Export and Import Price Shocks and Sovereign Debt

Enock W.N. Bulime*

November 10, 2026

Abstract

African governments rely heavily on commodity exports for fiscal revenue and spend heavily on commodity imports. Standard analyses summarize this exposure through the terms of trade: when export prices fall or import prices rise, the ratio deteriorates and debt increases. But export and import prices often fall simultaneously. I show that this common decline raises sovereign debt even though the terms of trade is unchanged. Using separately identified export and import price shocks for twenty African economies over 1990–2019, I find that a 10 percent export-price decline raises sovereign debt by 5.7 percentage points of GDP after four years; an equal import-price increase lowers it by 3.4 points; and a simultaneous 10 percent decline in both raises it by 6 points. Export-price shocks reduce government revenue with a delay, as commodity receipts take time to reach the budget; import-price increases lower debt largely through the GDP deflator: higher import prices raise the price level, expanding nominal GDP and reducing the debt ratio. Frameworks that summarize external price exposure through relative prices or one price at a time miss the debt that accumulates when both fall together, understating sovereign fiscal risk.

JEL classification numbers: E62, F41, H63, O55, Q33.
Keywords: export and import prices; sovereign debt; fiscal transmission; Africa; debt sustainability.

*Department of Economics, Kansas State University. Email: bulimeenock@ksu.edu.

1 Introduction

The terms-of-trade ratio summarizes external-price exposure as a single number: when export prices fall or import prices rise, the ratio deteriorates. Open-economy models (Mendoza, 1995; Schmitt-Grohé and Uribe, 2018) and empirical work on sovereign debt (Arezki and Brückner, 2012; Muhanji and Ojah, 2011) build in this compression: both prices enter only through their ratio, implying equal and opposite effects on any outcome. Di Pace, Juvenal, and Petrella (2025) challenge this for real activity outcomes: using separately identified export- and import-price shocks, they show that the two shocks move output, consumption, investment, the trade balance, and the real exchange rate differently, and that the equal-and-opposite restriction is rejected in a broad panel of developing economies. Do the same shocks move sovereign debt differently? If the restriction fails for debt as it does for real activity, existing sovereign-risk frameworks systematically understate the debt that accumulates in the most common type of external-price episode.

In three years of broad commodity price declines (1998, 2009, and 2015), between 91 and 98 percent of the 182 economies in the IMF Commodity Terms of Trade database (Gruss and Kebhaj, 2019) saw both their export and import prices fall at the same time. The standard terms-of-trade (ToT) ratio misses this for two reasons. First, it measures only how the two prices move relative to each other and ignores movements they share, so a simultaneous fall in both prices is invisible to it. Second, it constrains export- and import-price changes to have equal and opposite effects on debt by construction. The two prices affect debt through different channels: export-price declines reduce government revenue, while import-price changes affect expenditure and move the debt ratio through nominal GDP, valuation changes, and financing costs. I estimate the separate debt response to each price shock, test whether the two responses are equal and opposite as the ratio assumes, and measure the sovereign debt that simultaneous price declines generate but the ratio does not capture.

I use country-specific, separately identified export- and import-price shocks from Di Pace, Juvenal, and Petrella (2025) (hereafter DJP) for twenty African economies (Africa-20) over 1990–2019, and estimate the debt response at each horizon with panel local projections. Three features make Africa the appropriate setting: simultaneous price declines are frequent, separately identified export- and import-price shocks are available at the country level, and the IMF–World Bank Low-Income Country Debt Sustainability Framework (LIC-DSF) uses commodity-price stress scenarios to assess sovereign debt risks for these economies.

The estimates show two consequences of using the ratio alone. First, the ratio cannot distinguish which price changed. A 10 percent export-price decline raises public debt by 5.7 percentage points of GDP over four years, whereas a 10 percent import-price increase lowers debt by 3.4 percentage points. A regression on the terms-of-trade ratio alone instead assigns a debt increase of 3.9 percentage points to each type of price movement. It reverses the sign of the import-price response and collapses two shocks with opposite debt consequences into a single effect. I reject the symmetry restriction at every medium-run horizon in both the Africa-20 sample and DJP's broader 38-country panel.

Second, the ratio omits common price movements. A simultaneous 10 percent decline in export and import prices raises sovereign debt by 6 percentage points of GDP over four years, although the terms of trade does not change. These simultaneous price declines, which I call common busts, account for 57 percent of severe external-price episodes. A severe episode is one in which either the terms of trade or the sum of export and import price changes falls by more than 10 percent. Descriptive evidence supports the common-bust estimate. In the 168 cases where both prices fell and debt was observed four years later, the median four-year debt increase was 7.7 percentage points of GDP, while the median terms-of-trade change was only -3 percent. In the nine country-years in which both prices fell by more than 10 percent, the median debt increase was 13.1 percentage

points.[1]

Export-price declines raise debt partly through revenue losses that arrive with a lag; the fiscal response peaks one year after the shock (0.82 percentage points of GDP per standard-deviation shock, significant at the 5 percent level) and declines thereafter. Export-price declines raise debt by more than equal export-price increases reduce it (Table 4). On the import side, the fiscal balance and the debt ratio move in opposite directions. Higher import prices worsen the fiscal balance through procurement and subsidy costs, yet the debt-to-GDP ratio falls because higher prices raise nominal GDP faster than the debt stock grows. The paper separates responses that operate through the fiscal balance from those that arise through the nominal-GDP channel.

The symmetry rejection and the common-bust estimate hold under alternative identification strategies. Three shift-share designs, which combine predetermined trade shares with world prices, corroborate the baseline common-bust estimate without relying on DJP's structural rotation: an IMF CToT shift-share, a BACI (CEPII bilateral trade database)/Pink Sheet shift-share, and a DJP trade-weight shift-share. Identifying the separate effect of each price, however, still requires that rotation. A West Texas Intermediate (WTI) oil-price instrument for the six African oil exporters supports the sign of the export-price response at every horizon, including after partialling out global demand variation (VIX, U.S. real rates, dollar index).

Spatafora and Samaké (2012) and Levy, Ricci, and Werner (2020) are the closest precedents.[2] Spatafora and Samaké (2012) regress public debt on separate export and import commodity price indices using dynamic panel generalized method of moments (GMM) across up to 116 low- and middle-income countries. They find that debt rises with export prices only in low-income countries and with import prices in low-income commodity importers, a pattern that may reflect observed-price endogeneity, the GMM estimator's sensitivity to instrument choice, a larger and more heterogeneous country sample, or some combination. Structural identification reverses the import-price sign relative to their GMM estimates and identifies the denominator channel that drives it; what Spatafora and Samaké establish as an empirical pattern, this paper identifies as a causal effect. They infer from their revenue and expenditure estimates that co-movement of the two prices amplifies fiscal responses, but they do not estimate its effect on debt. Levy, Ricci, and Werner (2020) run panel local projections of debt on trade-weighted commodity export and import price indices across up to 194 countries. They find that import-price rises lower debt ratios in emerging-market commodity importers even as the fiscal balance worsens, a sign pattern I replicate for Africa, and Levy, Ricci, and Werner attribute it to asset sales rather than the primary balance. The fiscal decomposition and the deflator channel that explain this pattern are new. I extend both studies in four respects: I use price indices that cover both commodities and manufactured goods, not commodities alone (Section 4.1); I use DJP's country-specific structural shocks rather than treating observed prices as exogenous, and show that the debt responses hold across 50,000 draws from the full posterior of admissible rotations that I re-estimate independently; I test whether the two effects offset horizon by horizon, a restriction neither paper examines; and I construct the common-bust scenario that the ratio omits.

I also extend DJP's rejection of the symmetry restriction (established for real activity outcomes in Di Pace, Juvenal, and Petrella 2025) to the debt stock. The classic terms-of-trade literature

[1] In 1998 and 2001, sixteen and eighteen of the twenty economies, respectively, were in a common bust, yet the panel-mean terms-of-trade signal was positive.

[2] The closest precedent on spreads is Cohen, Riblier, and Rodrigues (2026), who construct export and import commodity-price shocks via a common global factor and find that both shocks raise sovereign spreads while the net component lowers them. The present paper differs in outcome (debt stock rather than spreads), identification (country-specific bilateral structural VARs rather than a common-factor decomposition), and the symmetry test and common-bust scenario.

emphasizes relative export and import prices (Harberger, 1950; Laursen and Metzler, 1950; Prebisch, 1950). Country-specific structural vector autoregression (VAR) evidence finds that the terms-of-trade ratio explains less than 10 percent of output variance (Schmitt-Grohé and Uribe, 2018), an empirical disconnect that DJP's separation of export- and import-price shocks addresses directly. For African economies, whose export baskets are concentrated in a few commodities and differ substantially across countries (Deaton, 1999), the two prices trace different fiscal channels. Arezki and Brückner (2012) and Muhanji and Ojah (2011) study external debt using a single export-price index with no import-price component. Ehrhart and Guérineau (2013) separate export and import commodity prices across 80 developing countries but study volatility and tax revenue, not debt dynamics.

I make three contributions. First, I estimate the consequences of separately identified export- and import-price shocks for the debt stock. Two shocks that both worsen the terms of trade can move public debt in opposite directions. A fiscal decomposition explains the divergence: export-price shocks raise debt through revenue losses, while import-price increases compress the debt ratio through the GDP deflator rather than the primary balance. Fiscal policy cannot directly offset this denominator channel, and a terms-of-trade regression cannot detect it. Second, I show that the terms of trade is not a sufficient statistic for external-price exposure in public-debt dynamics. Standard small-open-economy models impose this property by construction (Mendoza, 1995; Schmitt-Grohé and Uribe, 2018), and every empirical regression using the ratio inherits the assumption. I reject it at every medium-run horizon in both the Africa-20 sample and the broader 38-country panel of DJP. Because the restriction fails, estimating the two price effects separately changes the sign of the import-price response and the magnitude of both responses. Third, I show that the sufficient-statistic assumption and the LIC-DSF's netting provision both miss the common-bust scenario: a source of external-price risk invisible to the ratio that, over four years, generates 6 percentage points of debt accumulation. I provide the first estimate of the sovereign debt response to a simultaneous decline in export and import prices, constructed from separately identified shock responses.

The IMF–World Bank LIC-DSF uses commodity stress tests centered on export-price shocks and, where the same commodity drives both export receipts and import costs, allows import exposure to mitigate that shock (International Monetary Fund and World Bank, 2018).[3] Letting import-price co-movement mitigate the export-price shock is misleading in common-bust episodes: equal falls in export and import prices do not offset in the public-debt ratio. Such assessments determine borrowing access, loan terms, and debt-relief eligibility for commodity-dependent African economies. Omitting common price declines therefore understates projected debt stress, affecting G20 debt-relief decisions and fiscal-rule design across these economies. This conclusion applies to the structural fiscal channels identified over 1990–2019; generalizing to the post-pandemic environment, where the inflationary transmission of import prices may differ, requires caution.

Section 2 develops the framework, Section 3 describes the empirical model and identification, Section 4 presents the results, and Section 5 concludes.

2 Framework

The framework follows Di Pace, Juvenal, and Petrella (2025) and treats export and import prices separately instead of summarizing their effects through the terms-of-trade ratio. In the standard exportable–importable–nontradable model (Mendoza, 1995; Schmitt-Grohé and Uribe, 2018), the terms of trade is a sufficient statistic because export and import prices move in equal and oppo-

[3]The LIC-DSF applies a tailored commodity-price stress test when commodities account for at least 50 percent of exports. Where the same commodity is also imported, the framework may adjust the export-price stress for net import exposure.

site proportions by construction. DJP show that export- and import-price shocks move output, consumption, investment, the trade balance, and the real exchange rate differently. I apply this distinction to the public-debt ratio, whose response combines primary-balance effects with changes in interest costs, valuation, financing, and nominal GDP.

2.1 Notation and the Budget Identity

Countries are indexed by i and years by t. The baseline sample contains twenty African economies over 1990–2019. Response horizons are $h = 0, 1, \ldots, 8$, and $j \in \{x, m\}$ denotes export prices (x) or import prices (m). The operator Δ denotes the annual change and ln the natural logarithm.

Let $P_{x,it}$ and $P_{m,it}$ be the export and import price indices of country i in year t; write $x_{it} \equiv \Delta \ln P_{x,it}$ and $m_{it} \equiv \Delta \ln P_{m,it}$ for their annual log changes. The terms of trade is $\mathrm{ToT}_{it} \equiv P_{x,it}/P_{m,it}$, so $\Delta \ln \mathrm{ToT}_{it} = x_{it} - m_{it}$. Gross debt D_{it} evolves according to

$$D_{it} \;=\; (1 + r^{d}_{it})\, D_{i,t-1} \;-\; PB_{it} \;+\; SFA_{it}, \tag{1}$$

where r^{d}_{it} is the effective interest rate, PB_{it} the primary balance, and SFA_{it} the stock-flow adjustment, which collects valuation changes, debt transactions not recorded in the primary balance, and measurement differences. Iterating forward from the end of year $t-1$ to the end of year $t+h$ gives

$$D_{i,t+h} \;=\; \left[\prod_{\ell=0}^{h}(1 + r^{d}_{i,t+\ell})\right] D_{i,t-1} \;-\; \sum_{k=0}^{h} \omega^{(h)}_{it}(k)\,\big(PB_{i,t+k} - SFA_{i,t+k}\big), \tag{2}$$

where $\omega^{(h)}_{it}(k) \equiv \prod_{\ell=k+1}^{h}(1 + r^{d}_{i,t+\ell})$ is the interest accumulation factor between $t+k$ and $t+h$ on one unit of new borrowing. An empty product equals one, so $\omega^{(h)}_{it}(h) = 1$.

