

Supplemental Material to “Commodity Prices and Sovereign Risk”

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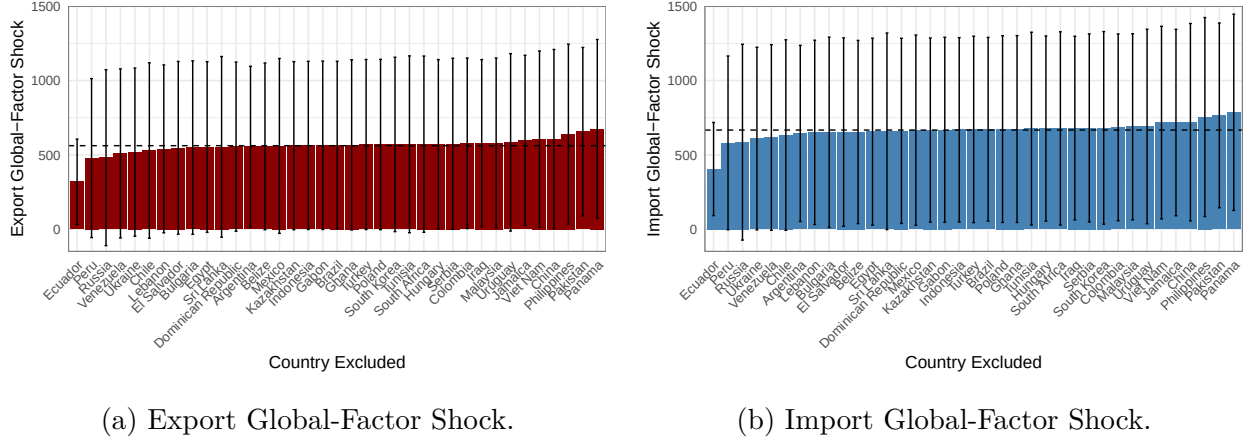
This Supplemental Material complements the appendices of the paper. Section [S1](#) reports robustness exercises for the empirical analysis of Section 2. Section [S2](#) discusses the measurement of net endowments and an alternative CES specification of preferences. Section [S3](#) reports the commodity crosswalk used to build the exposure parameters, and Section [S4](#) describes the estimation and solution algorithm in full. Section [S5](#) reports the default sets and bond-price schedules for all 37 countries and compares simulated debt with the data. Equation, section, table, and figure numbers without the prefix S refer to the paper.

S1 Empirical Robustness

S1.1 Leave-One-Country-Out Robustness

To assess whether the global-factor results are driven by individual countries, we re-estimate Column (3) of Table 1, excluding one country at a time. Figure [S1](#) shows that both export and import shock coefficients remain positive across all 37 exclusions, indicating that the baseline pattern is not driven by any single country.

Figure S1: Leave-One-Country-Out Estimates of Export and Import Global-Factor Shocks.



Note: The figure reports leave-one-country-out estimates of the export and import global-factor shock coefficients from Column (3) of Table 1. Each bar corresponds to a separate regression in which the country indicated on the horizontal axis is excluded from the estimation sample. Vertical lines report 95% confidence intervals based on standard errors clustered at the country level. The dashed horizontal line reports the corresponding full-sample estimate. All specifications include country and year fixed effects, real GDP growth, CPI inflation, nominal exchange-rate growth, and the underlying export and import global-factor exposure measures.

S1.2 Alternative Controls for Common Global Conditions

Our baseline specification includes year fixed effects, which absorb the aggregate response to the common commodity-price innovation together with other common annual developments. We assess the sensitivity of the results by replacing year fixed effects with explicit controls for these common factors, focusing on the net-gross specification in Table 2. Column (1) of Table S1 reproduces the baseline specification. Column (2) replaces year fixed effects with the standardized global commodity-price innovation, thereby controlling directly for the aggregate response to the common shock. Motivated by the emphasis in [Sosa-Padilla & Sturzenegger \(2023\)](#) on global risk conditions and U.S. interest rates, Column (3) further includes the annual log change in the average VIX and the annual change in the average U.S. ten-year Treasury yield. All specifications retain country fixed effects, country-level controls, the same estimation sample, and the same shock normalization.

The alternative specifications closely reproduce the baseline net-gross pattern. The net-shock coefficient is -52.6 basis points in the baseline, -54.4 when year fixed effects are replaced by the global commodity-price innovation, and -49.5 after additionally controlling for global financial conditions. The corresponding gross-shock coefficients are 614.6 , 595.0 ,

Table S1: Net and Gross Global-Factor Shocks: Alternative Controls for Common Global Conditions.

	Δ Spread		
	(1)	(2)	(3)
Net Global-Factor Shock	-52.6*** (18.4)	-54.4** (20.4)	-49.5** (19.0)
Gross Global-Factor Shock	614.6** (301.9)	595.0** (288.6)	536.7* (273.6)
Global Commodity-Price Innovation		-979.8** (482.6)	-893.5* (458.0)
Δ U.S. 10-Year Treasury Yield			-46.1* (23.6)
$\Delta \log$ VIX			214.7*** (47.4)
Observations	572	572	572
R ²	0.31	0.19	0.22
Country Fixed Effects	✓	✓	✓
Year Fixed Effects	✓		

Note: Column (1) reproduces the baseline specification with country and year fixed effects. Column (2) replaces year fixed effects with the standardized innovation to the global commodity-price factor. Column (3) further controls for the annual log change in the average CBOE Volatility Index (VIX) and the annual change in the average U.S. 10-year Treasury yield. All specifications include country fixed effects, controls for real GDP growth, CPI inflation, and nominal exchange-rate growth, and the underlying export and import global-factor exposure measures. The estimation sample and shock normalization are held fixed across columns. VIX and U.S. Treasury-yield data are from FRED.

and 536.7 basis points. Thus, both the signs and magnitudes of the net and gross responses remain close to the baseline when common annual developments are captured through explicit global controls rather than year fixed effects.

We retain year fixed effects in the baseline because they flexibly absorb common annual developments, leaving identification to differential responses across countries with different predetermined trade exposures. The alternative specifications show that the main results are robust to replacing this flexible treatment of common time variation with explicit controls for the common commodity-price innovation and global financial conditions.

S1.3 Sensitivity to Exchange-Rate Arrangements

We examine whether the baseline relationships persist under restricted exchange-rate flexibility using the de facto exchange-rate regime classification of [Ilzetzki et al. \(2019, 2022\)](#). We assign each country-year observation the regime recorded in December of the preceding year. This measures the arrangement before the current year’s commodity-price innovation, so that regime changes during the shock year do not determine sample assignment.

Table S2 reports the full-sample estimates alongside estimates for pegs and narrow horizontal bands (fine codes 1–4) and more flexible arrangements (fine codes 9–13). Crawling arrangements and narrow bands (codes 5–8), freely falling episodes (code 14), and observations with missing parallel-market information (code 15) or missing regime classifications

Table S2: Commodity-Price Exposure and Sovereign Spreads across Exchange-Rate Arrangements.

