad \forall i \in \{1, 2, \dots, N\}. \tag{8}$$

Throughout the paper, we refer to $\varepsilon^z_{i,t}$ as long-run shocks, due to their long-lasting impact on the growth rates of the endowments. Similarly, we call $\varepsilon^X_{i,t}$ 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.

### D. Heterogeneous 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,

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

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

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

The volatilities of $\epsilon^z_{\text{global},t}$ and $\epsilon^{\tilde{z}}_{i,t}$ are set to replicate both the unconditional standard deviation and the 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 $\epsilon^{\beta,z}_{i,t} \sim \text{i.i.d.}N(0, \sigma_{\beta,z})$. These shocks are both long-lived ($\rho^\beta_z \approx 1$) and uncorrelated with 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 of modeling country-specific exposure to shocks produces two benefits. First, it enables us to study an economy with ex ante symmetrically calibrated countries for which a well-defined equilibrium exists (Colacito, Croce, and Liu (2018)). Under the assumption of a permanent degree of heterogeneous exposure, there would be a degenerate ergodic distribution of the equilibrium variables.[6] Second, after simulating a history of heterogeneous exposure shocks

[6] 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).

($\epsilon_{i,t}^{\beta,z}$), we are able to study the characteristics of a cross section of countries that remain substantially heterogeneous in finite samples. We think of the $\beta_{i,t}^{z}$ coefficients as devices to capture the heterogeneity documented by Lustig, Roussanov, and Verdelhan (2011), Backus et al. (2010), Ready, Roussanov, and Ward (2017), and Hassan and Mano (2014) in a parsimonious, reduced-form manner.

### D.1. 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 markets and goods markets are assumed to be frictionless, the budget constraint of our agents can be written as

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

where $p_{i,t}$ denotes the price of good $i$ relative to that of good 1, $A_{i,t}(\zeta^{t})$ 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 $\zeta^{t+1}$ at time $t+1$. In equilibrium, the market for international state-contingent claims clears, implying that $\sum_i A_{i,t} = 0, \forall t$.

### D.2. 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 $j \in \{1, \ldots, N\}$, that maximize

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

subject to the economy-wide feasibility constraints

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

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 account for the nonseparability of the utility functions. In the Internet Appendix, we show that the first-order necessary conditions imply the allocations

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

$$x_{i,t}^{j} = \frac{1-\alpha}{\alpha} \frac{1}{N-1} \frac{S_{j,t}}{S_{i,t}} x_{i,t}^{i}, \quad \forall i \neq j \in \{1, 2, \ldots, N\},$$

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.

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

### D.4. Bilateral Imports and Exports

At each point in time, the exports of country 1 to 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 bilateral net exports scaled by total output are equal to

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

**Table II**
**Calibration**

This table reports our benchmark monthly calibration. Under the homogeneous exposure case, $\beta^z_{i,t} = 1\,\forall i, \forall t$.

| Description | Values |
|---|---|
| Preferences: | |
| Relative risk aversion [$\gamma$] | 6.50 |
| Intertemporal elasticity of substitution [$\psi$] | 1.60 |
| Subjective discount factor [$\delta^{12}$] | 0.98 |
| Degree of home bias [$\alpha$] | 0.98 |
| Endowments: | |
| Mean of endowment growth (%) [$12\mu$] | 2.00 |
| Short-run risk volatility (%) [$\sigma\sqrt{12}$] | 1.87 |
| Long-run risk volatility (%) [$\sigma_z/\sigma$] | 6.00 |
| Long-run risk autocorrelation (%) [$\rho^{12}$] | 0.78 |
| Cross-correlations of long-run shocks [$\rho_z$] | 0.93 |
| Cointegration speed [$(1-\tau)^{12}$] | 0.99 |

| | Exposure | |
|---|---|---|
| | Homogeneous | Heterogeneous |
| Cross-correlations of short-run shocks [$\rho_X$] | 0.20 | 0.40 |
| Orthogonalization: | | |
| Volatility of global long-run shocks (%) [$\sigma_z^{global}/\sigma$] | – | 5.80 |
| Volatility of local long-run shocks (%) [$\sigma_z^{global}/\sigma$] | – | 1.60 |
| Time-varying exposure: | | |
| Autocorrelation of $\beta^z_{i,t}$ [$(\rho_z^{\beta})^{12}$] | – | 0.99 |
| Volatility of shocks to $\beta^z_{i,t}$ (%) [$\sigma_{\beta,z}$] | – | 0.05 |

and

$$\frac{NX^j_{1,t}}{X_{1,t}} = \frac{(1-\alpha)\cdot\left(S^j_t - 1\right)}{\alpha(N-1) + (1-\alpha)\sum_{j\neq 1} S^j_t}, \tag{16}$$

respectively. Detailed derivations are reported in the Internet Appendix.

### *E. Calibration and Solution Method*

We detail our baseline monthly calibration in Table II. 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, Kiku, and Yaron (2008), Colacito and Croce (2013)). We set the home bias parameter ($\alpha$) 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 Section V of the Internet Appendix, we conduct a sensitivity analysis with respect to this parameter. We find a tension between home bias and the magnitude of risk premia. In particular, in the absence of other trade impediments, a low home bias generates modest currency risk premia.

The parameters governing the dynamics of the endowments' growth rates 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 such that they are consistent with the median values in our data set, which we discuss 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 heterogeneous exposure, this correlation is set to 0.40, a number that falls in the middle of the correlation range estimated among our 10 countries. When we consider the special setting without heterogeneous 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.[7] Both the amount of long-run risk ($\sigma_z/\sigma$) and the persistence of the long-run components ($\rho$) are consistent with the empirical estimates reported in Table I.