To separate the fiscal channels, write the primary balance as

$$PB_{it} \;=\; R^{x}_{it}(P_{x,it}) + R^{o}_{it} \;-\; S^{m}_{it}(P_{m,it}) - S^{o}_{it}, \tag{3}$$

where R^{x}_{it} denotes commodity-linked revenue, including royalties, profit taxes, and transfers from state-owned enterprises; S^{m}_{it} denotes import-priced expenditure, including procurement and subsidies; and R^{o}_{it} and S^{o}_{it} collect other revenue and expenditure. This decomposition is a simplifying device, not an exclusion restriction: other fiscal components can respond to prices through economic activity. Write their direct responses to the two prices as

$$\frac{\partial PB_{i,t+k}}{\partial \Delta \ln P_{x,it}} \;=\; \theta_{x,i}\,\phi_i(k), \qquad \frac{\partial PB_{i,t+k}}{\partial \Delta \ln P_{m,it}} \;=\; -\,\theta_{m,i}\,\psi_i(k), \tag{4}$$

where $\theta_{x,i} > 0$ is the eventual rise in commodity-linked revenue per unit of $\Delta \ln P_{x,it}$; $\theta_{m,i} > 0$ the eventual rise in import-priced spending per unit of $\Delta \ln P_{m,it}$; and $\phi_i(k)$ and $\psi_i(k)$ are the shares realized k years after the innovation. Commodity revenue arrives with a lag (Pieschacón, 2012; Medina and Soto, 2016), while import-priced spending adjusts on impact, so the two timing profiles differ. To isolate the primary-balance contribution, evaluate $\omega^{(h)}_{it}(k)$ on the counterfactual path where no price shock occurs, and call these baseline values $\bar{\omega}^{(h)}_{it}(k)$. Define the interest-weighted cumulative timing profiles

$$\Phi_{x,i}(h) \;\equiv\; \sum_{k=0}^{h} \bar{\omega}^{(h)}_{it}(k)\,\phi_i(k), \qquad \Psi_{m,i}(h) \;\equiv\; \sum_{k=0}^{h} \bar{\omega}^{(h)}_{it}(k)\,\psi_i(k).$$

The primary-balance contributions to the level of debt at horizon h are

$$B_{x,i}^{(h)} \;=\; -\,\theta_{x,i}\,\Phi_{x,i}(h), \qquad B_{m,i}^{(h)} \;=\; \theta_{m,i}\,\Psi_{m,i}(h), \tag{5}$$

where $B_{x,i}^{(h)}$ is the export-side contribution (the cumulative revenue loss discounted to horizon h) and $B_{m,i}^{(h)}$ is the import-side contribution (the cumulative rise in import-priced expenditure discounted to horizon h).

Equal unit increases in $\Delta \ln P_{x,it}$ and $\Delta \ln P_{m,it}$ have equal and opposite effects on the primary-balance contribution to debt at horizon h if and only if $\theta_{x,i}\,\Phi_{x,i}(h) = \theta_{m,i}\,\Psi_{m,i}(h)$.

The budget identity does not impose equality between the two primary-balance contributions. The equality requires both the exposures and their interest-weighted timing profiles to align, so it can hold at one horizon and fail at another.

2.2 From the Level of Debt to the Debt Ratio

A debt sustainability analysis (DSA) stresses the gross-debt ratio $d_{it} \equiv D_{it}/Y_{it}$, measured in percentage points of GDP, where Y_{it} is nominal GDP. The outcome is $d_{i,t+h} - d_{i,t-1}$, the cumulative change in the debt ratio from the year before the shock to horizon h.

Not every component of the debt ratio runs through the primary balance. Let $V_{j,i}^{(h)}$ denote the response to a unit change in $\Delta \ln P_{j,it}$ of all debt components other than the primary-balance contribution, including indirect fiscal responses, interest-rate effects, and stock-flow adjustments. Let $\mathcal{Y}_{j,i}^{(h)}$ be the response of nominal GDP to the same shock, and let $\bar{D}_{i,t+h}$ and $\bar{Y}_{i,t+h}$ be debt and GDP on the no-innovation path. A first-order expansion decomposes the total response of the debt ratio to price j into a primary-balance term and a remainder:

$$\underbrace{\frac{\partial(d_{i,t+h} - d_{i,t-1})}{\partial \Delta \ln P_{j,it}}}_{\text{total debt-ratio response}} \;=\; \underbrace{\frac{B_{j,i}^{(h)}}{\bar{Y}_{i,t+h}}}_{\text{primary balance}} \;+\; \underbrace{\left[\frac{V_{j,i}^{(h)}}{\bar{Y}_{i,t+h}} \;-\; \frac{\bar{D}_{i,t+h}}{\bar{Y}_{i,t+h}^{2}}\,\mathcal{Y}_{j,i}^{(h)}\right]}_{\text{remainder } \mathcal{Q}_{j,i}^{(h)}}, \qquad j \in \{x, m\}. \tag{6}$$

The remainder $\mathcal{Q}_{j,i}^{(h)}$ has two components: $V_{j,i}^{(h)}/\bar{Y}_{i,t+h}$, the effect of valuation changes, financing costs, and stock-flow adjustments on the debt numerator scaled by GDP; and $(\bar{D}_{i,t+h}/\bar{Y}_{i,t+h}^{2})\,\mathcal{Y}_{j,i}^{(h)}$, the effect of nominal GDP growth on the denominator. A price change can therefore worsen the primary balance and still reduce the debt ratio if the GDP or numerator effects in the remainder are large enough.

Summing equation (6) over j,

$$\sum_{j \in \{x,m\}} \frac{\partial(d_{i,t+h} - d_{i,t-1})}{\partial \Delta \ln P_{j,it}} \;=\; \frac{\left(B_{x,i}^{(h)} + B_{m,i}^{(h)}\right)}{\bar{Y}_{i,t+h}} \;+\; \mathcal{Q}_{x,i}^{(h)} + \mathcal{Q}_{m,i}^{(h)}. \tag{7}$$

Equality of the two primary-balance contributions sets only the first term to zero. If it holds but the remainders do not offset each other, a common price movement still moves the ratio. If it fails, the remainders can offset the first term exactly. The symmetry restriction must therefore be tested on the total debt-ratio response, not on the primary balance alone.

2.3 The Null Hypothesis

The framework generates a testable restriction: if the terms-of-trade ratio captures all external-price risk, then export- and import-price changes must have equal and opposite effects on the debt ratio. I show that this restriction must be evaluated on the total debt-ratio response: the remainder terms in equation (6) can break symmetry even when the primary balance satisfies it, and can restore it even when the primary balance does not.

Consider an equal log change $\bar{q}$ in both prices simultaneously. Because each price rises by $\bar{q}$, the total debt-ratio movement is the sum of the two partial responses from equation (6), each scaled by $\bar{q}$. The common change leaves the terms-of-trade ratio unchanged, so any mapping from prices to the debt ratio that runs only through $\ln(P_x/P_m)$ must set the sum of the two partial responses to zero.

The restriction takes the form of a null hypothesis on the panel local projection

$$d_{i,t+h} - d_{i,t-1} \;=\; \alpha_i^{(h)} + \beta_x^{(h)}\, x_{it} + \beta_m^{(h)}\, m_{it} + \mathbf{w}_{it}'\,\boldsymbol{\gamma}^{(h)} + \varepsilon_{it}^{(h)}, \tag{8}$$

where $\alpha_i^{(h)}$ is a country intercept specific to horizon h; $\mathbf{w}_{it}$ is a vector of lagged controls (specified in Section 3); $\boldsymbol{\gamma}^{(h)}$ is the corresponding coefficient vector; and $\varepsilon_{it}^{(h)}$ is the local-projection error. The coefficients $\beta_x^{(h)}$ and $\beta_m^{(h)}$ are the average cumulative responses of the debt ratio, in percentage points of GDP, to a unit log change in the export and import price, respectively, holding the other price and the controls fixed. Any terms-of-trade-only mapping imposes the restriction

$$H_0: \quad \beta_x^{(h)} + \beta_m^{(h)} = 0 \qquad \text{for each } h. \tag{9}$$

DJP test this restriction on real activity outcomes; I apply it to the debt ratio. Equation (7) shows that it cannot be verified from the primary balance alone, so I test it on the total debt-ratio response.

2.4 Relationship to a Terms-of-Trade-Only Regression

A terms-of-trade regression cannot recover the two separate price responses even with an arbitrarily large sample: it assigns zero weight to common price movements by construction, and it forces a single coefficient to represent two fiscally distinct effects. These are not sampling failures; they are consequences of the ratio specification itself.

When H_0 in equation (9) fails, a terms-of-trade regression is misspecified in two ways. First, it assigns no fitted effect to common price movements, predicting that a common bust leaves debt unchanged at every horizon.

To see the second misspecification, rotate the two price changes into a relative direction $v_{it}^r \equiv x_{it} - m_{it} = \Delta \ln \mathrm{ToT}_{it}$ and a common direction $v_{it}^c \equiv x_{it} + m_{it}$. Substituting $x_{it} = (v_{it}^r + v_{it}^c)/2$ and $m_{it} = (v_{it}^c - v_{it}^r)/2$ into equation (8) gives

$$d_{i,t+h} - d_{i,t-1} \;=\; \alpha_i^{(h)} + \frac{\beta_x^{(h)} - \beta_m^{(h)}}{2}\, v_{it}^r + \frac{\beta_x^{(h)} + \beta_m^{(h)}}{2}\, v_{it}^c + \mathbf{w}_{it}'\,\boldsymbol{\gamma}^{(h)} + \varepsilon_{it}^{(h)}. \tag{10}$$

The coefficient on v_{it}^c is zero exactly when H_0 holds. A terms-of-trade regression omits v_{it}^c; by the Frisch–Waugh–Lovell theorem, its slope is a weighted combination of the two separate price effects, where the weight $\lambda^{(h)}$ is the slope from regressing v_{it}^c on v_{it}^r after partialling out $\mathbf{w}_{it}$ and country fixed effects:

$$\lambda^{(h)} \;=\; \frac{\mathrm{Var}(\tilde{x}_{it}) - \mathrm{Var}(\tilde{m}_{it})}{\mathrm{Var}(\tilde{x}_{it}) + \mathrm{Var}(\tilde{m}_{it}) - 2\,\mathrm{Cov}(\tilde{x}_{it}, \tilde{m}_{it})}. \tag{11}$$

The numerator is the difference in residualized variances; the denominator is the residualized variance

of the terms-of-trade change. Stronger co-movement reduces the denominator and raises $|\lambda^{(h)}|$.

Second, the terms-of-trade coefficient therefore forces a single number to capture two fiscally distinct effects. Specifically, $\beta^{(h)}_{\text{ToT}}$, the coefficient a terms-of-trade regression recovers, satisfies

$$\beta^{(h)}_{\text{ToT}} \;=\; \frac{\beta^{(h)}_x - \beta^{(h)}_m}{2} \;+\; \frac{\beta^{(h)}_x + \beta^{(h)}_m}{2}\,\lambda^{(h)}, \tag{12}$$

where the first term is the average of the two separate effects and the second is a bias that is zero only when H_0 holds. Export-price changes are 2.79 times as volatile as import-price changes conditional on controls and country fixed effects,[4] so the bias term is nonzero and the ToT coefficient assigns the wrong sign to the import-price response.

3 Empirical Model

I describe the local-projection specification, explain why structural shocks are preferred over observed prices, and define the common-bust scenario used to evaluate the two failure modes of the terms-of-trade ratio.

3.1 The Estimating Equation

Equation (8) is estimated by panel local projection (Jordà, 2005; Jordà and Taylor, 2025) for each horizon $h = 0, \ldots, 8$. The control vector $\mathbf{w}_{it}$ contains two lags of the outcome, both price changes, domestic real GDP growth, CPI inflation, the change in the real effective exchange rate, and world real GDP growth, giving 14 controls. Lagged debt captures persistence, lagged prices capture collection lags, and lagged world growth absorbs the slow global cycle. The controls enter only in lags because contemporaneous growth, inflation, and the exchange rate are channels through which the price shocks move the debt ratio. Both prices must enter jointly, since dropping one assigns the common global factor to the retained price series.

The long-difference outcome $d_{i,t+h} - d_{i,t-1}$ sets $d_{i,t-1}$ as the pre-shock baseline, so $\hat{\beta}^{(h)}$ measures cumulative debt accumulation from the year before the shock rather than the debt level at $t+h$; this removes cross-country heterogeneity in pre-shock debt stocks that is unrelated to price shocks (Jordà and Taylor, 2025; Piger and Stockwell, 2023). I use panel local projections instead of a VAR or first-differenced GMM because they are robust to dynamic misspecification and reliable for inference under the persistence in this sample.[5] Fixed-effect local projections remain Nickell-biased (Mei, Sheng, and Shi, 2026; Jordà and Taylor, 2025), so I treat conventional intervals as descriptive; split-panel jackknife coefficients are smaller in absolute value, consistent with Nickell bias attenuating toward zero. I retain country effects rather than differencing them out, because each country's commodity export structure is correlated with its export-price exposure.[6] Two lags match the AR(2) description of the debt stock. Lag orders from one to four preserve the baseline sign and ranking; with five lags the two coefficients at $h = 2$ are of similar size and symmetry is no longer rejected. I report year fixed effects only as a diagnostic, because with $G = 20$ they also absorb much of the usable common variation in commodity prices. Global demand shocks account for approximately 30

[4] Estimated from the local-projection sample residuals, $\hat{\lambda} = 1.20$; the terms-of-trade regression on observed prices is therefore biased whenever the restriction fails. Section 4.2.3 and Table 3 measure the consequences with the structural estimates.

[5] Plagborg-Møller and Wolf (2021), Montiel Olea and Plagborg-Møller (2021), and Montiel Olea, Plagborg-Møller, Qian, and Wolf (2026) establish these properties in the panel setting.

[6] Because the controls include $d_{i,t-1}$, long differencing is a reparameterization, not a bias correction; retaining country effects preserves cross-sectional differences in commodity export structure (Ugarte-Ruiz, 2026).

percent of output variance at long horizons (Di Pace, Juvenal, and Petrella, 2025, Table 7), so year fixed effects remove valid identifying variation alongside global confounds. Contemporaneous primary balances, trade volumes, fiscal-rule indicators (Alonso and others, 2025), interest expenditure, and the current account enter only as robustness checks added one at a time, because each is a potential outcome of the price shocks.