	Δ Spread		
	Full sample	Pegs / Horizontal bands	More flexible
	(1)	(2)	(3)
Net Global-Factor Shock	-52.6*** (18.4)	-63.4** (27.1)	-23.2*** (7.10)
Gross Global-Factor Shock	614.6** (301.9)	304.0 (403.9)	241.7*** (68.6)
Observations	572	152	215
R ²	0.31	0.60	0.71
Country Fixed Effects	✓	✓	✓
Year Fixed Effects	✓	✓	✓

Note: The dependent variable is the annual change in the sovereign spread. Column (1) reproduces the baseline specification. Columns (2) and (3) restrict the sample using the de facto exchange-rate classification of [Ilzetzki et al. \(2019, 2022\)](#). Regimes are measured in December of year $t - 1$. Pegs and narrow horizontal bands comprise fine codes 1–4; more flexible arrangements comprise codes 9–13. Crawling arrangements and narrow bands (codes 5–8), freely falling episodes (code 14), and observations with missing parallel-market information (code 15) or missing regime classifications remain in column (1) but are excluded from columns (2) and (3). All specifications include real GDP growth, inflation, exchange-rate growth, export and import exposure controls, and country and year fixed effects. Shock normalization is unchanged across columns. Standard errors clustered by country are reported in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

remain in the full sample but are excluded from both restricted samples. All specifications retain the baseline controls, country and year fixed effects, country-clustered standard errors, and the original shock normalization.

The net-exposure coefficient remains negative and statistically significant in both restricted samples, with estimates of -63.4 under pegs and narrow horizontal bands and -23.2 under more flexible arrangements, compared with -52.6 in the full sample. The gross-exposure coefficient is positive in both subsamples, but is imprecisely estimated under pegged arrangements.

S2 Preferences and Measurement

S2.1 Measurement of Net Endowments

The aggregate environment depends on endowments and subsistence requirements through their difference, $\tilde{e}(j) = e(j) - \bar{c}(j)$. We approximate the corresponding value using observed commodity-level net trade. This mapping is not exact: the demand system implies

$$P_t(j)[e(j) - c_t(j)] = P_t(j)\tilde{e}(j) - \lambda(j) \left(E_t - \int_0^1 P_t(j')\bar{c}(j') dj' \right).$$

Observed net exports thus also reflect discretionary consumption above subsistence. Separately identifying these components would require detailed household expenditure data unavailable at the country–commodity level of our analysis.¹ The approximation does not rely on the consumption weights $\lambda(j)$ being small, since consumption baskets are concentrated (the largest weight is 0.26 for the median country); it requires discretionary consumption of the commodities to be small relative to net trade, as for necessities. Summing over goods, the omitted term is spending on commodities above subsistence, R_i . It matters when commodity exports and imports nearly offset, which is why Section 4.1 bounds the scale of trade income from below by R_i . Because the classification depends only on the sign of $\tilde{\nu}_i$, the relevant margin is $|\tilde{\nu}_i|$: one country, South Africa, lies within 0.02 of zero, and one more, Uruguay, lies within 0.05.

¹Estimates of committed consumption date back to [Stone \(1954\)](#) and have been applied to developing economies ([Lluch et al. 1977](#), [Atkeson & Ogaki 1996](#), [Ogaki et al. 1996](#)).

S2.2 Alternative Preferences: CES and the Role of Non-Homotheticity

Consider a homothetic CES aggregator with elasticity of substitution $\sigma_c > 0$. Its associated consumption price index is:

$$P_t^C = \left[\int_0^1 \lambda(j) P_t(j)^{1-\sigma_c} dj \right]^{1/(1-\sigma_c)}.$$

The expenditure share of good j is:

$$s_t(j) = \frac{\lambda(j) P_t(j)^{1-\sigma_c}}{\int_0^1 \lambda(j') P_t(j')^{1-\sigma_c} dj'}.$$

Differentiating the price index with respect to factor f_k gives:

$$\frac{\partial \ln P_t^C}{\partial f_k} = \int_0^1 s_t(j) a_k(j) dj.$$

Unlike under Cobb–Douglas preferences, these expenditure shares vary with relative prices:

$$\frac{\partial \ln s_t(j)}{\partial \ln P_t(j)} = (1 - \sigma_c)(1 - s_t(j)).$$

When $\sigma_c < 1$, an increase in a good's relative price raises its expenditure share. Consequently, a higher global factor places greater weight on commodities with larger factor loadings, strengthening the price-index response.

However, changing the elasticity of substitution does not introduce subsistence requirements. In the endowment environment considered here, homothetic CES therefore retains $\tilde{e}(j) = e(j) \geq 0$ and, when factor loadings are non-negative, cannot generate negative terms-of-trade exposure. Consumption can still decline if the price-index effect dominates the income effect, but the two channels cannot reinforce one another through negative net-endowment exposure.

Stone–Geary preferences allow such negative exposure while keeping the price index a geometric average with fixed weights $\lambda(j)$ and a tractable decomposition of the two channels. Under CES, by contrast, price-index exposure is itself state dependent because expenditure shares vary with relative prices.

S3 Commodity Crosswalk

This section reports the commodity crosswalk used to align price, trade, and production data when constructing the country-specific exposure parameters in the quantitative analysis (Section 4.1).

Table S3: Commodity Crosswalk – SITC, FAO, and Production Sources.

Commodity	SITC Rev. 1	FAO Item Code	Production Source
Cotton	263, 652, 6513, 6514	767, 328	FAO
Hard logs	2423	1863, 1867, 1657, 1670	FAO
Soft logs	2422	1862, 1866, 1651	FAO
Hard sawnwood	2433	1633, 1644	FAO
Soft sawnwood	2432	1632, 1643	FAO
Hides	2111, 2119	957, 919, 1025	FAO
Natural rubber	2311	836, 839	FAO
Wool	262, 6512	987	FAO
Coal	32	-	U.S. EIA
Natural gas	3411	-	U.S. EIA
Crude oil	331	-	U.S. EIA
Bananas	0513	486, 489	FAO
Barley	043	44	FAO
Beef	0111	947, 867	FAO
Chicken	0114	1058, 1080	FAO
Cocoa	072	661	FAO
Coffee	071	656	FAO
Corn	044	56	FAO
Fish	0311, 0312, 03201	1176	FAO
Groundnuts	2211	242	FAO
Lamb	0112	977, 1017	FAO
Olive oil	4215	261	FAO
Oranges	05111	490	FAO

(continued)

Commodity	SITC Rev. 1	FAO Item Code	Production Source
Palm oil	4222	257	FAO
Pork	0113	1035	FAO
Rapeseed oil	4217	271	FAO
Rice	042	27	FAO
Soybean oil	4212	237	FAO
Soybeans	2214	236	FAO
Sugar	0611, 0612	162	FAO
Sunflower seed oil	4216	281, 268	FAO
Tea	0741	667	FAO
Wheat	0410	15	FAO
Aluminum	684	-	USGS
Copper	2831, 682	-	USGS
Iron ores	2813	-	USGS
Lead	2834, 685	-	USGS
Nickel	28321, 683	-	USGS
Tin	2836, 687	-	USGS
Uranium	286, 688	-	U.N. Data
Zinc	2835, 686	-	USGS

Note: SITC Rev. 1 codes are used to extract trade flows from UN Comtrade. FAO item codes correspond to FAOSTAT crop and livestock production data. Production sources labeled USGS refer to the U.S. Geological Survey Minerals Yearbook; U.S. EIA refers to the U.S. Energy Information Administration. A dash (-) indicates that FAO or an alternative source is not the primary production provider for that category.

S4 Estimation and Solution Algorithm

This section describes the empirical construction of the exposure parameters $(\psi, \nu, \bar{Z}, \mu^C)$ and the numerical procedure used to compute the equilibrium. The structure mirrors the

calibration framework in Section 4.1. We first estimate the global factor structure and the exposure parameters from the data and then, conditional on these inputs, solve the government's optimization problem and the associated market-clearing bond price schedule.