We set $\rho_z^\beta = 0.999$ 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 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 $\tau$ so as to have an almost undetectable extent of cointegration, consistent with Rabanal, Rubio-Ramirez, and Tuesta (2011).[8] We present a sensitivity analysis with respect to this parameter in Section V of the Internet Appendix.

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, Croce, and Liu (2018) and 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

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

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

Section III of the Internet Appendix and show that it features very small approximation errors. All variables included in our dynare++ code are expressed in log-units.

## III. 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. In Section III.A, 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 III.B by introducing heterogeneous exposure to global long-run shocks. In Section III.C, we present novel empirical evidence that supports the distinctive channels at work in our model.

### *A. Homogeneous Exposure*

We set the number of countries in our model to 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 III. In Section IV of the Internet Appendix, we show that most of these moments are not sensitive to the number of countries, which implies 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 inspection of the model's mechanism with homogeneous exposure to Section IV of the Internet Appendix. 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. Because of the presence of highly cross-country-correlated long-run growth news, SDFs 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, Cochrane, and Santa-Clara (2006) puzzle.

This setting produces low and smooth risk-free rates. Across countries, the risk-free rates are as highly correlated as the SDFs. 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 Colacito and Croce (2013), Gourinchas and Rey (2014)). In Section IV of the Internet Appendix, we show that the model can also account for the almost complete lack of correlation between consumption growth differentials and exchange rate fluctuations (see, e.g., the Backus and Smith (1993)

**Table III**

## Simulated Moments with Heterogeneous Exposure

The table reports both empirical moments computed using the data set described in Section I and simulated moments from the model with both heterogeneous and homogeneous exposure. All parameters are set to their benchmark values reported in Table II. For the CRRA case, we set $\gamma = 1/6.5$. Panel A reports the moments for the dynamics of exogenous endowment growth rates. Panel B reports the moments of the consumption growth rate within each country. Panel C reports the cross-country moments for each country pair. Panel D reports the median moments for the risk-free rates ($r_f$), SDFs ($M$), NFA-to-output ($NFA/X$), slope coefficient of the UIP regressions ($\beta_{UIP}$), and average currency risk premium (E[$HML$]). In Panel E, we report cross-sectional standard deviations for the listed moments. CoV denotes the cross-sectional coefficient of variation.

| | | | Homogeneous | Heterogeneous | |
|---|---|---|---|---|---|
| | Data | *SE* | EZ | EZ | CRRA |
| Panel A: Endowment Growth | | | | | |
| Std($\Delta x$) | 2.10 | 0.26 | 1.93 | 1.95 | 1.95 |
| ACF$_1$($\Delta x$) | 0.21 | 0.13 | 0.29 | 0.30 | 0.35 |
| corr($\Delta x_t^h$, $\Delta x_t^f$) | 0.23 | 0.06 | 0.43 | 0.40 | 0.40 |
| Panel B: Single-Country Moments | | | | | |
| Std($\Delta c$) | 1.91 | 0.25 | 1.78 | 1.96 | 1.74 |
| ACF$_1$ ($\Delta c$) | 0.46 | 0.11 | 0.31 | 0.28 | 0.30 |
| Panel C: Bilateral Moments | | | | | |
| corr($\Delta c_t^h$, $\Delta c_t^f$) | 0.24 | 0.05 | 0.55 | 0.38 | 0.59 |
| Std ($\Delta e$) | 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 |
| Panel D: 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($\Delta x$) | 18.58 | 2.95 | 11.34 | 25.76 | 10.29 |
| ACF$_1$($NFA/X$) | 0.99 | 0.05 | 0.81 | 0.88 | 0.74 |
| $\beta_{UIP}$ | −0.94 | 0.48 | −5.54 | −4.62 | 0.78 |
| E($HML$) | 3.20 | 1.10 | 0.11 | 3.01 | 0.13 |
| Panel E: Cross-Sectional Standard Deviation | | | | | |
| Std($\Delta c$) | 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($\Delta x$) (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 |
| $\beta_{UIP}$ (CoV) | 0.87 | 0.29 | 1.30 | 1.16 | 0.58 |
| Std ($\Delta e$) (CoV) | 0.21 | 0.04 | 0.03 | 0.41 | 0.04 |

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 ($\beta_{UIP}$). 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. In 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 $\beta_{UIP}$ is excessively negative.

### *B. 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 heterogeneous exposure to endowment shocks (Lustig, Roussanov, and Verdelhan (2011), Hassan and Mano (2014)), (ii) NFA positions (Della Corte, Riddiough, and Sarno (2016)), or (iii) the level of their risk-free rate (Lustig, Roussanov, and Verdelhan (2011)).

*Conditional Heterogeneity:* Under the null of the model, the coefficients $\beta^i$ 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