The residual $\varepsilon_{it}^{(h)}$ is serially correlated by construction, because at horizon h the left-hand side spans h years of later innovations. I cluster standard errors by country, allowing arbitrary serial correlation within country (Cameron, Gelbach, and Miller, 2008). Driscoll–Kraay intervals complement the clustered standard errors by allowing for cross-sectional dependence (Driscoll and Kraay, 1998); two-way clustering is not well supported with twenty clusters. Petrella, Juvenal, and Di Pace (2026) use a heteroskedasticity- and autocorrelation-consistent estimator that adapts its bandwidth to the cross-sectional dimension (Bai, Choi, and Liao, 2024); with twenty clusters, Driscoll–Kraay inference together with the restricted wild cluster bootstrap (MacKinnon, Nielsen, and Webb, 2023) is more reliable. The results hold across fixed-effect, split-panel jackknife, lag-order, and structural-shock specifications. Unless stated otherwise, every table covers the twenty economies over 1990–2019 with country fixed effects and country-clustered standard errors; stars denote significance at the 10, 5, and 1 percent levels using $t(19)$ critical values. p(sym) denotes the p-value of the test that the export and import coefficients sum to zero: $H_0 : \beta_x + \beta_m = 0$ for price changes and $H_0 : \beta_{u_x} + \beta_{u_m} = 0$ for structural shocks.

3.2 Structural Shocks and the Common-Bust Scenario

Regressing debt on observed price changes confounds country-specific supply shocks with global demand fluctuations that move both prices and budgets simultaneously. I replace observed price changes with country-specific structural shocks that isolate each price channel, and define the common-bust scenario that the terms-of-trade ratio cannot detect.

The observed price changes share a large global component (Alquist, Bhattarai, and Coibion, 2020; Delle Chiaie, Ferrara, and Giannone, 2022). No single country moves world prices (Deaton and Miller, 1993), but a common demand factor can move prices and African budgets together through aid, remittances, and financing. I use DJP's structural export- and import-price shocks in place of the observed price changes, $u_{x,it}$ and $u_{m,it}$, used as published and with approximately unit variance in this sample:

$$d_{i,t+h} - d_{i,t-1} \;=\; \alpha_i^{(h)} + \beta_{u_x}^{(h)}\, u_{x,it} + \beta_{u_m}^{(h)}\, u_{m,it} + \mathbf{w}_{it}'\, \boldsymbol{\gamma}^{(h)} + \varepsilon_{it}^{(h)}, \tag{13}$$

The coefficient $\hat{\beta}_{u_j}^{(h)}$ is the response at horizon h to a structural shock in year t. Because the DJP price innovations have a median AR(1) of about 0.47, $\hat{\beta}_{u_j}^{(h)}$ reflects both the direct fiscal effect and the fiscal consequences of persistence. With lags of the shocks replacing lags of the price changes, DJP separate the country price innovations into export-price, import-price, global-demand, and global-supply disturbances using sign, relative-elasticity, and narrative restrictions (Di Pace, Juvenal, and Petrella, 2025; Antolín-Díaz and Rubio-Ramírez, 2018). These restrictions are credible for this sample: the price-taking assumption is plausible (Deaton and Miller, 1993), the narrative sign pattern matches the 2015–16 Brent collapse, and leave-Africa-out and WTI reconstructions recover the same direction of debt response. The shocks are sensitive to the choice of prior distributions (Giacomini, Kitagawa, and Read, 2022). DJP identify shocks on detrended real-activity variables, targeting cyclical co-movement, while the debt projections here use long differences of the debt stock, capturing medium-run accumulation. I construct long-difference outcomes ($d_{i,t+h} - d_{i,t-1}$) on the full panel before applying the estimation-sample restriction; the outcome window and the shock

sample therefore can differ at long horizons. Causal results are conditional on DJP's maintained assumptions: I treat their published posterior-median shocks as exogenous interventions, and the reported standard errors do not reflect uncertainty in DJP's posterior draws.

DJP's structural rotation narrows the common movement that the terms of trade discards but does not eliminate it. The common direction still accounts for 56 percent of the joint variance of the shocks, down from 68.2 percent in the raw price innovations, because DJP's sign restrictions require each shock to dominate its own price but do not impose zero cross-price correlation. The most frequent sign configuration in the shock data is a common bust (both u_x and u_m negative), which occurs in 30.3 percent of country-years, and is the more persistent of the two, with a median AR(1) of 0.468 against 0.312 for the terms-of-trade direction. In 1998 and 2001, the worst bust years in the sample, the panel-mean terms-of-trade shock had the opposite sign from the strongly negative common shock.[7] African import baskets include food and fuel subject to tariffs, subsidies, and administered prices, so a world-price shock maps imperfectly into domestic prices (Minot, 2011; Gilbert, 2011); that measurement gap attenuates the structural import coefficient.

To compare the two price responses in the same units, let $\mathbf{A}$ be the 2×2 impact matrix whose element a_{jl} is the response of $\Delta \ln P_{j,it}$ to a unit shock $u_{l,it}$. I estimate it by two $h = 0$ local projections of $\Delta \ln P_x$ and $\Delta \ln P_m$ on the structural shocks and the controls: $\hat{\mathbf{A}} = \left[\begin{smallmatrix} 0.096 & 0.020 \\ 0.017 & 0.036 \end{smallmatrix}\right]$, where rows index export and import prices and columns index export- and import-price shocks, with diagonal elements $\hat{a}_{xx} = 0.096$ and $\hat{a}_{mm} = 0.036$.[8] The off-diagonal elements are small but not negligible: with a diagonal $\hat{\mathbf{A}}$ the common-bust solution would give 9.1 points rather than 6.0. The shock vector solving $\hat{\mathbf{A}}\mathbf{u}^* = (-0.10, -0.10)'$ is $\mathbf{u}^* \approx (-0.52, -2.51)'$. A single-leg scenario rescales each response to a 10 percent move in its own price, $\tilde{\beta}_{u_j}^{(h)} = \beta_{u_j}^{(h)} \, 0.10/\hat{a}_{jj}$. A common bust cannot be built by adding the two single-leg responses, because each structural shock also moves the other price. The common-bust scenario instead solves $\mathbf{A}\mathbf{u}^* = \boldsymbol{\delta}$ for the shock combination $\mathbf{u}^*$ that lowers both prices by exactly 0.10 log points on impact, $\boldsymbol{\delta} = (-0.10, -0.10)'$, and its debt response is $\boldsymbol{\beta}^{(h)\prime}\mathbf{A}^{-1}\boldsymbol{\delta}$ with $\boldsymbol{\beta}^{(h)} \equiv (\beta_{u_x}^{(h)}, \beta_{u_m}^{(h)})'$.

4 Empirical Analysis

This section presents the main results. I describe the sample and data, report the reduced-form and structural estimates, translate the estimates into stress scenarios that expose the two failures of the terms-of-trade ratio, test for boom-bust asymmetry, examine whether structural shocks predict sovereign restructuring, and assess robustness to alternative identification strategies.

4.1 Data

The panel covers twenty African economies over 1990–2019, listed in Table A1. Countries are selected on data availability, not on debt distress, LIC-DSF eligibility, or borrowing. Of the 38 developing economies in the DJP panel, 20 are African, and 12 of these are on the World Bank LIC-DSF list. The price indices begin in 1990; with two lags of the differenced controls, the estimation sample

[7] Splitting the reduced-form sample by whether $\Delta \ln P_x$ and $\Delta \ln P_m$ share a sign confirms the sign divergence directly: the symmetry restriction is rejected strongly ($p < 0.011$ at every horizon) in the 84 percent of country-years where the two prices move together, but never in the divergent-move country-years, where the ratio is well defined. Because this split conditions on the regressors and the divergent-move subsample is only 16 percent of the panel, the test there has little power, and the absence of a rejection is descriptive.

[8] The symmetry test in equation (9) is stated in price units: $\beta_x^{(h)} + \beta_m^{(h)} = 0$. Its analog in shock units is $\beta_{u_x}^{(h)} + \beta_{u_m}^{(h)} = 0$. The two are equivalent only when the diagonal elements of $\mathbf{A}$ are equal. Here $\hat{a}_{xx} = 0.096$ and $\hat{a}_{mm} = 0.036$, so the shock-unit test is an approximation; I report it as a summary statistic alongside the price-unit test.

covers 1993–2019, yielding $N = 441$ country-year observations.

The sample of twenty countries follows from DJP's bilateral VAR identification, not from an independent sampling criterion: the structural shocks are available only for the economies in their panel. Sections 4.2.3 and 4.3 report results for DJP's broader 38-country panel and for shift-share designs that extend the rejection to a 51-country global sample; the symmetry rejection and the common-bust debt response hold in both broader samples.

The terms-of-trade ratio conflates two fiscally distinct channels: a rise in export prices (e.g., oil) raises commodity-linked fiscal receipts, whereas a rise in import prices (e.g., food) raises procurement costs; both enter the ratio but reach the fiscal accounts differently. The analysis requires separate export- and import-price series.

I use DJP's country-specific bilateral export and import price indices (Di Pace, Juvenal, and Petrella, 2025). Each index weights 62 disaggregated commodity and manufacturing price series by the country's lagged, time-varying trade shares and is measured in U.S. dollars deflated by the U.S. CPI, following the IMF Export and Import Price Manual (International Monetary Fund, 2009). Export baskets are more concentrated than import baskets: the mean Herfindahl index is 0.27 for exports and 0.07 for imports, which helps explain why a single import-price shock has a smaller effect than an equal export-price shock. No single commodity dominates the import basket for any country in the sample. Official unit-value terms-of-trade series are sparse for this sample and can conflate price changes with shifts in quality or composition; the bilateral indices avoid that problem. The structural export- and import-price shocks $u_{x,it}$ and $u_{m,it}$ are DJP's published posterior-median shocks, taken directly from their replication data; I do not re-estimate their VAR. All causal results are conditional on DJP's identification assumptions.

Two features of the price data shape the analysis. First, export prices are more than twice as volatile as import prices: the pooled standard deviation of annual changes is 0.137 for exports and 0.050 for imports, a ratio of 2.76, against 1.36 in the eighteen non-African economies of the same database. Second, the two prices often fall together. The first principal component (PC1) of the joint $(\Delta \ln P_x, \Delta \ln P_m)$ innovation accounts for 93 percent of the variance, but its loadings are $(0.975, 0.220)$: the common direction is predominantly an export-price factor that draws import prices with it.[9] In 1998, sixteen of twenty economies saw both indices fall by more than 5 percent; fourteen did so in 2009 and again in 2015. In each of those years the panel-mean terms-of-trade change stayed well inside conventional stress thresholds. I define a severe episode as one in which either the terms of trade or the sum of the two price changes falls by more than 10 percent. Fifty-seven percent of severe episodes are common declines that a terms-of-trade screen does not detect: of the 165 severe episodes, 94 are common-only (missed by the ToT screen), 64 are flagged by both screens, and 7 are detected only by the terms of trade. This screening gap is cyclical rather than secular. The median country-level terms-of-trade trend over 1990–2019 is mildly positive, so the aggregate price ratio has no material trend. Substantial debt accumulation follows common-decline episodes, which the ratio does not detect.

General-government gross debt, total revenue, total expenditure, the primary balance, and interest payments come from the April 2025 IMF World Economic Outlook (WEO) (International Monetary Fund, 2025). All fiscal aggregates are on the general-government perimeter; central-government series that exist for earlier years in some economies are not used in their place; the perimeter is therefore consistent across countries. Gross debt is the stock the LIC-DSF public analysis stresses (International Monetary Fund and World Bank, 2018). Gross rather than net debt is appropriate because debt-service obligations are assessed on the full liability stock, while liquid

[9] Di Pace, Juvenal, and Petrella (2025) report separate PC1 shares of 0.802 and 0.928 for each price series on detrended levels; the joint PC1 here is a different object and is not comparable to those figures.

assets are evaluated separately through staff judgment. A single WEO vintage avoids inconsistencies that arise when debt and GDP come from different releases. Mean gross debt is 54.5 percent of GDP, and debt varies more within countries over time than across countries, which motivates country effects and local projections.

The extreme values in the sample are WEO observations for Equatorial Guinea in 1991–94, just before its oil boom, and Sudan in 1992. Only three of them enter the estimation sample, and dropping Equatorial Guinea through 1995 and Sudan through 1993 changes the $h = 4$ scenarios by at most 0.2 points, to 5.7 points for the export bust, -3.6 for the import-cost shock, and 6.2 for the common bust (Table A9).

Nominal GDP, real GDP, the GDP deflator, the real effective exchange rate, and the bilateral dollar exchange rate come from the Global Macro Database (Müller, Xu, Lehbib, and Chen, 2025), which provides internally consistent national accounts and exchange-rate series and avoids the base-year revision discontinuities that affect WEO vintages for several Africa-20 economies. Effective tax rates on capital and labor income come from Bachas, Fisher-Post, Jensen, and Zucman (2026), whose macro-historical database constructs revenue-based effective rates from national accounts and government finance statistics; these are the only rates available at annual frequency for Africa over the study period. Domestic government debt issuance comes from the African Debt Database (Manger and others, 2025); coverage and cluster counts at each horizon are reported in the relevant tables. External disbursements to the public sector come from the World Bank International Debt Statistics (IDS) (World Bank, 2026b). The IDS dollar-denominated debt stock also supplies the predetermined foreign-currency share used in the valuation test; it is fixed at each country's 1990–95 average to limit endogeneity with the price shocks.

Sovereign restructuring years come from the Horn–Reinhart–Trebesch database of Chinese, Paris Club, and private-creditor restructurings (Horn, Reinhart, and Trebesch, 2022; Asonuma and Trebesch, 2016). The database counts symbolic Chinese restructurings but not symbolic Paris Club deals, and flags both the start and the end year of private-creditor deals; every one of the twenty economies has at least one restructuring year over 1990–2019. Default spells come from the Asonuma–Trebesch dataset (Asonuma and Trebesch, 2016).

4.2 Results

I report results in five steps: reduced-form estimates that document the raw price-debt relationship; structural estimates that isolate the causal effect of each price shock; stress scenarios that expose the two failures of the terms-of-trade ratio; a boom-bust asymmetry test; and a restructuring test that asks whether the same shocks predict sovereign distress.