Link to PCA Facts. The factorization:

$$\ln P_t(j) = \mu(j) + \mathbf{a}(j)^\top \mathbf{f}_t + \varepsilon_t(j)$$

mirrors the empirical evidence presented in Section 2. The factor dynamics are estimated on annual log prices of the 41 commodities over 1991–2019, each linearly detrended. On this panel, the first principal component explains 60.8% of the total variance and the first five components 87.1%, compared with 66.6% and 86.8%, respectively, in Section 2. These differences reflect the sample and preprocessing choices: Section 2 uses standardized log price levels over 1990–2023 without detrending.

The persistence of global shocks is captured by the process:

$$\mathbf{f}_t = \Phi \mathbf{f}_{t-1} + \mathbf{u}_t, \quad \mathbf{u}_t \sim \mathcal{N}(\mathbf{0}, \Sigma_u),$$

with eigenvalues of Φ strictly inside the unit circle.

The idiosyncratic components $\varepsilon_t(j)$ exhibit persistence, with a median first-order autocorrelation of 0.52 in the annual panel. The model abstracts from these components, as discussed in Section 3.1.

Two Exposure Vectors. Aggregate sensitivity to global shocks is governed by two exposure vectors:

$$(i) \text{ Price-index channel: } -\boldsymbol{\psi}^\top \mathbf{f}_t, \quad (ii) \text{ Terms-of-trade channel: } +\boldsymbol{\nu}^\top \mathbf{f}_t.$$

Export concentration in goods with high factor loadings $a_k(j)$ increases ν_k , while concentration of consumption weights $\lambda(j)$ on goods with high $a_k(j)$ increases ψ_k .

Calibration and Estimation Steps.

1. **Principal Component Estimation.** Let N denote the number of commodities (in our case, $N = 41$). Construct the $T \times N$ matrix of detrended log prices:

$$X = \left[\ln P_{t,j} - \widehat{\ln P}_{t,j} \right]_{t=1, \dots, T; j=1, \dots, N},$$

where $\widehat{\ln P_{t,j}}$ is a commodity-specific linear trend. We perform PCA on X (covariance PCA) to obtain the factor representation:

$$X \approx FA^\top, \quad F \in \mathbb{R}^{T \times K}, \quad A \in \mathbb{R}^{N \times K},$$

where the j th row of A , denoted $\mathbf{a}(j)^\top$, contains the loadings for commodity j , so that

$$\ln P_{t,j} \approx \widehat{\ln P_{t,j}} + \mathbf{a}(j)^\top \mathbf{f}_t.$$

Here $\mathbf{f}_t \in \mathbb{R}^K$ is the vector of factor realizations at time t . The factors are normalized to have zero mean and identity covariance, $\mathbb{E}[\mathbf{f}_t] = \mathbf{0}$ and $\mathbb{E}[\mathbf{f}_t \mathbf{f}_t^\top] = I_K$. This decomposition is used only to estimate factor dynamics. Country exposures are constructed using the loadings $a(j)$ estimated in Section 2 and rescaled to log-price units as described in Section 4.1.

2. **Factor Dynamics.** We estimate the autoregressive process:

$$\mathbf{f}_t = \Phi \mathbf{f}_{t-1} + \mathbf{u}_t, \quad \mathbf{u}_t \sim \mathcal{N}(\mathbf{0}, \Sigma_u),$$

by OLS, convert the estimated dynamics to the model's quarterly frequency, and discretize the resulting process into a finite Markov chain $\{f^m, P_f(m, m')\}$ using the method of [Tauchen \(1986\)](#).

3. **Inputs from Consumption and Trade Data.** We construct: (i) consumption weights $\lambda = (\lambda_1, \dots, \lambda_N)$ with $\sum_j \lambda_j = 1$; (ii) net endowments $\tilde{e} = (\tilde{e}_1, \dots, \tilde{e}_N)$, where $\tilde{e}(j) = e(j) - \bar{c}(j)$; and (iii) reference prices $\bar{P}_j$, the geometric mean of annual prices over 1991–2019 (Section 4.1). These objects map the observed consumption and trade structure into model inputs.

4. **Computation of Exposure Parameters.** We compute:

$$\begin{aligned} \boldsymbol{\psi} &= A^\top \lambda, & \boldsymbol{\nu} &= A^\top (\bar{P} \odot \tilde{e}), \\ \bar{Z} &= \sum_{j=1}^N \bar{P}_j \tilde{e}_j, & \mu^C &= \sum_{j=1}^N \lambda_j \mu(j), \end{aligned}$$

where $\odot$ denotes elementwise multiplication and the rows of A contain the loadings estimated in Section 2, rescaled to log-price units. The vectors $\boldsymbol{\psi}$ and $\boldsymbol{\nu}$ capture exposure through the price-index and terms-of-trade channels, respectively. The terms $\bar{Z}$ and μ^C determine baseline trade income and the baseline log consumption price

index. In the quantitative implementation, μ^C is held common across countries and scales only the level of consumption (Section 4.1).

5. **Country-Specific Calibration.** For each of the 37 countries, we use commodity-level production values V_{ij}^P , exports X_{ij} , and imports M_{ij} to construct the exposure pair $(\psi_i, \tilde{\nu}_i)$, as described in Appendix D. The terms-of-trade exposure, $\tilde{\nu}_i = \frac{\sum_j \bar{P}(j) \tilde{e}_i(j) a(j)}{\max\{|\bar{Z}_i|, R_i\}}$, with R_i spending on commodities above subsistence (equation (25)), weights each net trade position by its loading on the global factor. Its sign therefore captures the alignment of the trade basket with global commodity-price movements: a positive innovation raises trade income when $\tilde{\nu}_i > 0$ and lowers it when $\tilde{\nu}_i < 0$. These signs define exporter- and importer-type economies, respectively (equation (25)). The sign of consumption exposure, $\hat{\Delta}_i$, coincides with that of $\tilde{\nu}_i$ in 35 of the 37 countries (Sections 4.1 and 4.3). The price-index exposure ψ_i instead reflects the composition of domestic consumption.
6. **Solution of the Model.** Consistent with the empirical findings in Section 2 and the quantitative specification in Section 4, we solve the model with a single factor ($K = 1$). Thus, f , ψ , and ν are scalars throughout the remainder of this section. Given $(\psi, \nu, \bar{Z}, \mu^C)$ and the calibrated parameters, we solve for equilibrium recursively:

- (a) Initialize a bond price schedule $q^{(0)}(f, b') = 1/(1+r^*)$ and value functions $V^{r,(0)}(f, b)$ and $V^{d,(0)}(f)$.
- (b) For each state (f, b) , compute the repayment value:

$$V^r(f, b) = \max_{b'} \left\{ u(c^r(f, b, b')) + \beta \sum_{f'} P_f(f, f') V(f', b') \right\},$$

where

$$c^r(f, b, b') = \Gamma^c e^{-\mu^C - \psi f} [b - q(f, b') b' + \bar{Z} + \nu f].$$

The default value is:

$$V^d(f) = u(c^d(f)) - \phi(f) + \beta \sum_{f'} P_f(f, f') [\theta V(f', 0) + (1 - \theta) V^d(f')],$$

with

$$c^d(f) = \chi \Gamma^c e^{-\mu^C - \psi f} (\bar{Z} + \nu f).$$

- (c) Update $V(f, b) = \max\{V^r(f, b), V^d(f)\}$ and define the default region $\mathcal{D}(b) = \{f : V^d(f) > V^r(f, b)\}$.