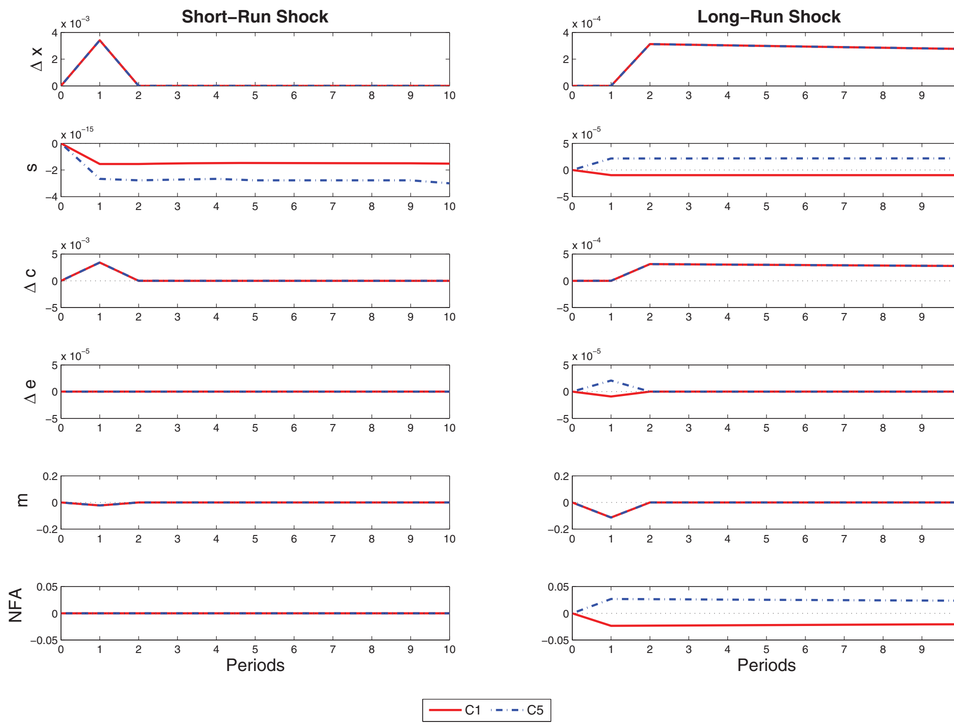


**Figure 1. Impulse response functions under heterogeneous wealth.** The left (right) panels report the response of endowment growth ($\Delta \log X^i$), relative Pareto weights with respect to country 3 ($\log S_j/S_3$), consumption growth ($\Delta c_i$), exchange rate growth ($\Delta e^i_3$), SDFs ($m_i$), and NFA ($A_i/X_i$) to a one-standard-deviation short-run (long-run) global shock. All panels correspond to the case in which the economy consists of five countries ($i = 1, \ldots, 5$). The exchange rate is measured with respect to country 3, which implies 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. (Color figure can be viewed at wileyonlinelibrary.com)

plausible range given our results in Panel B of Table I. All other countries have an exposure coefficient equally spaced in this range.[9]

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.

[9] In Section I of the Internet Appendix, we show that estimating the $\beta^i$ coefficients using long-run risk innovations, as opposed to levels, yields similar results. In Section VII of the Internet Appendix, we show that when we apply our estimation procedure to simulated data, we can reproduce the cross section of the calibrated $\beta^i$ 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 a lower interest rate for country 1. 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, Riddiough, and Sarno (2016)). 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):

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

where $\mathcal{M}_{t+k} = \prod_{l=0}^{k} Q_{t+l}$ and $NX_{i,t+k} = \sum_{j=1}^{N} p_{j,t} x_{i,t}^{j} - p_{i,t} X_{i,t}$, $\forall i \in \{1, \dots, N\}$. 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 can think of countries 1 and 5 as Switzerland and Australia, respectively, as 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 SDF 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,