4.2.1 Reduced-Form Estimates

A regression of debt on the observed price changes, equation (8), gives an import coefficient of -66.7 percentage points of GDP per unit log change at $h = 2$, against -9.7 for exports (Table 1). An equal 10 percent decline in both prices is associated with a 7.6-point increase in debt at $h = 2$, significant at the 1 percent level; the reduced-form symmetry restriction is rejected from $h = 1$ onward but not at $h = 0$. Stacked Frisch–Waugh–Lovell tests strongly reject joint irrelevance of each price block. The large import coefficient does not capture a direct fiscal channel; it captures co-movement: in global commodity episodes both prices fall together, so import-price declines predict debt rises not because lower import prices cause debt accumulation but because they signal the simultaneous export-price collapse that does. Large import-price innovations coincide with these episodes (the 2008 energy shock, the 2010–11 food-price surge, and the 2014–16 collapse), which move both prices

together and leave the terms of trade nearly unchanged. Adding year fixed effects moves the $h = 2$ import coefficient to 39.22, with an interval that spans zero. The structural shocks, orthogonal to global demand and supply, attenuate the import response even without year effects, while the export-price coefficient is stable across all specifications. I use the structural shocks in all subsequent specifications.

Table 1: Reduced-form local projections: gross government debt

	Horizon h (years)								
	$h = 0$	$h = 1$	$h = 2$	$h = 3$	$h = 4$	$h = 5$	$h = 6$	$h = 7$	$h = 8$
Export price ($\hat{\beta}_x^{(h)}$)	−12.57**	−14.95**	−9.66**	−17.82**	−26.25***	−22.44***	−19.39***	−20.91***	−15.80***
	(5.72)	(5.80)	(4.07)	(7.04)	(9.04)	(6.11)	(5.90)	(6.42)	(5.45)
Import price ($\hat{\beta}_m^{(h)}$)	−2.51	−37.81**	−66.67***	−56.23**	−55.72***	−47.99**	−49.68***	−48.52**	−35.40*
	(19.63)	(16.59)	(18.86)	(22.86)	(17.76)	(18.99)	(16.70)	(18.49)	(17.61)
$p(H_0 : \hat{\beta}_x + \hat{\beta}_m = 0)$	0.368	0.001	<0.001	0.001	<0.001	0.001	<0.001	<0.001	0.002
Observations	441	441	441	441	441	421	401	381	361

Notes: Twenty African economies, 1990–2019. Long-difference local projections of general-government gross debt with country fixed effects. Units: percentage points of GDP per unit $\Delta \ln P$. Standard errors clustered at the country level in parentheses. The symmetry row tests whether the export- and import-price coefficients sum to zero at each horizon. Restricted wild cluster bootstrap inference is in Table **??**. *, **, *** denote significance at the 10, 5, and 1 percent levels, two-tailed, using $t(G - 1)$ critical values.

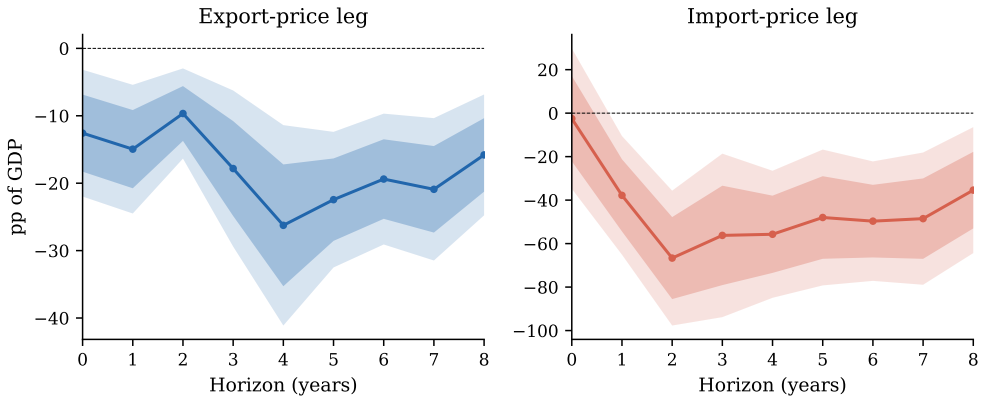


Figure 1: Reduced-form long-difference local projections of general-government gross debt, both prices entered jointly. Left: export-price response; right: import-price response. Darker and lighter shading are 68 and 90 percent pointwise country-clustered confidence bands. Units: percentage points of GDP per unit $\Delta \ln P$.

4.2.2 Structural Estimates

Throughout this section a positive structural export-price shock ($u_x > 0$) is a surprise increase in export prices, and a positive import-price shock ($u_m > 0$) is a surprise increase in import prices. Both are normalized to approximately unit variance in this sample, so coefficients $\beta_{u_x}^{(h)}$ and $\beta_{u_m}^{(h)}$ from equation (13) measure the debt response per standard-deviation shock. Negative coefficients on u_x therefore mean that export-price increases lower debt, or equivalently that export-price declines raise it.

A one-standard-deviation export-price decline raises gross debt at every horizon through $h = 4$, by 5.49 points of GDP at $h = 4$ (Table 2; Figure 2). A one-standard-deviation import-price increase lowers debt by 1.24 points at $h = 4$, individually significant at the 5 percent level at $h = 2$ and at the 10 percent level at $h = 3$ and $h = 4$, and indistinguishable from zero at $h = 0$ and $h = 1$. I reject the restriction that the two responses sum to zero at the 1 percent level from $h = 0$ through $h = 4$. The

$p(H_0)$ row in Table 2 reports the p-value for $H_0 : \hat{\beta}_{u_x}^{(h)} + \hat{\beta}_{u_m}^{(h)} = 0$, the standardized-shock version of the symmetry restriction. The same restriction stated in price units, equation (9), evaluated through the common-bust interval in Table 3, is rejected from $h = 1$ onward (the $h = 0$ interval is $[-1.65, 5.00]$). Across the 50,000 posterior draws of the Africa-20 VAR, 76.6 percent reject at $h = 1$ and at least 94.2 percent at $h = 2$ to $h = 4$, against 16.4 percent at $h = 0$. Driscoll–Kraay standard errors, which allow for cross-sectional dependence, reject at the 5 percent level through $h = 6$, and a restricted wild cluster bootstrap with Webb weights through $h = 7$ (Table A10). The joint rejection at every horizon reflects the magnitude of the export-price response: $\hat{\beta}_{u_x}^{(h)}$ is large and negative throughout, while $\hat{\beta}_{u_m}^{(h)}$ is never large enough in absolute value to satisfy $\hat{\beta}_{u_m} = -\hat{\beta}_{u_x}$, even at horizons where it is individually insignificant. Di Pace, Juvenal, and Petrella (2025) reject the analogous restriction $b_x = -b_m$ for real activity outcomes. The rejection survives dropping outcome years after 2019, leaving the common-bust estimate at 5.7 points. Split-panel jackknife and pooled estimators, which address Nickell bias in different ways, give the same sign pattern and ranking. When each structural response is rescaled to a 10 percent move in its own price using the impact-matrix formula in Section 3.2 (Table 3), a fall in export prices raises debt by 5.7 points after four years and a rise in import prices lowers it by 3.4 points. Ten percent is the common-bust scenario size used throughout. The LIC-DSF Guidance Note expresses its revenue and growth elasticities per 10-percentage-point price move, although its default shock is calibrated as one standard deviation from IMF commodity price-forecast distributions.

Table 2: Structural price shocks (DJP): gross government debt

	Horizon h (years)								
	$h = 0$	$h = 1$	$h = 2$	$h = 3$	$h = 4$	$h = 5$	$h = 6$	$h = 7$	$h = 8$
$\hat{\beta}_{u_x}$: structural export shock	-2.56^{***}	-3.83^{***}	-4.09^{***}	-4.29^{***}	-5.49^{***}	-5.35^{***}	-4.17^{***}	-3.47^{***}	-2.52^{***}
	(0.85)	(0.94)	(0.96)	(1.11)	(1.19)	(1.21)	(0.94)	(0.74)	(0.68)
$\hat{\beta}_{u_m}$: structural import shock	-0.14	-0.95	-1.48^{**}	-1.14^{*}	-1.24^{*}	-0.99	-0.74	0.19	0.23
	(0.71)	(0.70)	(0.70)	(0.66)	(0.70)	(1.09)	(0.76)	(0.99)	(0.92)
$F(H_0 : \hat{\beta}_{u_x} + \hat{\beta}_{u_m} = 0)$	8.38	18.21	24.90	21.93	33.54	10.17	17.04	7.77	3.10
$p(H_0 : \hat{\beta}_{u_x} + \hat{\beta}_{u_m} = 0)$	0.009	<0.001	<0.001	<0.001	<0.001	0.005	<0.001	0.012	0.094
Observations	441	441	441	441	441	421	401	381	361

Notes: Structural local projections using DJP export- and import-price shocks. Long-difference specification with country fixed effects and baseline controls. Units: percentage points of GDP per one standard deviation of the shock. Standard errors clustered at the country level in parentheses. Symmetry rows report the clustered $F(1, 19)$ statistic and p-value for $H_0 : \hat{\beta}_{u_x} + \hat{\beta}_{u_m} = 0$ (equal one-standard-deviation responses), not the reduced-form restriction in Table 1. For comparison, Di Pace, Juvenal, and Petrella (2025, Table 4) reject the algebraically related restriction $b_x = -b_m$ for real activity with F-statistics from 4.12 (consumption) to 39.19 (real exchange rate). *, **, *** denote significance at the 10, 5, and 1 percent levels, two-tailed, using $t(G - 1)$ critical values.

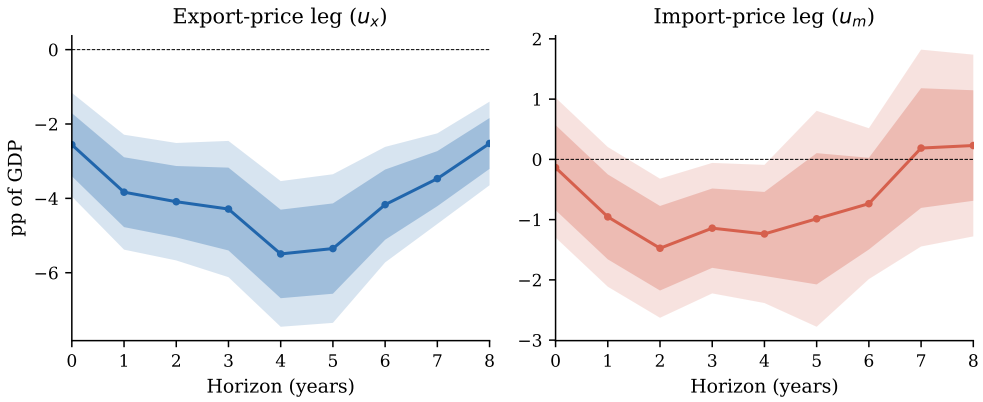


Figure 2: Impulse responses of general-government gross debt to DJP structural export-price shocks (left panel) and import-price shocks (right panel), entered jointly. Shaded areas as in Figure 1. Units: percentage points of GDP per one-standard-deviation shock.

4.2.3 Two Failures of the Terms of Trade

The structural estimates establish the sign and magnitude of each price channel. I now use them to construct debt scenarios that expose where and by how much the terms-of-trade ratio misleads a risk assessment. Two failure modes emerge: a source-identification failure, in which the ratio misattributes the separate export and import effects, and a common-shock failure, in which a simultaneous decline in both prices generates debt accumulation that the ratio records as zero.

Table 3 and Figure 3 translate the structural estimates into debt scenarios and compare them with the terms-of-trade regression. The composition map predicts the debt change for any combination of export- and import-price movements from the two structural coefficients, separating the two price effects rather than collapsing them into the ratio. The stress map applies this decomposition horizon by horizon.

Table 3: The two failure modes at every horizon, $h = 0$–4 (source-identification failure and common-shock failure)

	Source failure			Common-shock failure
h	Export bust	Import cost	ToT-only	Common bust
$h = 0$	2.66***	−0.38	1.56***	1.67
	[0.93, 4.38]	[−4.18, 3.43]	[0.52, 2.59]	[−1.65, 5.00]
$h = 1$	3.98***	−2.62	2.40***	4.39**
	[2.07, 5.90]	[−6.42, 1.17]	[1.17, 3.62]	[0.88, 7.90]
$h = 2$	4.25***	−4.05**	2.26***	5.83***
	[2.29, 6.21]	[−7.83, −0.27]	[1.51, 3.02]	[2.36, 9.30]
$h = 3$	4.45***	−3.14*	3.05***	5.10***
	[2.19, 6.72]	[−6.68, 0.41]	[1.71, 4.38]	[1.91, 8.28]
$h = 4$	5.71***	−3.40*	3.93***	5.96***
	[3.28, 8.13]	[−7.16, 0.35]	[2.12, 5.73]	[2.73, 9.20]

Notes: Percentage points of GDP; Figure 3 plots the $h = 4$ row. Brackets are 95 percent delta-method intervals (normal critical value, $\hat{\mathbf{A}}$ held fixed, country-clustered covariance of the two leg coefficients, $G = 20$); the $h = 4$ wild cluster bootstrap interval in the text re-estimates $\hat{\mathbf{A}}$ in each draw. Stars are from $t(G-1)$ tests on the same delta-method standard errors. *, **, *** denote significance at the 10, 5, and 1 percent levels.