(d) Compute default probabilities $\delta(f, b') = \sum_{f' \in \mathcal{D}(b')} P_f(f, f')$ and update bond prices:

$$q^{(1)}(f, b') = \frac{1 - \delta(f, b')}{1 + r^*}.$$

(e) Iterate until convergence of $q(f, b')$ and the value functions.

At convergence, equilibrium policies, prices, and default probabilities jointly satisfy optimality and market clearing.

Overall, Steps (1)–(5) discipline the model by constructing exposure parameters from international price, trade, and production data, while Step (6) solves the model numerically via value function iteration. The equilibrium highlights how global price shocks interact with trade exposure to determine consumption, debt, and default risk. Throughout, we set $K = 1$.

S5 Cross-Country Quantitative Results

This section complements Appendix F of the paper with the default sets and bond-price schedules of all 37 countries, and with a comparison of simulated debt levels with the data.

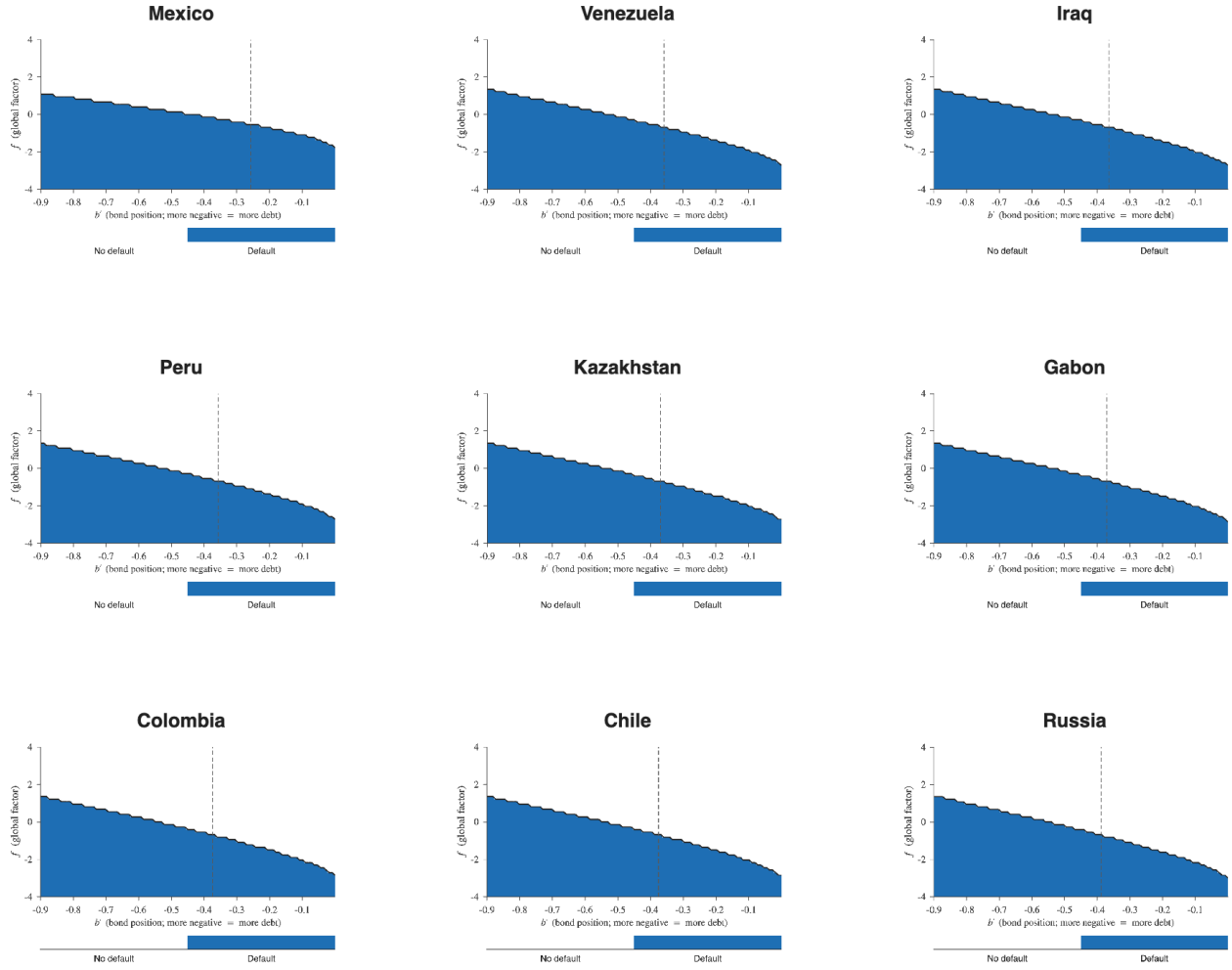
S5.1 Default Sets for the 37 Countries

Figures S2–S5 report the default sets $\mathcal{D}(b)$ for all countries in the sample, grouped by exposure type and ordered within each group by the magnitude of their terms-of-trade exposure $|\tilde{\nu}_i|$. Each heatmap shows the combinations of debt b and global factor f for which default is optimal. The slope of the frontier is governed by the country’s type, while its steepness reflects the magnitude of exposure to the global factor.

The shape of the default frontier reflects the relative strength of the two channels. In exporter-type countries, the frontier slopes downward in f : default occurs at low realizations of the global factor, where adverse terms-of-trade shocks reduce real income and increase default incentives. In importer-type countries, the frontier slopes upward: default occurs at high realizations of the global factor, where increases in the price of the consumption basket compress real income.

The steepness of the frontier increases with exposure to the global factor. Countries with larger absolute exposure display sharper transitions between repayment and default re-

Figure S2: Default Sets – Exporter-type Countries, Part 1.



gions, consistent with a stronger sensitivity of consumption and default incentives to global shocks. This pattern is particularly visible for highly exposed economies such as Mexico and Venezuela among exporters, and Ukraine and Hungary among importers.

Figure S3: Default Sets – Exporter-type Countries, Part 2.