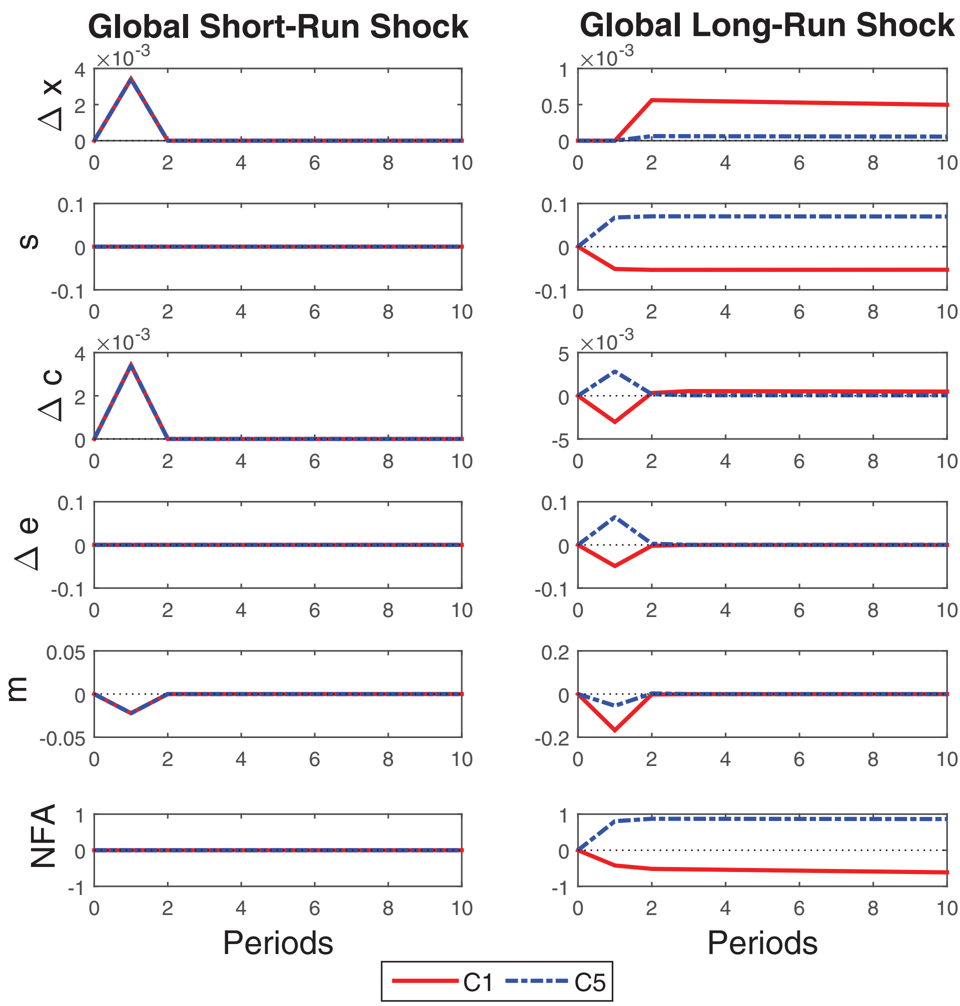


**Figure 2. Impulse response functions under heterogeneous exposure.** The left (right) panels report the response of endowment growth ($\Delta \log X^i$), relative Pareto weights with respect to country 3 ($\log S_j/S_3$), consumption growth ($\Delta c_i$), exchange rate growth ($\Delta e_3^i$), SDFs ($m_i$), and NFA ($A_i/X_i$) to a one-standard-deviation short-run (long-run) global shock. All panels correspond to the case in which the economy consists of five countries ($i = 1, \dots, 5$). 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 ($-0.65$). (Color figure can be viewed at wileyonlinelibrary.com)

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 previous 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 C 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 D). This 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 E). The implied average currency returns have an annual spread of about 3% (bottom panel), consistent with the estimated unconditional HML-FX of Lustig, Roussanov, and Verdelhan (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 III, we report simulated moments from our model with and without heterogeneous exposure and recursive preferences. In all cases, we report median values from our cross section of five countries. In the absence of heterogeneous exposure (first column),

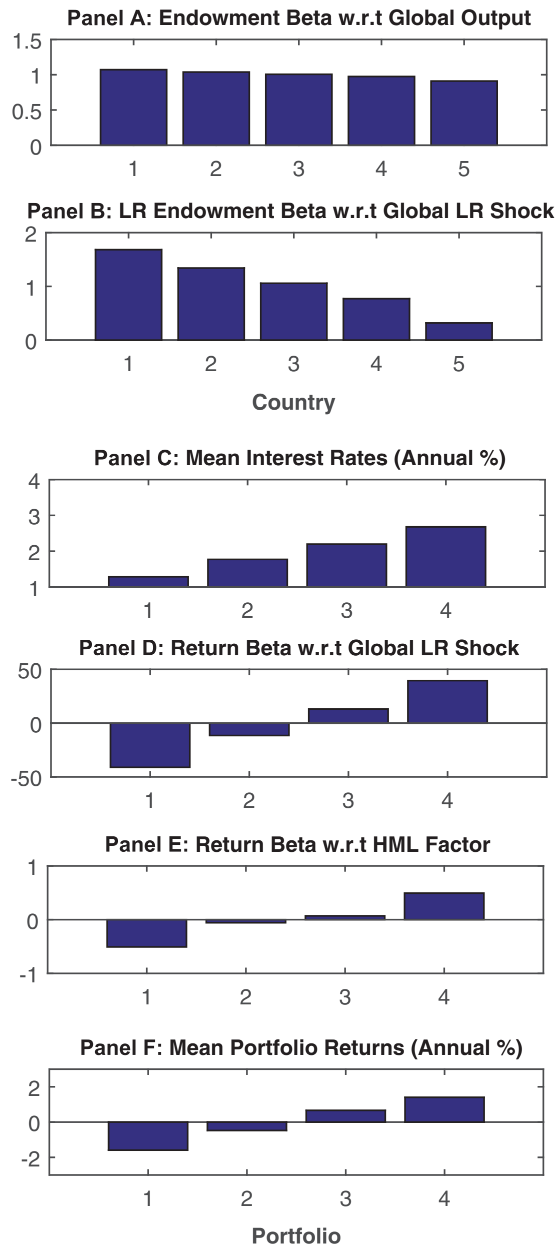


**Figure 3. Cross-sectional risk (EZ case).** The top panel shows the exposure of our five countries to global output $(1+\beta^i_{\Delta y})$ as defined in equation (2). The second panel depicts the five countries' exposure to the global long-run component $(1+\beta^z_i)$. The bottom four panels show 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 of the parameters are calibrated to the monthly values reported in Table II. Statistics are the averages across 300 simulations of 250 monthly observations. (Color figure can be viewed at wileyonlinelibrary.com)

the moments are identical to those obtained in the last column of Table IA.VII in the Internet Appendix.

All moments remain basically unchanged across columns, implying that the introduction of heterogeneous 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 $\beta_{UIP}$ is excessively negative.

Second, heterogeneity introduces more cross-country variation in the exposure to shocks and hence reduces the correlation of the SDFs. 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.

Finally, heterogeneity increases risk-sharing motives and makes the adjustment of both NFA positions and 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, especially net exports.

*The Role of Heterogeneity for Cross-Sectional Moments:* In the bottom panel of Table III, we focus on key cross-sectional moments of both quantities and prices. First, 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 $\beta_{UIP}$, we analyze whether the model can reproduce the coefficients of variation in our empirical cross section. Under homogeneous exposure, the result is negative. With heterogeneous 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 reporting 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 III show 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 produce no immediate movements in the marginal utilities of our