The first failure arises from the source of the price movement: each single-leg scenario rescales the structural shock vector to produce a 10 percent movement in its own price. Because the off-diagonal elements of the impact matrix $\mathbf{A}$ also move the other price, the implied cross-price effect is approximately 5.6 percent for the export shock (which slightly raises import prices) and −1.8 percent for the import shock (which slightly lowers export prices). These cross-price movements are

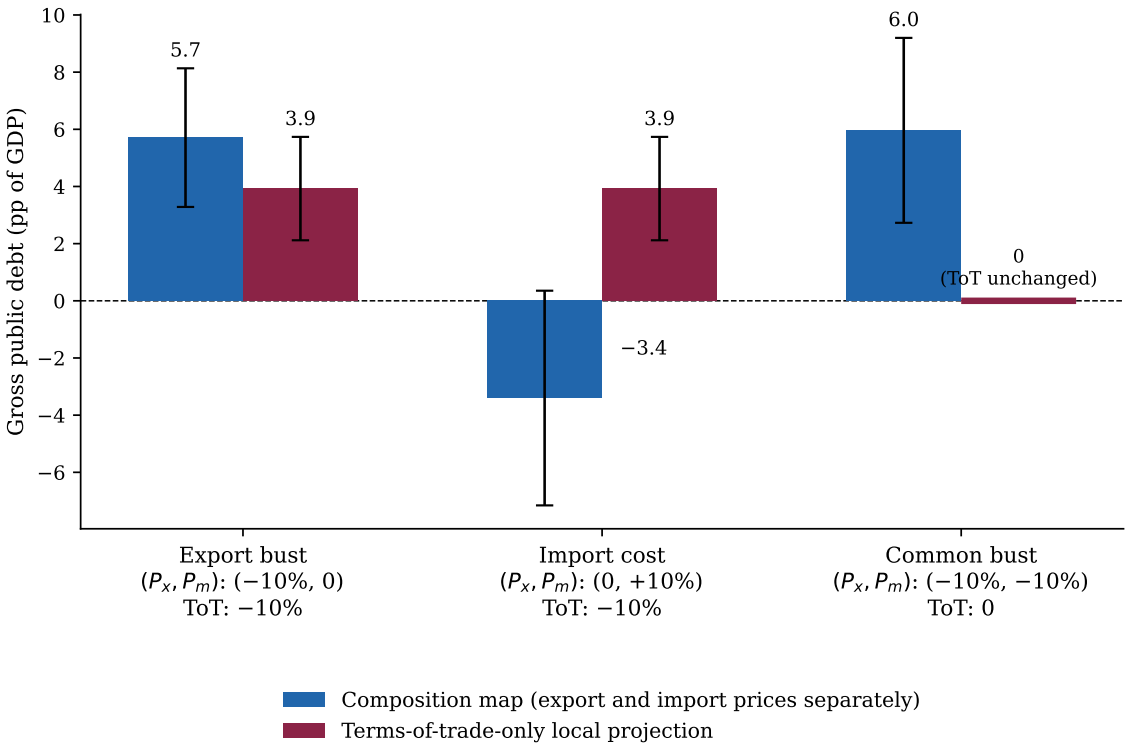


Figure 3: Debt responses to three price scenarios using the DJP structural estimates. Bars show the $h = 4$ change in gross public debt (percentage points of GDP). Left and center bars: single-leg shocks scaled to a 10 percent movement in their own price (cross-price effects from $\hat{\mathbf{A}}$ are reported in Table 3). Right bar: common bust (both prices fall 10 percent). Navy bars: composition-map prediction; maroon bars: terms-of-trade-only projection. Error bars are 95 percent delta-method confidence intervals (impact matrix held fixed).

reported in Table 3. The bars in Figure 3 represent a pure export-price shock and a pure import-price shock, not two scenarios with identical terms-of-trade movements.[10] At $h = 4$ the first raises debt by 5.7 points of GDP (95 percent interval $[3.3, 8.1]$) and the second lowers it by 3.4 points (interval $[-7.2, 0.4]$). Single-leg intervals use the normal 1.96 multiplier; stars throughout use $t(19)$ critical values. The terms-of-trade regression assigns each scenario an increase of 3.9 points (using the same structural-shock controls as the baseline; with price-change controls the figure is 3.4, consistent with equation (12)). The opposite signs reach statistical significance at $h = 2$, where the import-cost interval is $[-7.8, -0.3]$, and appear only in the point estimates at other horizons.

The second failure concerns common movements. Di Pace, Juvenal, and Petrella (2025) show theoretically that a common shifter of export and import prices leaves the terms-of-trade ratio invariant yet still affects the economy through the external position; the estimates here measure that effect for debt. When both prices fall by 10 percent, the terms of trade remains unchanged and the terms-of-trade regression predicts no change in debt. The structural estimates imply a rise of 6 points of GDP (95 percent interval $[2.7, 9.6]$, wild cluster bootstrap re-estimating $\hat{\mathbf{A}}$ in each draw), with an interval excluding zero from $h = 1$ onward. Allowing for cross-sectional dependence widens the interval: with Driscoll–Kraay standard errors it is $[-0.6, 12.5]$ at $h = 4$, insignificant at the 10 percent level, and excludes zero only at $h = 2$ (Table A10), consistent with common busts affecting many economies in the same years: the same co-movement that generates the debt response also

[10]The common-bust scenario is constructed by inverting $\mathbf{A}$ to find the shock vector that drives both prices down by 10 percent simultaneously.

widens the DK interval.[11] A terms-of-trade stress test therefore projects no debt accumulation in the episodes in which accumulation is greatest.

The gap between the composition-map prediction and the terms-of-trade prediction, computed over a grid of $(\Delta \ln P_x, \Delta \ln P_m)$, is largest along the common-price direction, the variation the ratio discards, and along export-led busts, where terms-of-trade projections understate debt accumulation. Historical Africa-20 price realizations cluster in these two regions. These estimates apply directly to any specified combination of export- and import-price changes in a debt sustainability assessment.

The pattern is not specific to Africa. In DJP's full 38-country panel, the reduced-form symmetry restriction is rejected at the 1 percent level from $h = 2$ onward, and the reduced-form import coefficient outside Africa is -20.6. Leave-one-country-out forecasts of four-year debt changes that use the two prices separately outperform a terms-of-trade-only forecast in 17 of 20 held-out economies.

4.2.4 Busts Cost More Than Booms Save

The structural estimates treat positive and negative shocks symmetrically by construction. I test whether this symmetry holds in the data: does an export-price boom reduce debt by as much as an export-price bust raises it? An asymmetry would matter for stress testing, since it means the fiscal cost of a price collapse exceeds the fiscal savings from a comparable boom.

Equation (13) imposes equal effects of export booms and busts. Splitting u_x into positive and negative parts relaxes this restriction (Table 4). I define boom savings as $-\hat{\beta}^{+}$ (debt decline per unit positive shock) and bust costs as $-\hat{\beta}^{-}$ (debt increase per unit negative shock). An export bust raises debt by 9.5 points at $h = 4$, while a boom of equal size lowers it by 2.5 points, significant only at the 10 percent level; I reject equality of the two at the 10 percent level. Dropping the 2008–09 and 2014–16 crisis years raises the bust cost to 12.3 points at $h = 4$ and leaves the boom saving at 2.5. The asymmetry is stronger for fiscal capacity: on debt-to-revenue, bust costs exceed boom savings at every horizon, and equality is rejected at $h = 1$ to $h = 3$. The debt-to-revenue scenarios are imprecise, however: an export bust raises the ratio by 49.7 points at $h = 4$ (95 percent interval $[1.3, 98.0]$), while the import-cost and common-bust responses, 19.8 and 6.7 points, have intervals that include zero.

The 2015 commodity collapse illustrates where the asymmetry matters most. In that year, fourteen country-years had both prices falling by more than 5 log points. The asymmetric map projected an average four-year rise in debt of 18.8 points; the symmetric map projected 11.6; the terms-of-trade regression projected 4.5; realized debt rose by 23.5 points. The 2008–09 bust is the contrasting case: realized debt rose by only 4.9 points, close to the terms-of-trade projection of 4.9 and well below the asymmetric map. That bust was short-lived, with a 2010–12 recovery compressing the four-year change within the measurement window. The asymmetry shows up in prolonged collapses, not brief ones.

Out-of-sample forecasts confirm the same pattern. These comparisons use coefficients estimated on the full sample, so they assess in-sample fit rather than out-of-sample forecast accuracy. The leave-one-country-out forecasts in Panel B are out of sample: using the two prices separately lowers the forecast error relative to the terms-of-trade-only model in 17 of 20 held-out economies, though by only 3.5 percent on average.

[11] The panel-mean common component was strongly negative in 2001 and 1998 (-2.060 and -1.682). Equatorial Guinea and Cameroon in 2015 show the same gap at the country level: common-component shocks of -4.015 and -3.723 against terms-of-trade signals of -0.703 and 0.197. The 2019 synchronized bust (20 of 20 countries) is not an artifact of a single narrative episode: DJP's narrative restrictions attach a 2019 negative export-price shock to only three Africa-20 economies.

Table 4: Structural export-price sign split: gross debt

h	Boom saves	Bust costs	p(boom)	p(bust)	$p(\beta^+ = \beta^-)$
0	2.27	2.94	0.260	0.129	0.850
	(1.96)	(1.85)			
1	1.96	6.35***	0.155	0.001	0.061
	(1.32)	(1.58)			
2	2.38	6.39***	0.139	0.007	0.209
	(1.54)	(2.11)			
3	1.72	7.74***	0.150	0.005	0.049
	(1.15)	(2.42)			
4	2.51*	9.51***	0.054	0.006	0.079
	(1.22)	(3.09)			

Observations: 441 at every h; $G = 20$ countries

Notes: Sign-split local projections of general-government gross debt on positive and negative parts of the DJP export-price shock, $u_x^+ = \max(u_x, 0)$ and $u_x^- = \min(u_x, 0)$, with coefficients $\hat{\beta}^+$ and $\hat{\beta}^-$, with linear u_m and the baseline structural controls, Africa-20, 1990–2019. Parentheses are country-clustered standard errors. $p(\beta^+ = \beta^-)$ tests equal-magnitude boom and bust effects. Units: percentage points of GDP per unit structural shock. *, **, *** denote significance at the 10, 5, and 1 percent levels, two-tailed, using $t(G-1)$ critical values.

4.2.5 Export-Price Shocks and Sovereign Restructuring

Debt accumulation and sovereign distress are related but distinct outcomes. A country can accumulate debt without entering restructuring, and restructuring can occur without large prior debt accumulation if creditors lose confidence or fiscal space erodes quickly. I therefore ask whether the same export- and import-price shocks that raise debt also predict sovereign restructuring events. I estimate the local projection

$$R_{i,[t,t+h]} \;=\; \alpha_i^{(h)} + \gamma_{u_x}^{(h)}\, u_{x,it} + \gamma_{u_m}^{(h)}\, u_{m,it} + \mathbf{w}_{it}'\, \boldsymbol{\delta}^{(h)} + \varepsilon_{it}^{(h)}, \tag{14}$$

where $R_{i,[t,t+h]}$ equals one if a restructuring year occurs in $[t, t+h]$ and zero otherwise, and $\mathbf{w}_{it}$ collects the same controls as equation (13). The coefficients $\gamma_{u_x}^{(h)}$ and $\gamma_{u_m}^{(h)}$ measure the change in the probability of a restructuring event per unit shock; they are distinct from the debt-accumulation coefficients $\beta_{u_x}^{(h)}$ and $\beta_{u_m}^{(h)}$ of equation (13). A positive structural export-price shock raises the probability of a restructuring event within one year at the 5 percent level under a linear probability model and at the 1 percent level under a fixed-effects conditional logit (Tables 5 and A6). Import-price shocks are insignificant at every horizon and do not predict restructuring events. The estimate is stable across sub-periods with no significant era interaction. This is the first result establishing the asymmetric role of the two price shocks in African sovereign distress, as distinct from debt accumulation.

The sign of the export-price coefficient is positive: export booms, not busts, precede restructuring years. This reflects the timing of African debt relief. Restructuring agreements in Africa (with Paris Club creditors, Chinese official lenders, and private bondholders) are typically concluded as commodity export revenue recovers, because creditors are willing to negotiate when debt service is again feasible. The Horn–Reinhart–Trebesch database records the agreement year, which coincides with the recovery rather than the price trough, so the positive coefficient on export booms is a timing feature of the data, not a contradiction of the debt response. Splitting the database into distress-onset years and agreement years confirms this: the export-price coefficient is negative at onset and positive at agreement, and the pooled positive sign reflects that agreement years outnumber onset years in the database.[12]

[12]Agreement years include Paris Club and Chinese deals, symbolic Chinese restructurings, and completions of private-creditor deals. Distress-onset years number sixteen over the sample, so the split is directional rather than

Table 5: Structural export-price shocks and sovereign debt restructuring

	Horizon h				
	$h=0$	$h=1$	$h=2$	$h=3$	$h=4$
Panel A: Restructuring year in $[t, t+h]$, none at $t-1$					
$\hat{\gamma}_{u_x}$	0.007	0.063**	0.043*	0.038*	0.012
	(0.021)	(0.026)	(0.021)	(0.019)	(0.020)
$\hat{\gamma}_{u_m}$	−0.011	0.003	−0.016	−0.015	−0.014
	(0.014)	(0.012)	(0.018)	(0.021)	(0.024)
$p(\hat{\gamma}_{u_x} + \hat{\gamma}_{u_m} = 0)$	0.875	0.010	0.213	0.411	0.942
N	437	437	437	437	437
Panel B: In default					
$\hat{\gamma}_{u_x}$	0.002	−0.002	−0.009	−0.010	−0.012
	(0.006)	(0.003)	(0.008)	(0.010)	(0.011)
$\hat{\gamma}_{u_m}$	0.006	0.005	0.000	−0.000	0.000
	(0.005)	(0.007)	(0.007)	(0.008)	(0.006)
N	471	471	471	471	471

Notes: Linear probability model (LPM) local projections on DJP structural export- and import-price shocks, twenty African economies, 1990–2019. Panel A: the outcome equals one if a restructuring year occurs between t and $t+h$, among country-years without one at $t-1$. A restructuring year is a Paris Club or Chinese agreement (including symbolic Chinese restructurings) or the start or end year of a private-creditor restructuring in the Horn–Reinhart–Trebesch database; every one of the twenty economies has at least one. Panel B: the outcome equals one during any ongoing default or restructuring episode. N is smaller in Panel A because it drops country-years with a restructuring year at $t-1$. Country fixed effects; standard errors clustered by country ($G=20$). *, **, *** denote significance at the 10, 5, and 1 percent levels, two-tailed, using $t(G-1)$ critical values.