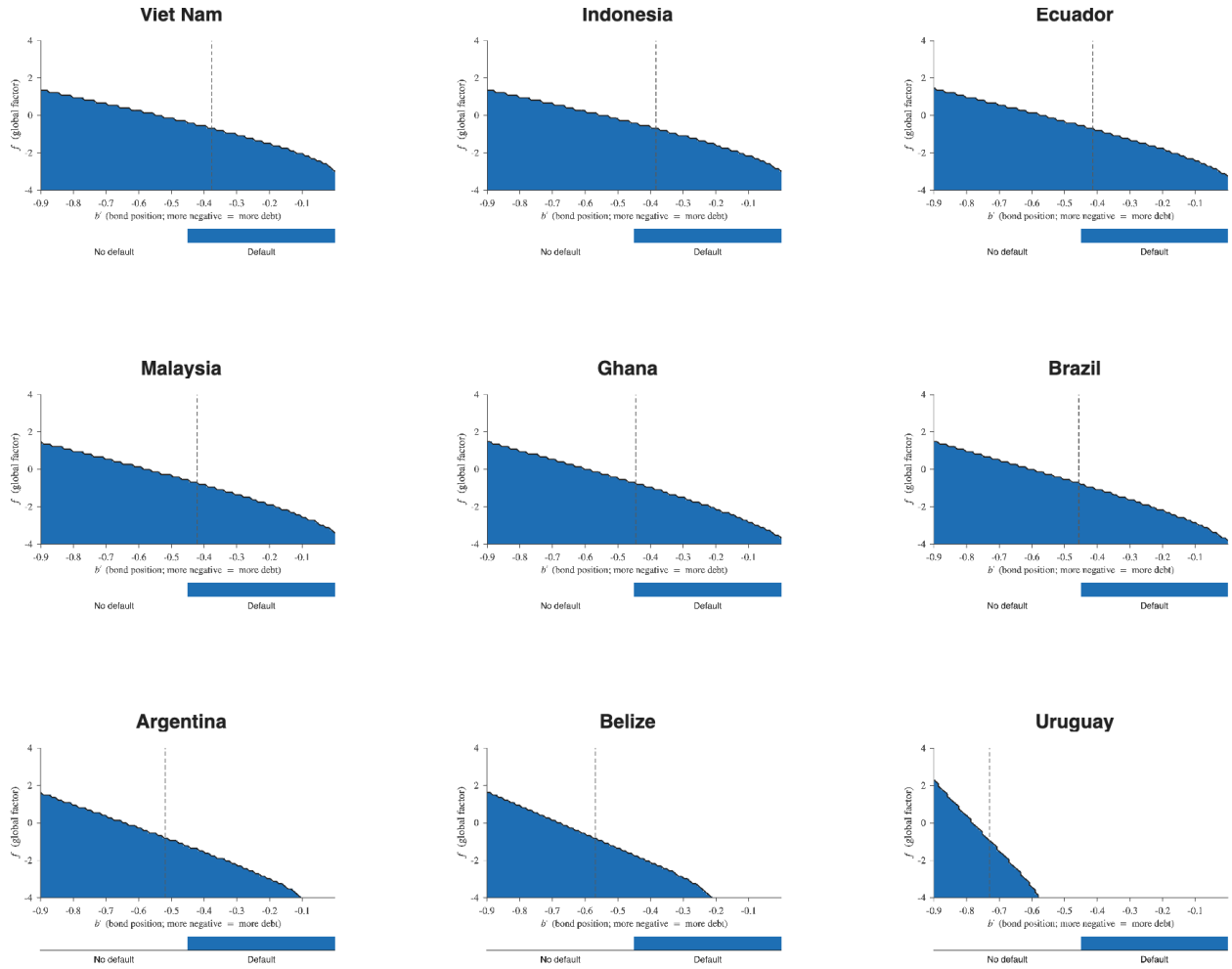


Figure S4: Default Sets – Importer-type Countries, Part 1.

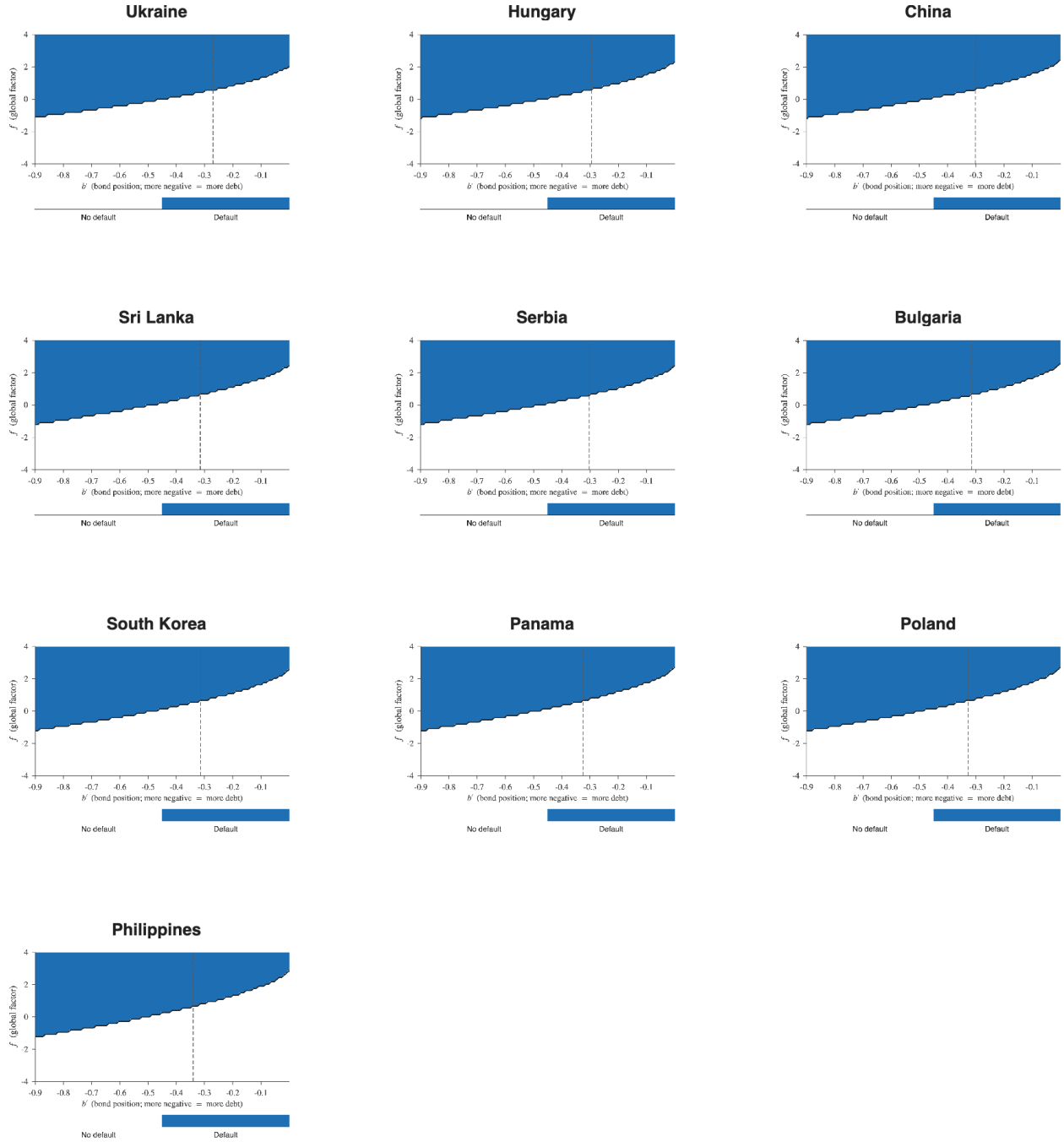
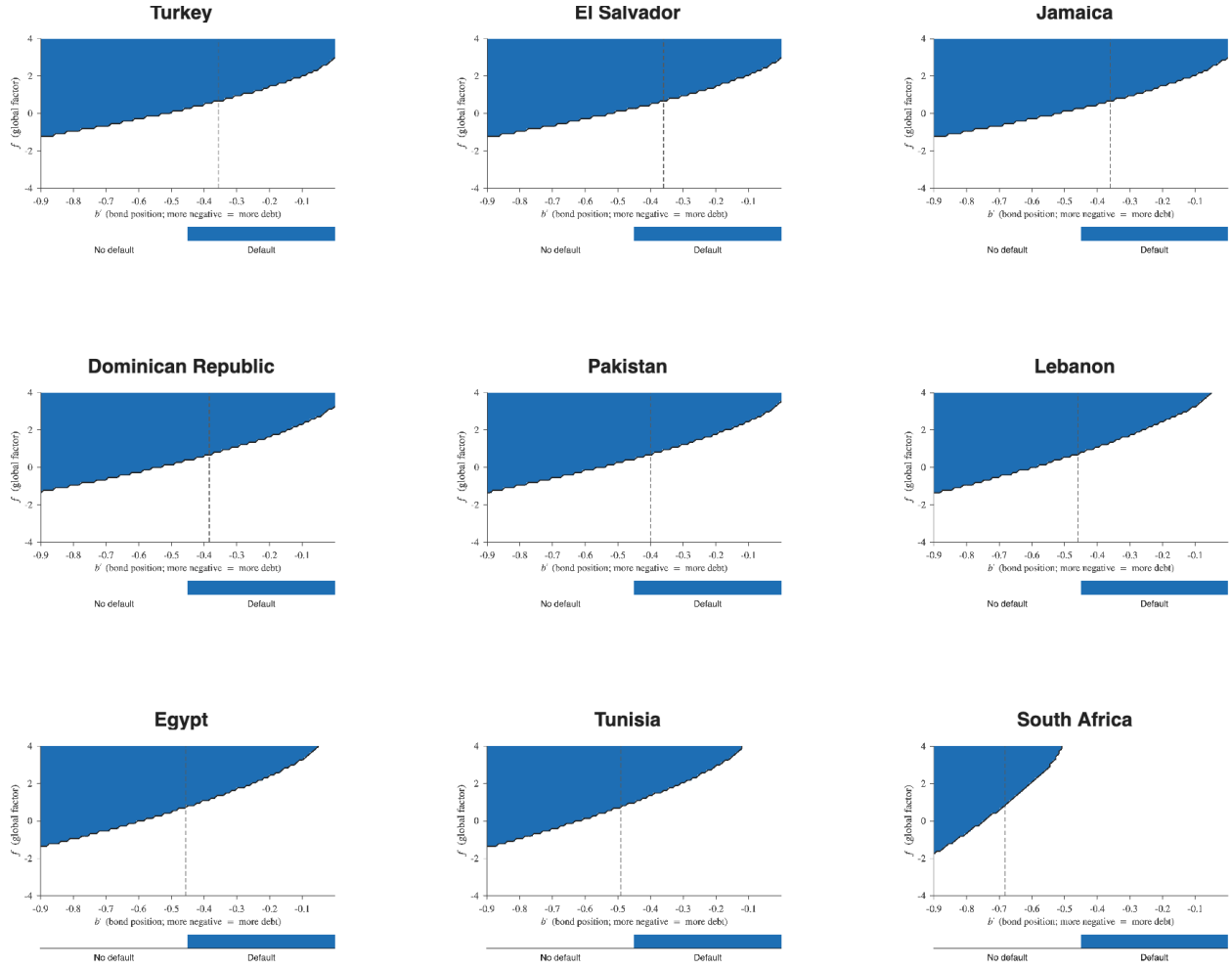