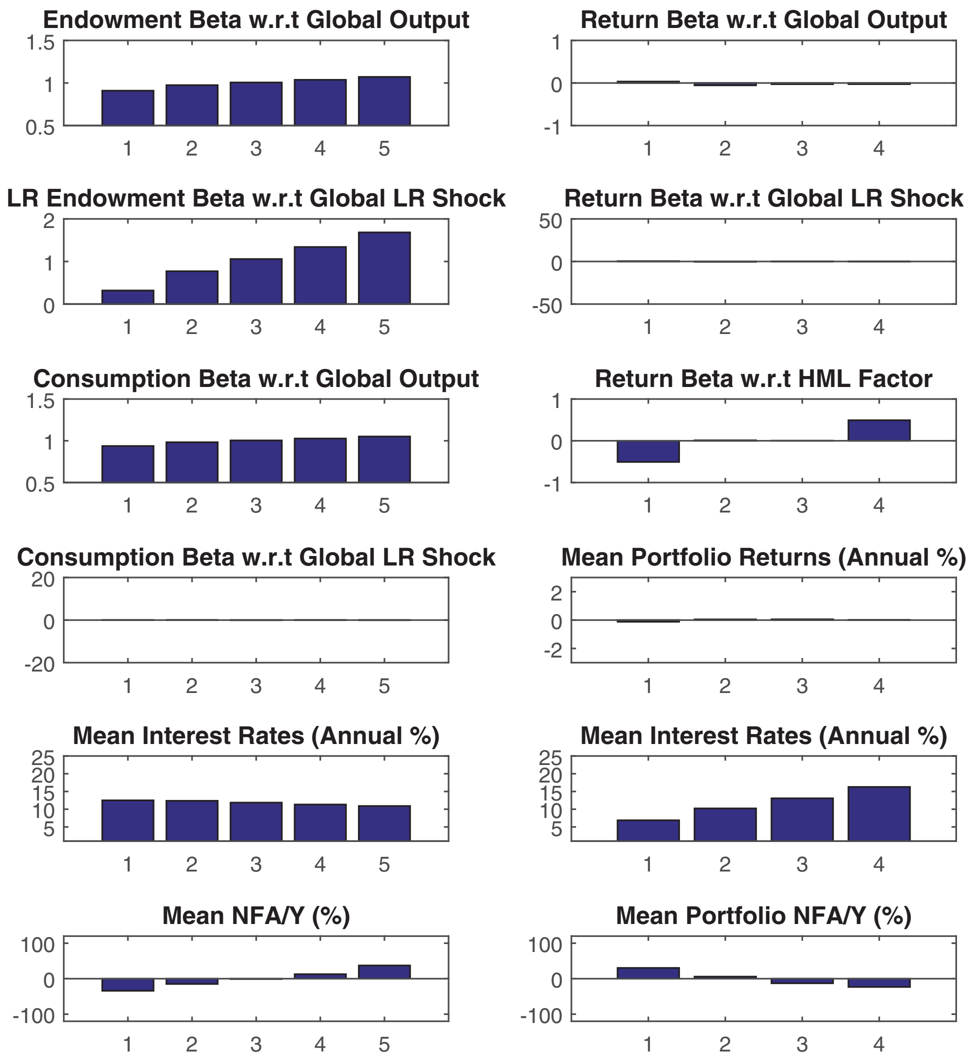


**Figure 4. Cross-sectional risk (CRRA case).** The panels on the left show the cross-sectional characteristics of five countries sorted on their exposure to the global long-run shock $(1 + \beta_i^z)$. The panels on the right show the cross-sectional characteristics of four portfolios formed by sorting our countries each period on their lagged interest rate differential with respect to a fixed numeraire country. All of the parameters are calibrated to the monthly values reported in Table II, 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). (Color figure can be viewed at wileyonlinelibrary.com)

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 heterogeneous.

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 III 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 the 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.

### *C. 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 U.S.-based investor is close to flat 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 shows 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,

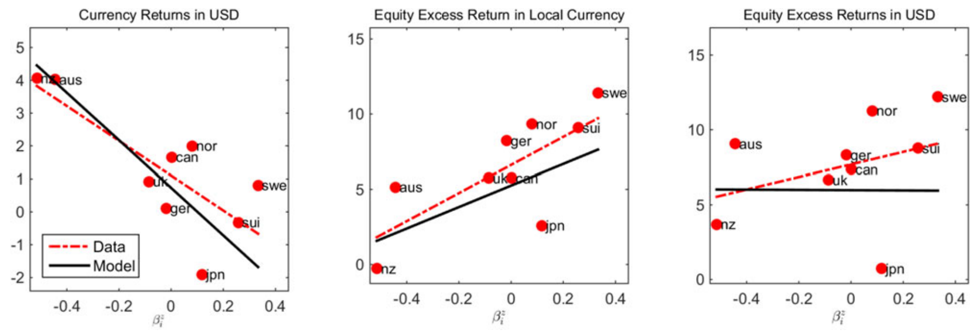


**Figure 5. Equity and FX excess returns.** In each panel, the horizontal axis corresponds to the estimated exposure coefficient $\beta_i^z$. For each country in our data set, we use the estimates provided in Table I. The left panel links exposure to global long-run risk to average currency returns expressed in USD. The middle (right) panel corresponds 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 II and features heterogeneous exposure to global long-run growth shocks. The equity excess return in local units is defined as $r_{d,t}^{ex,i} = \lambda r_{c,t}^{ex,i} + \epsilon_t^i, i \in \{1, 2, \ldots, N\}$, 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. (Color figure can be viewed at wileyonlinelibrary.com)

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 $\beta$ 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 $\beta$s 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-à-vis* the USD) 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 shows that our model's prediction lines up with the data, as countries like Australia and New Zealand (which have the lowest $\beta$s) have negative average NFA positions and countries like Switzerland (which has one of the largest $\beta$s) have positive average NFA positions.