4.3 Identification Beyond the DJP Rotation

The baseline results rest on DJP's identification strategy. Four exercises test whether the findings depend on that specific rotation: propagating uncertainty across the full posterior of admissible rotations, three shift-share designs that share no price source or trade weights with DJP, a WTI instrument for oil exporters, and raw episode counts. All four converge: the export-price coefficient is negative, the import-price coefficient is negative at medium-run horizons, and the symmetry restriction is rejected from $h=1$ onward.

First, I address the rotation uncertainty inherent in sign-restricted SVARs. Fry and Pagan (2011) show that element-wise posterior medians across admissible rotation matrices can produce a rotation belonging to no admissible model. To propagate this uncertainty through to the local projections, I re-estimate the Africa-20 bilateral VARs over 1990–2018—the period for which reliable African public debt data are available—imposing DJP's narrative sign restrictions, and draw 50,000 times over the full posterior of admissible Q rotation matrices. For each draw I reconstruct the country-specific structural shocks and run the local projection. At $h=2$ to 4, the export coefficient is negative in 99.6 to 100 percent of draws, the import coefficient in 98.3 to 99.6 percent, and symmetry is rejected at the 5 percent level in 94.2 to 97.9 percent. The median export coefficient across all 50,000 draws is -2.17 at $h=2$ and -3.14 at $h=4$, attenuated relative to the DJP baseline of -5.49 at $h=4$ as expected when averaging over all admissible rotations rather than the best-fitting one, but the sign and the symmetry rejection are consistent throughout. Shocks rebuilt from common-component prices or leave-Africa-out factors reproduce the sign and ranking of the export-price coefficient at $h=4$. Following Juvenal and Petrella (2024), who document that global financial conditions drive commodity prices, I residualize the published shocks on changes in the Chicago Board Options Exchange Volatility Index (VIX) and the U.S. real rate before the projection; this moves the $h=4$ export coefficient only from -4.55 to -4.52 on a common sample, and symmetry still fails at the 1 percent level.

precise.

Second, three shift-share instruments interact predetermined trade shares with world commodity prices (Goldsmith-Pinkham, Sorkin, and Swift, 2020; Borusyak, Hull, and Jaravel, 2022). The first uses CEPII BACI trade shares (Gaulier and Zignago, 2010) and World Bank Pink Sheet prices (World Bank, 2026a) and shares no input with DJP; the second uses DJP's 1985–89 trade weights; the third uses the IMF CToT weights (Gruss and Kebhaj, 2019). All have strong first stages (first-stage F-statistics reported in Table A2), the Anderson–Rubin test strongly rejects $\beta_x = \beta_m = 0$, and the first two reject symmetry at every horizon (Table A2). For the BACI/Pink Sheet design the rejection at $h = 2$ and $h = 4$ survives exposure-robust standard errors clustered by commodity (Table A3). The third gives a common-bust response of 8.2 points at $h = 4$, within 2.2 points of the structural estimate of 6 (Table A4). Because these instruments do not remove global demand, they reproduce the reduced-form pattern of the two price coefficients; they show that common declines raise debt.

Third, for the six oil exporters, I instrument $\Delta \ln P_x$ with an annual WTI series (Federal Reserve Bank of St. Louis, 2026) (country-clustered first-stage F of 210, with only six clusters; no DJP input). The instrumented export-price increase raises revenue, improves the primary balance, and lowers debt from impact through four years (Table A5). The fiscal pattern matches the structural one: revenue and the primary balance improve most on impact and at $h = 1$, then diminish while debt continues to decline.

Fourth, in the nine country-years in which both prices fell by more than 10 percent, the median rise in debt from $t - 1$ to $t + 4$ was 13.1 points of GDP with no controls.

The principal reduced-form findings also hold under restricted wild cluster bootstrap inference. Wild cluster bootstrap p-values are consistent with the asymptotic ones, and dropping any single country leaves the $h = 2$ import coefficient between -79.0 and -59.5. One specification overturns the symmetry rejection. Adding U.S. inflation, the ten-year real rate, the VIX, and a broad dollar index (controls that proxy for U.S. monetary conditions, which may independently affect African sovereign borrowing costs), together with domestic exchange-rate terms, halves the sample to 220 observations, all in 2009–2019, because the dollar index is available only from 2007 and enters with two lags. On that smaller sample, the baseline specification still rejects symmetry at the 5 percent level. The four global variables vary only by year, so on an eleven-year sample their current values and lags span year fixed effects exactly; the specification is therefore equivalent to adding year fixed effects, which absorb much of the usable common variation in commodity prices (Section 3.1). The symmetry rejection survives on this restricted sample; the failure to reject on the global-variable specification reflects absorbed identifying variation, not a fragility in the baseline.

4.4 Transmission

Knowing that export- and import-price shocks raise debt by different amounts is not enough: a risk analyst or policymaker also needs to know through which channels the debt accumulates. I decompose each debt response into a primary-balance contribution and a remainder that captures valuation, financing, and nominal-GDP effects, and trace the two price shocks through revenue, the GDP deflator, and the exchange rate.

Equation (6) decomposes each debt response into a primary-balance contribution and a remainder. A positive export-price shock raises government revenue, with the largest response at $h = 1$, while expenditure exhibits no consistent response and debt declines (Figure 4).

Corporate income tax receipts and the effective tax rate on capital income also respond to the shock (Bachas, Fisher-Post, Jensen, and Zucman, 2026), pointing to transmission through export-related capital income. High informality in African labor markets limits cross-sector adjustment: the effective tax rate on labor income is statistically indistinguishable from zero at every horizon, while capital income responds significantly, so the adjustment operates through sector-specific capital

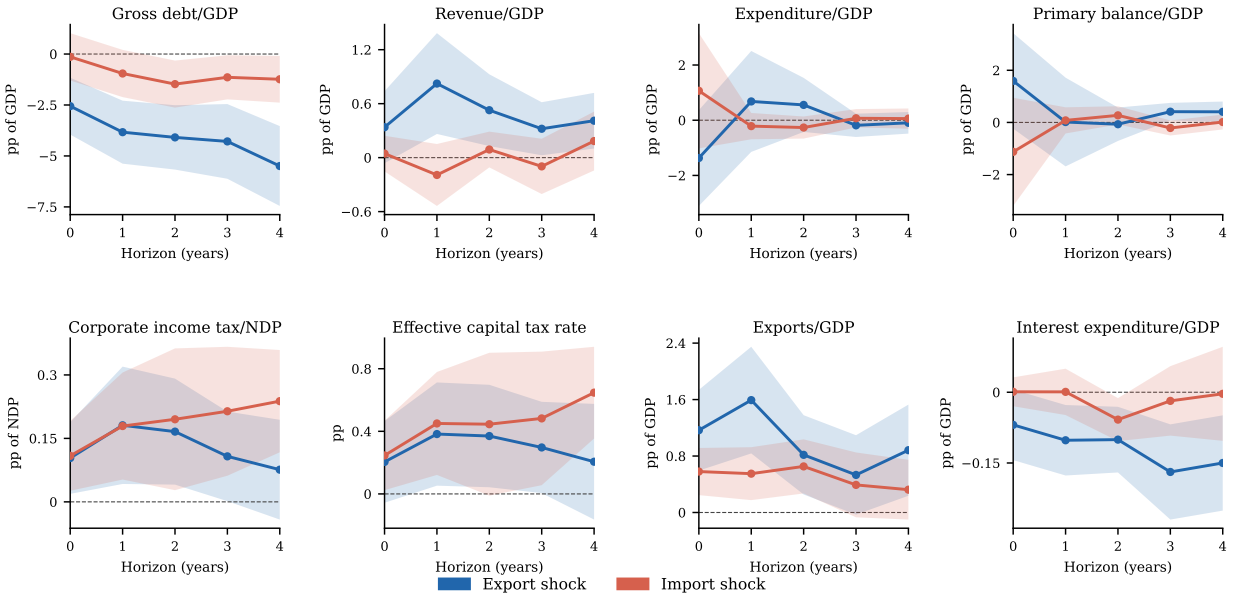


Figure 4: Fiscal transmission of DJP structural export-price shocks (blue) and import-price shocks (red), eleven outcomes, horizons $h = 0$ to $h = 4$. Units: percentage points per one-standard-deviation shock, of GDP or, for the four tax-revenue panels, of net domestic product (NDP); the effective capital tax rate is in percentage points. Darker and lighter shading are 68 and 90 percent pointwise confidence bands.

rents rather than wages.

Cumulative debt-to-revenue multipliers range from -5.5 at $h = 1$ to -8.4 at $h = 4$, and a joint test rejects the null of no debt response across the tested horizons. The primary balance accounts for 43 percent of the debt response at $h = 4$ and between 38 and 62 percent across horizons; these shares are approximate because the components are estimated on different samples and the decomposition does not implement the exact interest-weighted identity of equation (6). In the Jordà and Taylor (2025) stacked counterfactual, setting the estimated revenue response to zero attenuates the debt response from -5.4 to -4.5 at $h = 4$, both bootstrap means (the point estimate in Table 2 is -5.49; Figure A1). The revenue path is not jointly significant (stacked path test across $h = 0$–4), so setting it to zero is a modest intervention in the sense of Jordà and Taylor (2025); it does not measure that channel's share of the debt response.

An import-price increase worsens the primary balance yet reduces the debt ratio: the primary-balance contribution and the total debt response have opposite signs. Among the fiscal components, corporate income tax receipts rise, while goods-and-services taxes and interest expenditure show no clear response. Domestic issuance falls by 1.47 percentage points of GDP on impact, with no offsetting increase in external disbursements. The debt response is jointly significant across horizons; the revenue response is not.

At $h = 4$, the primary-balance contribution raises debt by 0.97 percentage points of GDP, while the remainder reduces it by 2.21 points. I test two candidate channels for the remainder. First, foreign-currency revaluation: interactions with predetermined foreign-currency debt shares are insignificant, and setting the structural dollar-rate response to zero does not attenuate the debt response. Second, the denominator: in the restricted sample excluding high-inflation episodes (absolute log deflator change above $\ln 1.4$ in any year from $t - 2$ to $t + h$; see footnote below), the GDP-deflator response rises from 0.006 log points on impact to 0.023 at $h = 4$ while the real output response is negligible, reflecting nominal-GDP expansion that reduces the debt ratio. In the full sample, which keeps the high-inflation episodes, the deflator response is indistinguishable from zero

at every horizon. The common-bust response at $h = 4$ in that sample divides approximately equally between numerator and denominator contributions of 2.97 and 3.07 points (Table A7; the sample and controls differ from Table 3, so the total is 6.03 rather than 5.96).[13] The denominator evidence and the null result of the valuation test together narrow the remainder to nominal-GDP pass-through and debt transactions not recorded in the primary balance; the available data do not separately identify these two components. In the sample excluding high-inflation episodes, the denominator channel accounts for the majority of the total import-price debt response at $h = 4$: the 0.023 log-point deflator response, applied to mean debt of 54.5 percent of GDP, implies approximately -1.25 percentage points from nominal-GDP growth alone, against a primary-balance cost of 0.97 points, more than reversing the fiscal effect. Across the two structural shocks, the export shock accounts for approximately 95 percent of their joint explanatory power for debt at $h = 4$.

4.5 Implications for Debt Stress Tests

The empirical results have direct implications for how sovereign debt risk is assessed in practice. The IMF–World Bank LIC-DSF is the primary framework through which commodity-price risk enters official debt sustainability assessments for low-income countries. I evaluate two specific features of that framework against the evidence: the reliance on single-leg export-price stress tests, and the Paragraph 71 netting provision that allows co-moving import prices to offset assessed export stress.

The LIC-DSF applies a tailored commodity-price stress test when commodities account for at least 50 percent of exports. The default is a one-standard-deviation decline in export prices. Paragraph 71 of the Guidance Note allows a net export figure for the commodity, which may be applied; where the country has large commodity imports subject to the same price shock, a customized re-run should be done (International Monetary Fund and World Bank, 2018). The 2024 Supplement leaves these provisions in effect and adds guidance only on climate change, domestic public debt, and debt restructuring (International Monetary Fund and World Bank, 2024). The September 2026 review leaves the commodity calibration unchanged (International Monetary Fund and World Bank, 2026).

The estimates identify two failures in the LIC-DSF stress design. First, sovereign risk frameworks and empirical models that compress external-price exposure into the terms-of-trade ratio miss the full common-bust debt accumulation: when both prices fall equally the ratio records no change, yet debt accumulates by 6 percentage points. Second, the Paragraph 71 import-price netting provision moves in the wrong direction in common-bust episodes: netting a co-moving import-price decline against the export stress reduces assessed exposure when the data show that assessed exposure should not fall. The same-commodity net export figure in Paragraph 71 adjusts exposure to a particular export-price shock; the present results do not challenge that adjustment within its intended scope. They show that simultaneous export- and import-price declines do not offset in the debt ratio, so netting cannot substitute for evaluating the common-bust scenario separately.

The structural import-price response is separately identified and negative at $h = 4$. The decomposition establishes where it operates: the primary-balance contribution and the total debt response have opposite signs, so the response arises outside the primary balance. The valuation tests provide no clear evidence of a foreign-currency revaluation channel. A government that raises tax

[13] The exclusion criterion is an absolute annual log change in the GDP deflator exceeding $\ln(1.4)$, corresponding to an annual deflator change above 40 percent, which flags recognized hyperinflation and large devaluation episodes; an observation is dropped when any year from $t-2$ to $t+h$ meets it. The criterion flags the Democratic Republic of Congo (1991–97 and 1999–2001), Sudan (1991–95, 2011, 2012, 2014, and 2018), Nigeria (1992–95), Madagascar (1994–95), Malawi (1995–96), Equatorial Guinea (1998 and 2000), Ghana (1995), Algeria (1991), and Chad, Côte d'Ivoire, and Gabon in 1994, the year of the CFA (Communauté Financière Africaine) franc devaluation. The cumulative CPI-based measure provides a complementary check.

rates, as Petrella, Juvenal, and Di Pace (2026) document in strong-institution settings, still faces this stock response operating through the GDP deflator and debt transactions not recorded in the primary balance. Because this stock response operates through the denominator rather than fiscal flows, it does not appear in standard flow-based fiscal monitoring and should be tracked separately in stress projections.