Figure S5: Default Sets – Importer-type Countries, Part 2.



S5.2 Bond Price Schedules for the 37 Countries

Figures S6–S9 report the equilibrium bond price schedules $q(f, b')$ for all countries. Each panel shows, for a given country, the price at which lenders are willing to purchase new debt b' as a function of the current global factor f . Each panel plots bond-price schedules for six realizations of the global factor, chosen for each country as those whose schedule falls from one to zero within the displayed debt range; the color bar reports each value in standard deviations from the mean. Colors range from dark red for low f to light blue for high f and identify factor realizations rather than country types. The dashed vertical line marks the country's mean simulated debt position, and the horizontal axis is restricted to $b' \leq 0$. The ordering of the schedules highlights the contrast across types: in exporter-type economies, the red schedules lie farther to the right, indicating that default risk emerges at lower levels of indebtedness when the factor is low; in importer-type economies, this ordering is reversed.

Bond prices decline as b' becomes more negative, reflecting the higher default risk associated with greater borrowing. The dependence of prices on the global factor mirrors the structure of default incentives. In exporter-type countries, bond prices are increasing in f , so spreads decline when global commodity prices rise. In importer-type countries, bond prices are decreasing in f , implying that spreads increase with the global factor.

The sensitivity of bond prices to f varies across countries and increases with exposure to the global factor. Countries with larger absolute exposure display a wider separation between price schedules at low and high realizations of f , consistent with a stronger response of default risk to global shocks.

The bond price rises with b' in every country, but the range of debt over which it does so varies considerably, from 0.03 to 0.20 across the sample, with a median of 0.14.² That range is governed by exposure. Where consumption responds strongly to the global factor, the default decision depends on the realization of f as well as on the level of debt: at a given b' there is a genuine probability of default in bad states and none in good ones, so the default probability varies smoothly with issuance. Where exposure is small, consumption barely moves with f and default is close to deterministic given b' , so the schedule approaches a step function.³

²The range is defined as the interval of b' over which q rises from 10% to 90% of its maximum value.

³The transition width has a Spearman correlation of 0.76 with absolute terms-of-trade exposure $|\tilde{\nu}_i|$, compared with 0.71 for absolute consumption exposure $|\hat{\Delta}_i|$ and 0.34 for price-index exposure ψ_i . Exposure is also associated with the location of the schedule: the value of b' at which default risk begins to be priced has a Spearman correlation of 0.58 with $|\hat{\Delta}_i|$ and 0.89 with $|\tilde{\nu}_i|$. Since debt corresponds to negative b' , these positive correlations imply that more exposed sovereigns begin paying a risk premium at lower levels of indebtedness and face a more gradual transition in bond prices as debt increases.

Figure S6: Bond Price Schedules – Exporter-type Countries, Part 1.