Additionally, the model predicts that countries with extreme $\beta$s should experience a greater fluctuation in their NFA positions, as they trade a large amount of securities to provide and receive insurance against their exposure to global risk (the reallocation channel). In contrast, countries with $\beta$s 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.

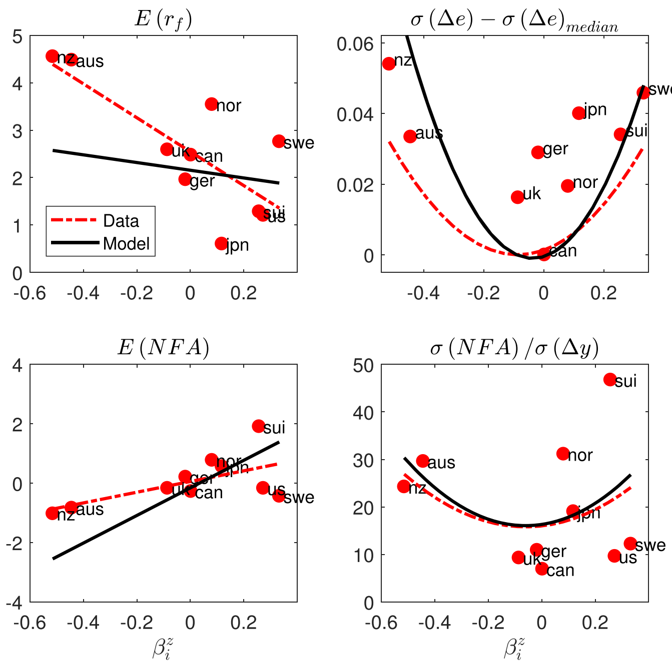


**Figure 6. Cross-sectional patterns.** In each panel, the horizontal axis corresponds to the estimated exposure coefficient $\beta_i^z$. For each country in our data set, we use the estimates provided in Table I. The top-left panel links exposure to global long-run risk to the average risk-free rate in the cross section of countries. The top-right panel corresponds to the volatility of the growth rate of the real exchange rate of each country against the USD. 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 II and features heterogeneous exposure to global long-run growth shocks. (Color figure can be viewed at wileyonlinelibrary.com)

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-$\beta$ countries) should experience an outflow (inflow) of resources at the occurrence of a negative (positive) global long-run shock. Equivalently, the $\lambda^{NFA}$ coefficient should be positive for countries with low $\beta$ and negative for countries with high $\beta$. We test this negative link in our cross section of $\beta$-sorted

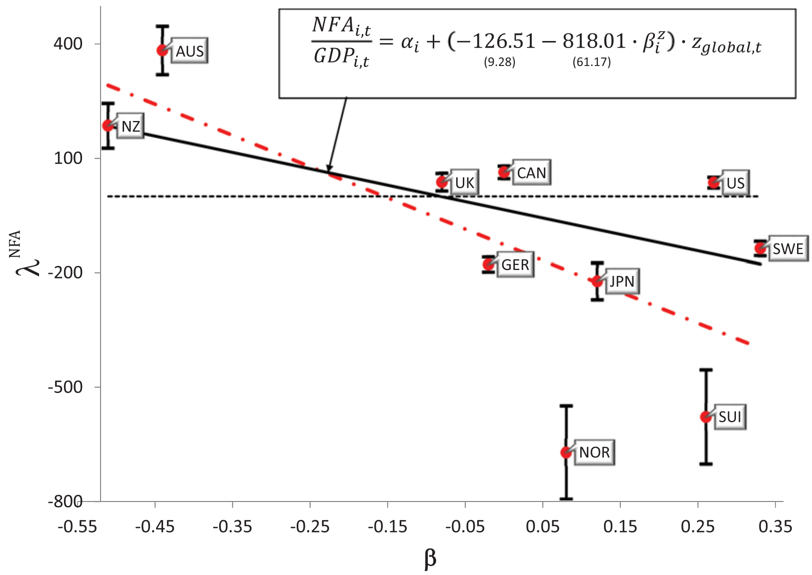


**Figure 7. NFA exposure.** Each dot represents the estimated sensitivity of a country's NFA-to-GDP ratio with respect to global long-run risk plotted (see equation (17), coefficient $\lambda_i^{NFA}$). For each dot, the vertical line represents the 90% confidence interval associated to the estimated 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 $\beta$s are reported in Table I. Standard errors are adjusted for heteroskedasticity. (Color figure can be viewed at wileyonlinelibrary.com)

countries by jointly estimating the following system of equations via GMM:

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

Under the null hypothesis of the model, the estimated $\vartheta_1^{NFA}$ in equation (18) should be negative.

We illustrate our results in Figure 7. Countries with low $\beta$, such as Australia and New Zealand, have positive estimated $\lambda^{NFA}$ coefficients in equation (17), whereas countries with high $\beta$, such as Switzerland and Sweden, have negative estimated $\lambda^{NFA}$ coefficients. For the countries with the most extreme exposure to global shocks, we can typically reject the null hypothesis that their estimated $\lambda^{NFA}$ coefficient is equal to 0. Furthermore, the negative cross-sectional link between $\beta$ and $\lambda^{NFA}$ 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 $\lambda_i^{FX}$ in