In the twelve LIC-DSF economies alone, the structural ranking holds and the stress-map estimates are of similar size (Tables A12 and A13).

The two channels do not operate through the same mechanism. For the export shock, the primary balance accounts for 38 to 62 percent of the debt response across horizons, with revenue losses as the leading component. For the import shock, the primary-balance contribution and the total debt response have opposite signs: fiscal flows worsen but debt falls, because nominal-GDP growth in the denominator more than offsets the fiscal cost. Standard flow-based fiscal monitoring captures the first mechanism and misses the second.

5 Conclusion

The terms-of-trade ratio discards common price movements that generate debt responses in this sample. When export and import prices fall by the same proportion, the ratio records no change and any risk assessment anchored to the ratio predicts no debt movement; yet debt rises by 6 percentage points of GDP over four years. In the nine country-years in which both prices fell by more than 10 percent, the median four-year debt increase was 13.1 points, with no controls or identification. The results establish two distinct failures. The first concerns the source of the price movement: export- and import-price shocks move debt in opposite directions, so compressing them into their ratio reverses the import-price sign and assigns the wrong magnitude to both. The second concerns common movements: equal declines in both prices leave the ratio unchanged while debt accumulates.

The two price responses transmit through different channels. On the export side, the fiscal channel operates through revenue, peaking one year after the shock; the primary balance accounts for less than half of the total debt response. On the import side, a remainder outside the primary balance, larger than the primary-balance contribution and of the opposite sign, dominates the debt response. The export shock also predicts a sovereign debt restructuring event within one year, significant at the 5 percent level under a linear probability model and the 1 percent level under a fixed-effects conditional logit: export booms, not busts, precede restructuring years, consistent with agreements concluded as export revenue recovers.

Analysts who stress the terms-of-trade ratio understate risk for export-led busts and miss common busts entirely. Analysts who stress export prices directly, as the LIC-DSF does, handle export-led busts correctly but still miss common busts when the Paragraph 71 netting provision offsets the export stress with a co-moving import-price decline. The September 2026 LIC-DSF review retains single-leg commodity shock tests and adds a gross financing needs threshold; the 2.97-point numerator contribution of a common bust (Table A7) could push projected GFN above the new threshold while a ratio-based baseline projects no accumulation (International Monetary Fund and World Bank, 2026).

The analysis has three limitations. First, the design uses identified surprise price shocks; anticipatory borrowing around expected price paths lies outside the identified window. Second, country-specific sovereign spreads are not separately identified, though the four-year debt accumulation is established through revenue, stock-flow adjustments, and the GDP denominator independently of the interest channel. Third, the analysis provides a diagnostic rather than a calibrated structural model; establishing that the ratio fails, and measuring by how much, is a prerequisite for quantitative

welfare analysis, not a substitute for it. A calibrated small-open-economy model with separate export and import prices and a government sector would quantify the symmetry failure and explain the revenue timing, the sign reversal of the import-price response, and the boom-bust asymmetry. Better currency-composition data on sovereign balance sheets would test whether the import-price remainder reflects foreign-currency revaluation or debt transactions not recorded in the primary balance. Higher-frequency sovereign spread data would allow direct estimation of the interest-rate channel. Extending the bilateral price series past 2019 would test whether the post-pandemic commodity cycles, which generated common price declines of different composition, preserve the asymmetry or reveal that the sign of the import-price response depends on the source of the global shock.

References

Alonso, Virginia, Clara Arroyo, Ozlem Aydin, Vybhavi Balasundharam, Hamid R. Davoodi, Gabriel Hegab, W. Raphael Lam, Anh Nguyen, Natalia Salazar Ferro, Galen Sher, Alexandra Solovyeva, and Nino Tchelishvili, 2025, "Fiscal Rules Dataset: 1985–2024," version 1.0 (Washington, DC: International Monetary Fund), September 25.

Alquist, Ron, Saroj Bhattarai, and Olivier Coibion, 2020, "Commodity-Price Comovement and Global Economic Activity," *Journal of Monetary Economics* 112: 41–56.

Antolín-Díaz, Juan, and Juan F. Rubio-Ramírez, 2018, "Narrative Sign Restrictions for SVARs," *American Economic Review* 108 (10): 2802–2829.

Arezki, Rabah, and Markus Brückner, 2012, "Commodity Windfalls, Democracy and External Debt," *Economic Journal* 122 (561): 848–866.

Asonuma, Tamon, and Christoph Trebesch, 2016, "Sovereign Debt Restructurings: Preemptive or Post-Default," *Journal of the European Economic Association* 14 (1): 175–214.

Bachas, Pierre, Matthew Fisher-Post, Anders Jensen, and Gabriel Zucman, 2026, "Globalization, Capital Taxation, and Development: Evidence from a Macro-Historical Database," *American Economic Journal: Applied Economics*, forthcoming.

Bai, Jushan, Sung Hoon Choi, and Yuan Liao, 2024, "Standard Errors for Panel Data Models with Unknown Clusters," *Journal of Econometrics* 240 (2): 105004.

Borusyak, Kirill, Peter Hull, and Xavier Jaravel, 2022, "Quasi-Experimental Shift-Share Research Designs," *Review of Economic Studies* 89 (1): 181–213.

Cameron, A. Colin, Jonah B. Gelbach, and Douglas L. Miller, 2008, "Bootstrap-Based Improvements for Inference with Clustered Errors," *Review of Economics and Statistics* 90 (3): 414–427.

Cohen, Naomi, Venance Riblier, and Diego Rodrigues, 2026, "Commodity Prices and Sovereign Risk," Working Paper, May.

Deaton, Angus, 1999, "Commodity Prices and Growth in Africa," *Journal of Economic Perspectives* 13 (3): 23–40.

Deaton, Angus, and Ron Miller, 1993, "Commodity Prices and Macroeconomic Management in Africa," Institute for Policy Reform Working Paper IPR71 (Princeton: Research Program in Development Studies, Princeton University).

Delle Chiaie, Simona, Laurent Ferrara, and Domenico Giannone, 2022, "Common Factors of Commodity Prices," *Journal of Applied Econometrics* 37 (3): 461–476.

Di Pace, Federico, Luciana Juvenal, and Ivan Petrella, 2025, "Terms-of-Trade Shocks Are Not All Alike," *American Economic Journal: Macroeconomics* 17 (2): 24–64.

Driscoll, John C., and Aart C. Kraay, 1998, "Consistent Covariance Matrix Estimation with Spatially Dependent Panel Data," *Review of Economics and Statistics* 80 (4): 549–560.

Ehrhart, Hélène, and Samuel Guérineau, 2013, "Commodity Price Volatility and Tax Revenue: Evidence from Developing Countries," Banque de France Working Paper No. 423 (Paris: Banque de France), February.

Federal Reserve Bank of St. Louis, 2026, "Spot Crude Oil Price: West Texas Intermediate (WTI) [WTISPLC]," retrieved from FRED, Federal Reserve Bank of St. Louis, `https://fred.stlouisfed.org/series/WTISPLC`.

Fry, Renée, and Adrian Pagan, 2011, "Sign Restrictions in Structural Vector Autoregressions: A Critical Review," *Journal of Economic Literature* 49 (4): 938–960.

Gaulier, Guillaume, and Soledad Zignago, 2010, "BACI: International Trade Database at the Product-Level. The 1994–2007 Version," CEPII Working Paper No. 2010-23 (Paris: CEPII), October.

Giacomini, Raffaella, Toru Kitagawa, and Matthew Read, 2022, "Narrative Restrictions and Proxies,"

Journal of Business & Economic Statistics 40 (4): 1415–1425.

Gilbert, Christopher L., 2011, "Grains Price Pass-Through, 2005–09," in Adam Prakash (ed.), *Safeguarding Food Security in Volatile Global Markets* (Rome: Food and Agriculture Organization of the United Nations), pp. 122–143.

Goldsmith-Pinkham, Paul, Isaac Sorkin, and Henry Swift, 2020, "Bartik Instruments: What, When, Why, and How," *American Economic Review* 110 (8): 2586–2624.

Gruss, Bertrand, and Suhaib Kebhaj, 2019, "Commodity Terms of Trade: A New Database," IMF Working Paper WP/19/21 (Washington, DC: International Monetary Fund).

Harberger, Arnold C., 1950, "Currency Depreciation, Income, and the Balance of Trade," *Journal of Political Economy* 58 (1): 47–60.

Horn, Sebastian, Carmen M. Reinhart, and Christoph Trebesch, 2022, "Hidden Defaults," *AEA Papers and Proceedings* 112: 531–535.

International Monetary Fund, 2009, *Export and Import Price Index Manual: Theory and Practice* (Washington, DC: International Monetary Fund), prepared jointly with the International Labour Office, OECD, Eurostat, United Nations Economic Commission for Europe, and World Bank.

International Monetary Fund, 2025, *World Economic Outlook* database, April 2025 vintage (Washington, DC: International Monetary Fund).

International Monetary Fund and World Bank, 2018, *Guidance Note on the Bank-Fund Debt Sustainability Framework for Low Income Countries*, IMF Policy Paper (Washington, DC: International Monetary Fund), December 2017, released 2018.

International Monetary Fund and World Bank, 2024, *Supplement to the 2018 Guidance Note on the Bank-Fund Debt Sustainability Framework for Low Income Countries* (Washington, DC: International Monetary Fund), July.

International Monetary Fund and World Bank, 2026, *Review of the Bank-Fund Debt Sustainability Framework for Low-Income Countries—Proposed Reforms*, IMF Policy Paper No. PP/26/034 (Washington, DC: International Monetary Fund), September, staff report completed 31 July 2026 and considered by the Executive Board on 9 September 2026.

Jordà, Òscar, 2005, "Estimation and Inference of Impulse Responses by Local Projections," *American Economic Review* 95 (1): 161–182.

Jordà, Òscar, and Alan M. Taylor, 2025, "Local Projections," *Journal of Economic Literature* 63 (1): 59–110.

Juvenal, Luciana, and Ivan Petrella, 2024, "Unveiling the Dance of Commodity Prices and the Global Financial Cycle," *Journal of International Economics* 150: 103913.

Laursen, Svend, and Lloyd A. Metzler, 1950, "Flexible Exchange Rates and the Theory of Employment," *Review of Economics and Statistics* 32 (4): 281–299.

Levy, Antoine, Luca Antonio Ricci, and Alejandro Mariano Werner, 2020, "The Sources of Fiscal Fluctuations," IMF Working Paper No. WP/20/220 (Washington: International Monetary Fund).

MacKinnon, James G., Morten Ø. Nielsen, and Matthew D. Webb, 2023, "Cluster-Robust Inference: A Guide to Empirical Practice," *Journal of Econometrics* 232 (2): 272–299.

Manger, Mark, David Mihalyi, Ugo Panizza, Niccolò Rescia, Christoph Trebesch, and Ka Lok Wong, 2025, "Africa's Domestic Debt Boom: Evidence from the African Debt Database," CEPR Discussion Paper No. 20747 (London: Centre for Economic Policy Research).

Medina, Juan Pablo, and Claudio Soto, 2016, "Commodity Prices and Fiscal Policy in a Commodity Exporting Economy," *Economic Modelling* 59 (C): 335–351.

Mei, Ziwei, Liugang Sheng, and Zhentao Shi, 2026, "Nickell Bias in Panel Local Projection: Financial Crises Are Worse Than You Think," *Journal of International Economics* 160: 104210.

Mendoza, Enrique G., 1995, "The Terms of Trade, the Real Exchange Rate, and Economic Fluctuations," *International Economic Review* 36 (1): 101–137.

Minot, Nicholas, 2011, "Transmission of World Food Price Changes to Markets in Sub-Saharan Africa," IFPRI Discussion Paper No. 1059 (Washington, DC: International Food Policy Research Institute).

Montiel Olea, José Luis, and Mikkel Plagborg-Møller, 2021, "Local Projection Inference Is Simpler and More Robust Than You Think," *Econometrica* 89 (4): 1789–1823.

Montiel Olea, José Luis, Mikkel Plagborg-Møller, Eric Qian, and Christian K. Wolf, 2026, "Double Robustness of Local Projections and Some Unpleasant VARithmetic," *Econometrica* 94 (4): 1313–1343.

Muhanji, Stella, and Kalu Ojah, 2011, "External Shocks and Persistence of External Debt in Open Vulnerable Economies: The Case of Africa," *Economic Modelling* 28 (4): 1615–1628.

Müller, Karsten, Chenzi Xu, Mohamed Lehbib, and Ziliang Chen, 2025, "The Global Macro Database: A New International Macroeconomic Dataset," NBER Working Paper No. 33714 (Cambridge, MA: National Bureau of Economic Research).

Petrella, Ivan, Luciana Juvenal, and Federico Di Pace, 2026, "Commodity Prices and Fiscal (Pro)Cyclicality," Collegio Carlo Alberto Working Paper No. 757 (Turin: Collegio Carlo Alberto), April; earlier version IDB Working Paper No. IDB-WP-1739 (Washington, DC: Inter-American Development Bank), October 2025.

Piger, Jeremy, and Thomas Stockwell, 2023, "Differences from Differencing: Should Local Projections with Observed Shocks be Estimated in Levels or Differences?" Preprint, SSRN. `http://dx.doi.org/10.2139/ssrn.4530799`.

Pieschacón, Anamaria, 2012, "The Value of Fiscal Discipline for Oil-Exporting Countries," *Journal of Monetary Economics* 59 (3): 250–268.

Plagborg-Møller, Mikkel, and Christian K. Wolf, 2021, "Local Projections and VARs Estimate the Same Impulse Responses," *Econometrica* 89 (2): 955–980.

Prebisch, Raúl, 1950, *The Economic Development of Latin America and Its Principal Problems* (Lake Success, NY: United Nations Department of Economic Affairs, Economic Commission for Latin America), E/CN.12/89/Rev.1.