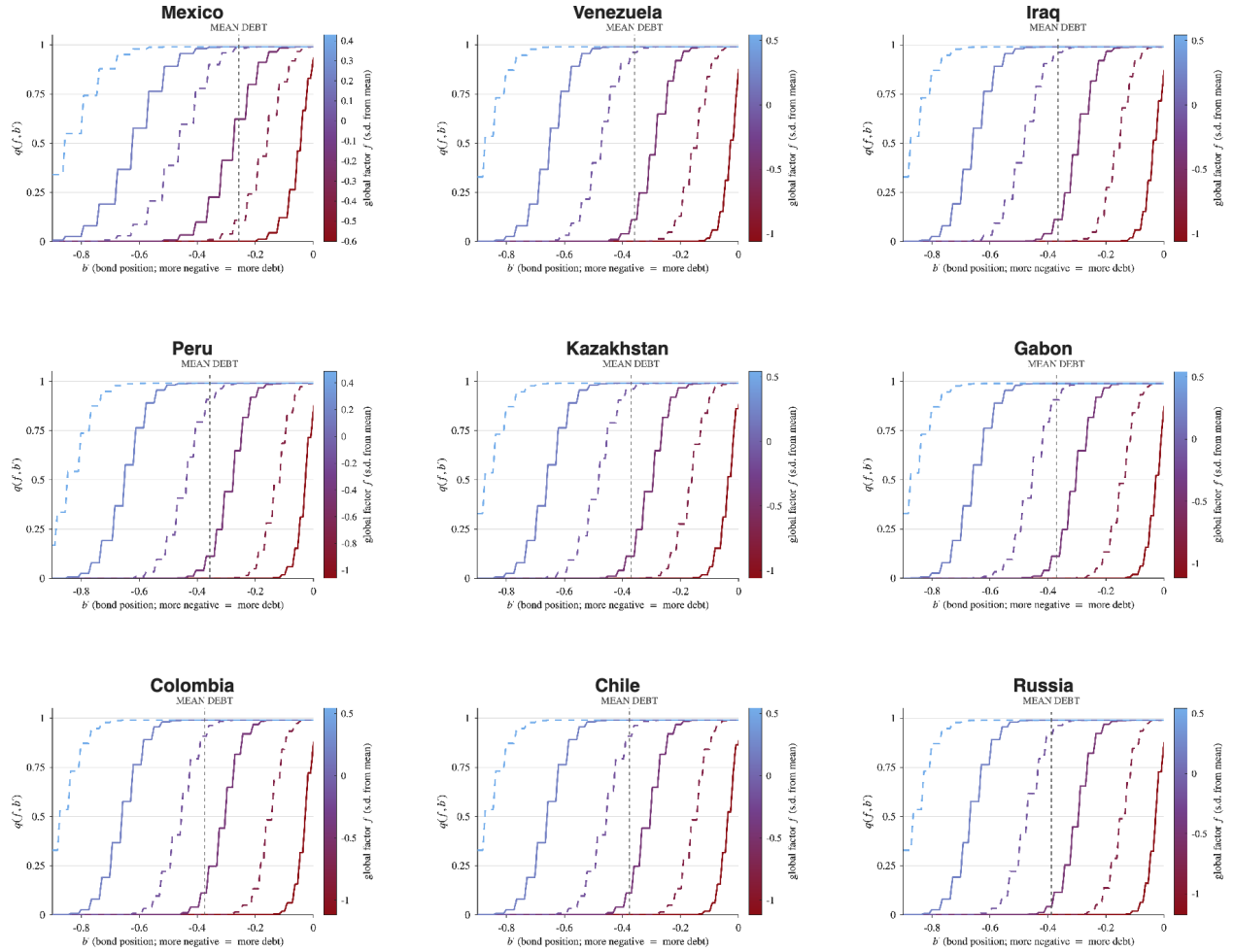


Figure S7: Bond Price Schedules – Exporter-type Countries, Part 2.

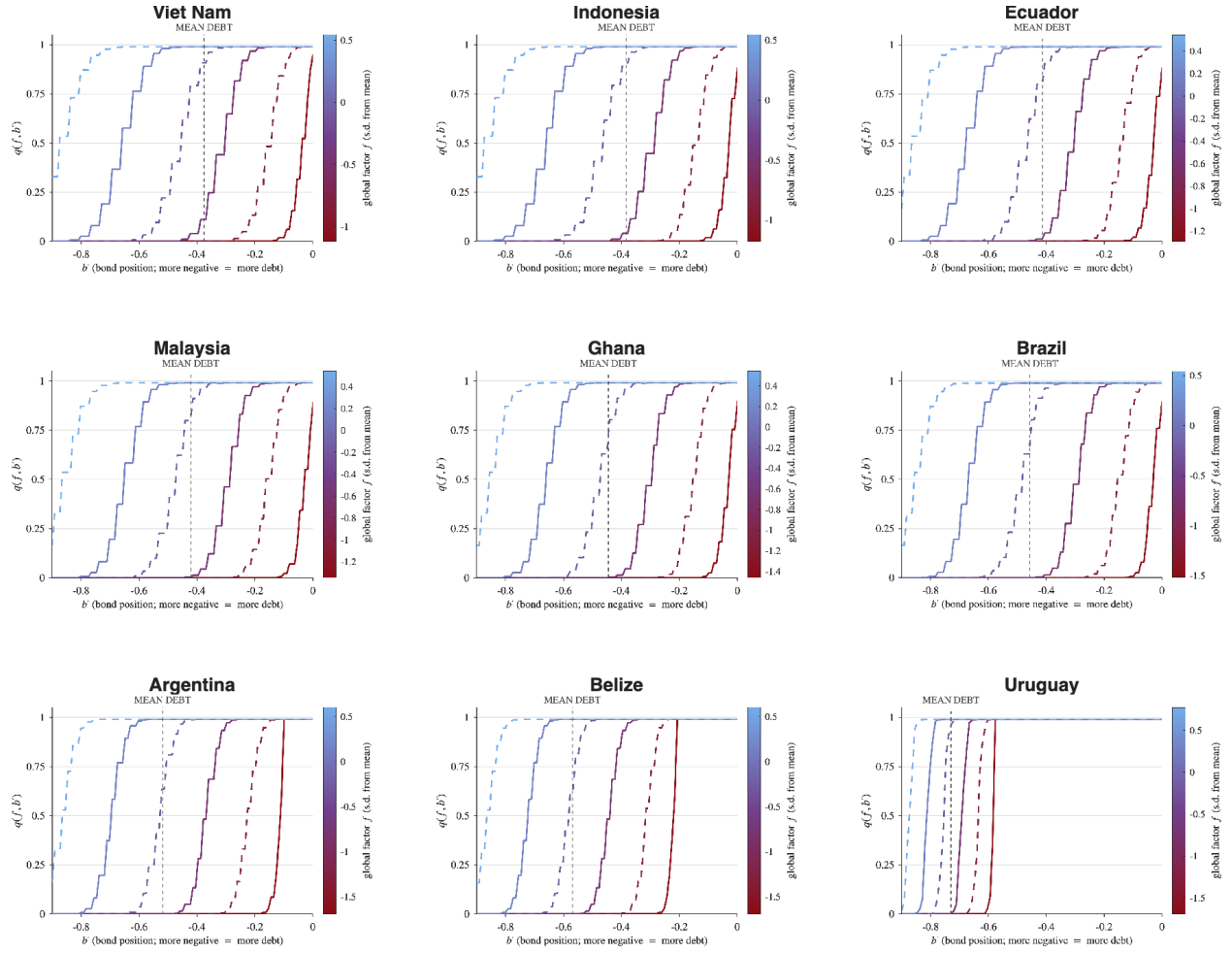


Figure S8: Bond Price Schedules – Importer-type Countries, Part 1.