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

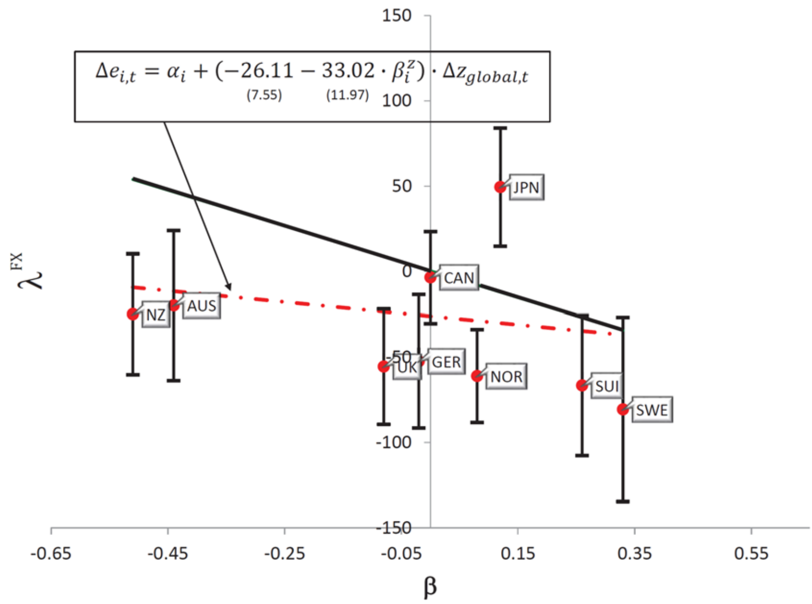


**Figure 8. Currency exposure.** Each dot represents the estimated sensitivity of the growth rate of the exchange rate of a country's currency relative to the USD with respect to a global long-run shock (see equation (19), coefficient $\lambda_i^{FX}$). For each dot, the vertical line represents the 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 $\beta$s are reported in Table I. Standard errors are adjusted for heterokedasticity. (Color figure can be viewed at wileyonlinelibrary.com)

We then estimate the slope coefficient $\vartheta_1^{FX}$ in the following system of equations:

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

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 $\beta$) experience stronger appreciation in response to negative global growth shocks. This is equivalent to a negative $\vartheta_1^{FX}$. In our cross section, the estimate of this coefficient is 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.[10]

[10] 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 (2011) 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. In the Internet Appendix, we document that, for the more recent 2007 to 2008 Great Recession, both the exchange rate and the NFA behaved exactly as predicted by the model (see Figure IA.3, Panels A and B, in the Internet Appendix.)

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 IV we focus on portfolios constructed according to countries' $\beta$s, as opposed to focusing on each country separately.

Specifically, we construct three portfolios based on the results of our analysis in Table I. The first portfolio (labeled "Low $\beta_i^z$") pools the estimates of $\lambda_i^{FX}$ and $\lambda_i^{NFA}$ for all countries with $\beta$s that are negative and statistically different from zero (Australia and New Zealand). The second portfolio (labeled "Medium $\beta_i^z$") contains countries whose $\beta$s are not statistically different from zero (the United Kingdom, Germany, Canada, Norway, and Japan). The third portfolio (labeled "High $\beta_i^z$") consists of countries with $\beta$s that are positive and statistically different from zero (the United States, Switzerland, and Sweden).

Columns (1) and (5) of Table IV are the portfolio counterparts of the country-level estimated coefficients depicted in Figure 7. Columns (2), (6), and (7) show that our results are robust to the exclusion of Japan from portfolio 2 and the 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 United States, since the United States 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 supportive of 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, Roussanov, and Verdelhan (2011) and Lustig, Roussanov, and Verdelhan (2014) to macroeconomic fundamentals such as international consumption dynamics.

### *D. 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 so using the time series of long- and short-run shocks estimated in Section I.

To carry out this empirical exercise, we first need a functional form for each country's SDF. In a single-country setup, this function is generally obtained via a log-linear approximation (e.g., Bansal, Kiku, and Yaron (2008)). Because our model does not have a closed-form approximation, we proceed numerically using a second-order polynomial projection approach. Specifically, we first simulate the model over a grid of values of $\psi$ and $\gamma$. We then use a projection to

**Table IV**

## Conditional Analysis

Unless otherwise specified in the last four rows of the table, the "Low $\beta_i^z$" portfolio contains Australia and New Zealand; the "Medium $\beta_i^z$" portfolio contains the United Kingdom, Germany, Canada, Norway, and Japan; and the "High $\beta_i^z$" portfolio contains Switzerland, the United States, and Sweden. The United States is excluded from the analysis in columns (1) to (4) because the USD is the base currency for this part of the analysis. The coefficients reported in columns (1) and (2) are obtained by pooling the coefficients $\lambda_i^{FX}$ in equation (19) for the countries of the corresponding portfolios. The coefficients in columns (3) and (4) repeat the same analysis after augmenting equation (19) with the component of the domestic long-run shock that is orthogonal to the global long-run shock, $\Delta(z_{i,t} - \beta_i^z \cdot z_{global,t})$, and the component of the U.S. long-run shock that is orthogonal to the global long-run shock, $\Delta(z_{\mathrm{US},t} - \beta_{\mathrm{US}}^z \cdot z_{global,t})$. The coefficients reported in columns (5) to (7) are obtained by pooling the coefficients $\lambda_i^{NFA}$ in equation (17) for the countries of the corresponding portfolios. The coefficients in columns (8) to (10) repeat the same analysis, after augmenting equation (17) with the nonglobal component of the local long-run risk, $z_{i,t} - \beta_i^z \cdot z_{global,t}$. In all specifications, the sample is 1987 to 2013. All standard errors are adjusted for heteroskedasticity. The average exposures to the global long-run risk ($\beta_i^z$) are reported in Table IA.X of the Internet Appendix.