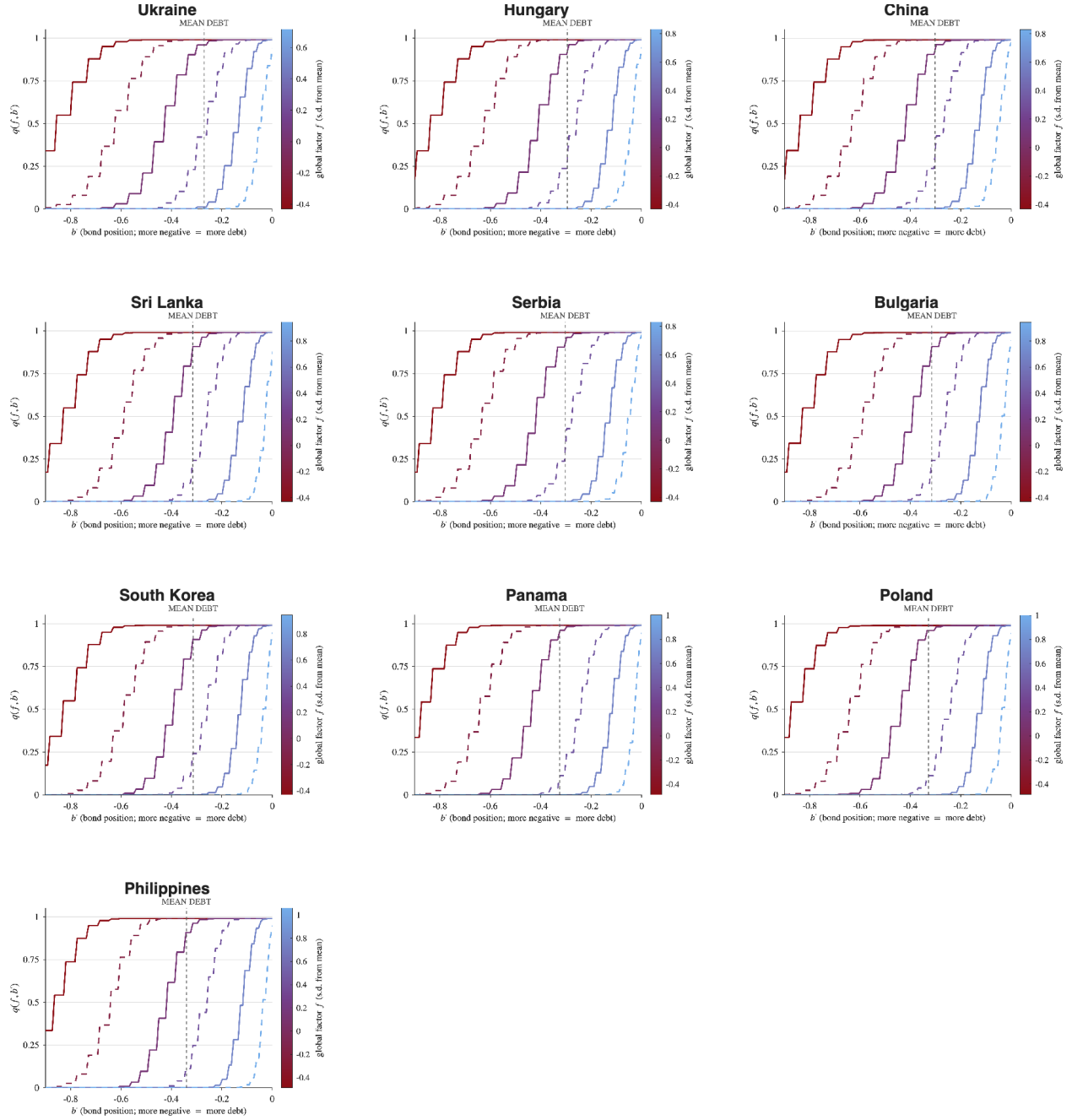
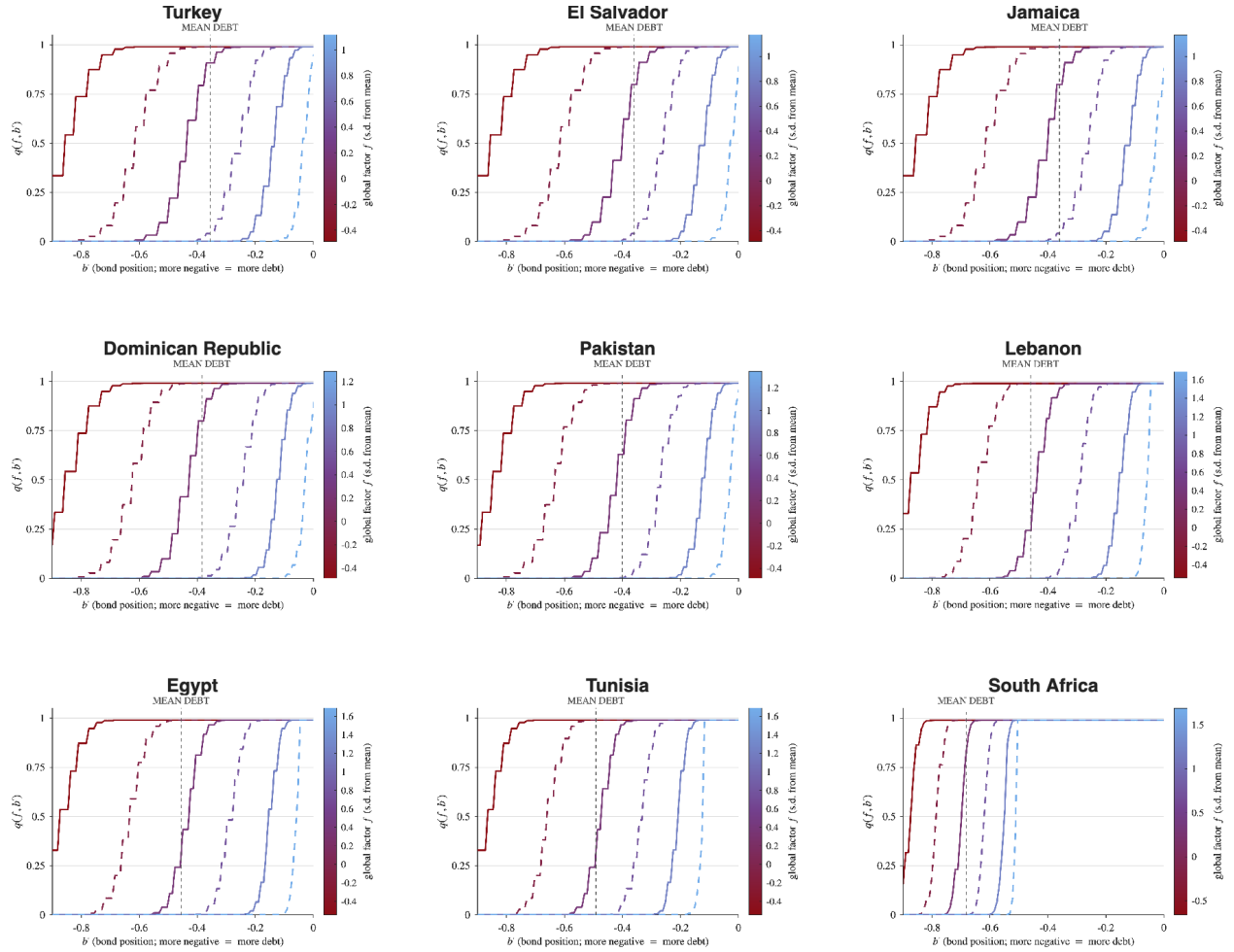


Figure S9: Bond Price Schedules – Importer-type Countries, Part 2.



S5.3 Debt Levels and the Data

Model debt represents external sovereign borrowing against commodity income, priced by risk-neutral lenders. Observed general government debt-to-GDP ratios cover both domestic and external liabilities and scale them by total output. These measures therefore differ in both coverage and normalization. The calibration does not target their correspondence: the estimation uses the level and volatility of sovereign spreads and the frequency of default, with no debt moment entering the criterion. The exporter–importer classification is also independent of debt moments, being determined by the sign of $\tilde{\nu}_i$.

To compare relative variability, we report the within-country coefficient of variation of debt. This statistic is invariant to constant rescaling, although differences in coverage and movements in the empirical GDP denominator still limit comparability. Across the 37 countries, its median is 0.54 in the model and 0.38 in the data. Simulated debt is thus more variable, but its relative variability is of the same order of magnitude as that observed. The greater variability may partly reflect the model’s one-period debt structure, which requires the entire outstanding stock to be repaid or refinanced each period.

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