| | Exchange Rate | | | | Net Foreign Assets | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|
| Portfolio | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) | (10) |
| Low $\beta_i^z$ | −26.52 | −26.52 | −27.51 | −27.51 | 522.53 | 522.53 | 522.53 | 544.72 | 544.72 | 544.72 |
| | (16.05) | (16.05) | (13.91) | (13.91) | (87.20) | (87.20) | (87.20) | (186.93) | (186.93) | (186.93) |
| Medium $\beta_i^z$ | −23.85 | −38.49 | −27.31 | −52.15 | 30.61 | 30.61 | −7.50 | 10.41 | 10.41 | −8.02 |
| | (6.08) | (9.35) | (12.06) | (12.72) | (6.29) | (6.29) | (4.32) | (6.51) | (6.51) | (4.52) |
| High $\beta_i^z$ | −67.16 | −67.16 | −84.88 | −84.88 | −155.05 | −145.05 | −155.05 | −140.73 | −140.90 | −140.73 |
| | (23.75) | (23.75) | (20.43) | (20.43) | (17.21) | (17.34) | (17.21) | (11.67) | (16.05) | (11.67) |
| Include All | – | – | – | – | Y | N | N | Y | N | N |
| Exclude JPN | N | Y | N | Y | N | N | Y | N | N | Y |
| Exclude USA | Y | Y | Y | Y | N | Y | N | N | Y | N |
| Control local shocks | N | N | Y | Y | N | N | N | Y | Y | Y |

obtain an approximate functional form of each country's SDF in terms of $\psi$, $\gamma$, 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, the preference parameters, and the model's shocks and observables. This regression yields a specification for each country's SDF:

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

This approach provides a very accurate approximation of the equilibrium SDFs with $R^2$s on the order of 99%. We provide more details on this procedure in Section VI of the Internet Appendix.

We feed the time series of the shocks estimated in Section I into the SDFs reported in (21). Using GMM, we then estimate the EIS ($\psi$) and the risk-aversion coefficient ($\gamma$) using the following moment restrictions:

(1) The second moment of each country's exchange rate growth rate (labeled "FX vol" in Table V),

$$\frac{1}{T}\sum_{t=1}^{T}\left(m^i_t - m^{US}_t\right)^2 - \left(\Delta e^i_t\right)^2 = 0. \tag{22}$$

(2) Each country's NFA position (labeled "NFA" in Table V),

$$\frac{1}{T}\sum_{t=1}^{T}\exp\left\{m^i_{t+1}\right\}A^i_{t+1} - NX^i_t - A^i_t = 0. \tag{23}$$

(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 "Market returns" and "Risk-free rates," respectively, in Table V),

$$\frac{1}{T}\sum_{t=1}^{T}\exp\{m^i_t\}R^i_{k,t} - 1 = 0, \quad \forall k = s, f. \tag{24}$$

(4) The Euler equations for the pricing of the currency excess returns of the six HML and IMX portfolios ("HML (6 portf)" and "IMX (6 portf)" in Table V), the excess return of the sixth over the first HML and IMX portfolios ("HML (6-1)" and "IMX (6-1)" in Table V), and the excess return of a strategy that is short in the United States and long in each of the remaining nine currencies in the G10 countries ("RFX (G-10)" in Table V).

**Table V**

**Testing Moment Restrictions**

This table reports the estimated preference parameters $\psi$ and $\gamma$ associated with nine 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 $p$-values displayed in brackets underneath. The $J$-statistics reported in the second panel of the table refer to the full set of moment restrictions ( "$J$-stat (Full)") and to the subset of moment restrictions that is limited to asset returns ( "$J$-stat (Pricing)"). The numbers in brackets underneath each $J$-stat denote the corresponding $p$-value. The sample is 1987 to 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 (equation (22) in the main text, labeled "FX Vol."); the average intertemporal condition of each G-10 country's NFA position (equation (23) in the main text, labeled "NFA"); the Euler equation for the currency excess returns of each G-10 currency relative to the United States (labeled "RFX (G10)"), for the currency excess returns of the six 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).

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) |
|---|---|---|---|---|---|---|---|---|---|
| $\psi$ | 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) |
| $\gamma$ | 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 | | | | | | | | √ | √ |

As a result, we have a total of up to 62 moment restrictions to estimate the $\psi$ and $\gamma$ parameters in a cross section of 10 countries.

### *D.1. Discussion*

Table V 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 they are imprecisely estimated (column (1)). A test of the null hypothesis that investors have time-additive CRRA preferences ($\gamma = 1/\psi$)

cannot be rejected at conventional confidence levels. The results improve when we add the cross section of NFA positions to our estimation (columns (2) and (3)): the preference parameters are now sharply identified, and we can 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 $\psi$ and $\gamma$ of approximately 1 and 4, respectively (columns (4), (5), and (6)). The results in column (7) show 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 0 (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 V 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 those 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.

## IV. 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 heterogeneous 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 a number of important economic implications that can be empirically assessed. We find good, but not perfect, alignment between our model and the data.

Future developments should extend this setting to international real business cycle models to study the role of international investment flows and international frictions for the cross section of currency risk premia. Investigation of the roles of trading frictions, portfolio composition, and market incompleteness offers other promising directions for future research. In these settings, it is important to explore the accuracy of both local and global approximations, in the spirit of Rabitsch, Stepanchuk, and Tsyrennikov (2015).

While we can match the cross-sectional coefficient of variation for many variables of interest, we cannot fully replicate their cross-country spread. Future research needs to develop models that better fit the dynamics of asset prices and quantities in our cross section of countries. Furthermore, future models will need to introduce more realistic trading impediments, so that the tensions between strong consumption home bias and currency risk premia can be alleviated.

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## Supporting Information

Additional Supporting Information may be found in the online version of this article at the publisher's website:

**Appendix S1**: Internet Appendix.
Replication Code