Schmitt-Grohé, Stephanie, and Martín Uribe, 2018, "How Important Are Terms-of-Trade Shocks?" *International Economic Review* 59 (1): 85–111.

Spatafora, Nikola, and Issouf Samaké, 2012, "Commodity Price Shocks and Fiscal Outcomes," IMF Working Paper No. 12/112 (Washington, DC: International Monetary Fund).

Ugarte-Ruiz, Alfonso, 2026, "Panel Local Projections without Fixed-Effects: A Cumulative-Difference OLS Estimator," SSRN Working Paper 7020198; also BBVA Research Working Paper No. 26/09 (Bilbao: BBVA Research), June.

World Bank, 2026a, *Commodity Price Data (The Pink Sheet)*, monthly prices (Washington, DC: World Bank), `https://www.worldbank.org/en/research/commodity-markets`, accessed October 2026.

World Bank, 2026b, *International Debt Statistics* database (Washington, DC: World Bank), `https://www.worldbank.org/en/programs/debt-statistics/ids`, accessed October 2026.

A Data and Sample

This section documents the sample, data sources, and variable construction underlying the main results. Table A1 lists the twenty economies, their main exports, income groups, and LIC-DSF debt sustainability status.

Table A1: Country coverage

Country	ISO3	Primary export	Income group	LIC-DSF	Debt rating	DSA pub.
Algeria	DZA	Oil and gas	Upper-middle	No	—	—
Burkina Faso	BFA	Gold, cotton	Low	Yes	Moderate	02/26
Cameroon	CMR	Oil, cocoa	Lower-middle	Yes	High	03/25
Chad	TCD	Oil	Low	Yes	High	10/25
Côte d'Ivoire	CIV	Cocoa, oil	Lower-middle	Yes	Moderate	08/25
Dem. Rep. Congo	COD	Minerals	Low	Yes	Moderate	09/25
Egypt	EGY	Oil, gas	Lower-middle	No	—	—
Equatorial Guinea	GNQ	Oil	Upper-middle	No	—	—
Gabon	GAB	Oil	Upper-middle	No	—	—
Ghana	GHA	Gold, cocoa, oil	Lower-middle	Yes	High	12/25
Kenya	KEN	Tea, horticulture	Lower-middle	Yes	High	10/24
Madagascar	MDG	Vanilla, minerals	Low	Yes	Moderate	02/25
Malawi	MWI	Tobacco, tea	Low	Yes	In distress	10/25
Mauritius	MUS	Services, sugar	Upper-middle	No	—	—
Morocco	MAR	Phosphates	Lower-middle	No	—	—
Niger	NER	Uranium, gold	Low	Yes	High	03/26
Nigeria	NGA	Oil	Lower-middle	No	—	—
Senegal	SEN	Gold, fish	Lower-middle	Yes	Moderate	06/23
South Africa	ZAF	Gold, minerals	Upper-middle	No	—	—
Sudan	SDN	Oil, gold	Low	Yes	In distress	06/21

Notes: Every country has non-missing bilateral price indices in all 30 years, 1990–2019. Income group is the World Bank 2023 classification. LIC-DSF: IMF–World Bank Low-Income Country Debt Sustainability Framework. Debt rating is the overall rating from the World Bank Debt Toolkit DSA country table, as of end-April 2026; "—" indicates economies not covered by the framework. Sample selection follows bilateral price-data availability, not DSA status. Uranium, Niger's main export, is not among the commodities DJP price, so commodities are 25 percent of Niger's export index in the 1990s and 73 percent in the 2010s.

B Identification Details

This section details the identification results in Section 4.3: robustness of the debt response across posterior draws of the Africa-20 bilateral VAR, a sign check on the structural shocks for a known price episode, the three shift-share designs, the WTI instrument for oil exporters, raw common-bust descriptives, the sovereign-restructuring regression, and the persistence and placebo checks. Across all variants below, the export-price coefficient is negative, the symmetry restriction is rejected, and the common-bust estimate remains positive and significant unless otherwise noted.

Shock sign check. DJP's sign restrictions imply that a negative structural export-price shock should accompany the 2014–16 oil-price collapse. The structural export shock is negative in each of the six oil exporters during that episode: the panel mean is $\bar{u}_x = -1.71$ in 2015 and -1.39 in 2016, confirming that the rotation produces economically sensible shocks.

Posterior draws. The baseline in the main text uses DJP's published posterior-median shocks. Fry and Pagan (2011) show that element-wise medians across posterior draws can produce a rotation that belongs to no admissible model, so the projection is also estimated on each of the 50,000 stored draws of the Africa-20 bilateral VAR. The export coefficient is negative in at least 96.9 percent of draws at every horizon from $h = 0$ to $h = 4$, and the import coefficient in at least 94.6 percent

at every horizon from $h = 1$ to $h = 4$. The symmetry restriction is rejected at the 5 percent level in 94.2 to 97.9 percent of draws at $h = 2$ to 4, in 76.6 percent at $h = 1$, and in 16.4 percent at $h = 0$. Across draws the export coefficient is smaller in magnitude than the baseline (median -3.14 against -5.49 at $h = 4$), so the draws support the sign and the symmetry rejection rather than the magnitude.

Shift-share instruments. Each design interacts predetermined export and import shares in commodity c with world price changes, $z_{x,it} = \sum_c s_x[i,c]\,\Delta \ln P^{\text{world}}_{c,t}$ and $z_{m,it} = \sum_c s_m[i,c]\,\Delta \ln P^{\text{world}}_{c,t}$, and (z_x, z_m) instrument $(\Delta \ln P_x, \Delta \ln P_m)$ by two-stage least squares after removing the effects of the baseline controls and country effects. The DJP trade-weight design uses the 1985–89 average of DJP's own country-by-commodity weights and commodity-price panel, but not their bilateral indices or rotation. The BACI/Pink Sheet design uses the 1995–99 average of each country's trade composition from CEPII's BACI database (Gaulier and Zignago, 2010) and World Bank Pink Sheet prices (World Bank, 2026a), without sharing any input with DJP.[14] The export-instrument series from the two designs correlate at 0.85 and the import-instrument series at 0.86. At $h = 4$ the Sanderson–Windmeijer conditional F statistics are 24.3 and 22.2 for the first and 17.3 and 22.7 for the second, the Kleibergen–Paap statistics are 10.3 and 8.8, and the Anderson–Rubin test strongly rejects $\beta_x = \beta_m = 0$ in both. Neither design applies DJP's global-demand orthogonalization; accordingly, both reproduce the reduced-form pattern of the two price coefficients instead of isolating the structural effects.[15] Adding third and fourth lags of debt attenuates the $h = 4$ common-bust estimate by 15 to 20 percent in all three designs, but each estimate remains significant.

Table A2: Full-panel shift-share (Bartik) LP-IV, $h = 4$, Africa-20 (1990–2019)

	DJP weight-tensor (1985–89 shares)		BACI/Pink Sheet (1995–99 shares)
Own-equation F, $\Delta \ln P_x$	242.7		61.3
Own-equation F, $\Delta \ln P_m$	176.6		124.6
Corr. with DJP u_x / u_m	0.57 / 0.49		0.56 / 0.49
Corr. with the other Bartik instrument (z_x/z_m)		0.85 / 0.86	
SW conditional F, $\Delta \ln P_x$ / $\Delta \ln P_m$	24.3 / 22.2		17.3 / 22.7
Kleibergen–Paap rk Wald F	10.3		8.8
Anderson–Rubin p-value (joint)	<0.001		<0.001
2SLS $\hat{\beta}_x^{(4)}$ (SE)	−19.61 (10.87)		−18.66 (10.51)
2SLS $\hat{\beta}_m^{(4)}$ (SE)	−84.76 (40.90)		−87.64 (48.95)
p-value, $\beta_x^{(4)} + \beta_m^{(4)} = 0$	0.0016		0.0090
N	441		441

Notes: Both columns instrument the bilateral $\Delta \ln P_x / \Delta \ln P_m$ pair with a predetermined-share Bartik instrument (z_x, z_m); Column 1 uses DJP's own weight tensor (1985–89 shares), Column 2 uses CEPII BACI/World Bank Pink Sheet shares (1995–99), sharing no input with DJP's package. Country-clustered standard errors in parentheses; units are log-point price movements, not one-standard-deviation structural shocks.

Exposure-robust inference and Rotemberg weights. With standard errors clustered by commodity, the exposure-robust standard errors of Borusyak, Hull, and Jaravel (2022), or by year, the symmetry restriction is still rejected at the 5 percent level at $h = 2$ and $h = 4$. Oil carries a Rotemberg weight of 0.80 in the export instrument (Goldsmith-Pinkham, Sorkin, and Swift, 2020),

[14] The 1995–99 base period is predetermined for observations from 2000 onward but overlaps the 1990–99 part of the estimation window. Results are qualitatively unchanged when the sample is restricted to 2000–2019.

[15] In a pre-trend check, the import-price shift-share instrument predicts debt changes over $t-3$ to $t-1$, but that association disappears with year fixed effects or outside 2000–2010; on the baseline DJP structural shocks, the import shock is null at every lead ($k = 1, 2, 3$). The pre-trend is a property of the shift-share instrument, not the DJP shocks.

but its concentration in six exporters does not drive the structural estimates: the symmetry rejection is strongest outside the oil group and the export-price coefficient is negative in non-oil economies at every horizon (Table A11 and Table A3).

Table A3: BACI/Pink Sheet shift-share design: exposure-robust inference and Rotemberg weights

Panel A: LP-IV estimates and symmetry test under alternative clustering

	$h = 2$			$h = 4$		
	Country	Commodity	Year	Country	Commodity	Year
$\hat{\beta}_x$		−16.94			−18.66	
standard error	(6.68)	(14.13)	(10.80)	(10.51)	(11.55)	(12.47)
$\hat{\beta}_m$		−55.54			−87.64	
standard error	(20.88)	(44.30)	(38.85)	(48.95)	(30.29)	(38.24)
$p(H_0 : \beta_x + \beta_m = 0)$	<0.001	0.027	0.040	0.017	<0.001	0.004
Clusters	20	26	27	20	26	27

Panel B: Largest Rotemberg weights

	Export instrument				Import instrument			
	Commodity	α_k	$\hat{\beta}_k$, $h = 2$	$\hat{\beta}_k$, $h = 4$	Commodity	α_k	$\hat{\beta}_k$, $h = 2$	$\hat{\beta}_k$, $h = 4$
1	oil	0.797	−23.5	−17.1	oil	0.432	−51.0	−78.8
2	cotton	0.112	−8.2	−78.8	rice	0.171	−25.6	0.5
3	cocoa	0.076	−0.2	−51.9	wheat	0.167	62.2	−139.2
4	natural gas	0.070	−11.5	−14.9	maize	0.090	44.7	−27.0
5	phosphate	−0.037	−183.3	−181.0	sugar	0.082	−448.9	−205.8
Negative weights (number, share of $\sum \lvert\alpha_k\rvert$)		11 of 20, 0.067				0 of 7, 0.000		
Effective number of shocks		5.13				4.43		

Notes: BACI/Pink Sheet shift-share LP-IV of the debt ratio (column 2 of Table A2). Panel A: standard errors clustered by country, by commodity across all years (the exposure-robust standard errors of Borusyak, Hull, and Jaravel (2022), computed on the per-commodity instrument components), and by year. Symmetry p-values use $t(G-1)$ critical values, with G the number of clusters, so the country-clustered values differ from the normal-based p-values in Table A2. Panel B: Rotemberg weights of Goldsmith-Pinkham, Sorkin, and Swift (2020) for each leg, with the other leg's instrument held as published; $\hat{\beta}_k$ is the just-identified estimate using commodity k's component alone, and $\sum_k \alpha_k \hat{\beta}_k$ equals the 2SLS coefficient exactly. Weights do not depend on h because the sample is the same at $h = 2$ and $h = 4$. The effective number of shocks is the inverse Herfindahl index of the commodity weights. Oil enters the export and the import instruments as separate components.

IMF CToT shift-share instrument. The third design uses fixed weights from the IMF CToT database (Table A4), with country effects only (adding year effects collapses the import-price first stage). Its export-instrument series correlates at 0.87 with the DJP trade-weight instrument and its import-instrument series at 0.83, despite sharing no inputs. As a check on generalizability, rebuilding the BACI/Pink Sheet instruments on a global panel of 51 commodity-exporting countries with Global Macro Database debt (Müller, Xu, Lehbib, and Chen, 2025) reproduces the same sign pattern and strongly rejects symmetry at $h = 2$ ($N = 1{,}318$; $G = 51$); outside the Africa-20 the pattern holds in the subsample of 38 countries.

WTI instrument for oil exporters. For the six oil exporters, the instrument is the annual log change in the WTI spot price (Federal Reserve Bank of St. Louis, 2026), common to the six in each year; year effects would absorb it, so the specification has country effects only. Leave-one-country-out estimates preserve the negative sign, as does dropping 2008–09 or 2014–16. Residualizing WTI on lagged VIX, U.S. real rates, and a broad dollar index before forming the instrument also preserves the sign ($N = 72$; Table A5). The design is a sign check on the oil subsample; the two columns of Table A5 use different units, so the comparison concerns sign, not magnitude.

Table A4: A third public shift-share instrument: IMF commodity terms of trade, Africa-20 (1990–2018)

	$h = 0$	$h = 4$	$h = 8$
2SLS $\hat{\beta}_x$ (SE)	−12.48 (7.84)	−19.91** (8.58)	−10.33 (11.55)
2SLS $\hat{\beta}_m$ (SE)	−6.09 (20.87)	−62.37* (30.99)	−117.35*** (32.86)
Common bust, both −10% (SE)	1.86 (1.52)	8.23*** (2.71)	12.77*** (2.63)
First-stage and validation, country FE (primary spec)			
Own-equation F, $\Delta \ln P_x$ / $\Delta \ln P_m$		65.6 / 107.9	
Same, with year FE added (not usable, see notes)		51.6 / 5.9	
DJP u_x / u_m on own CToT instrument, F		47.9 / 40.4	
Corr. with DJP weight-tensor instrument (z_x/z_m)		0.87 / 0.83	
Min-eigenvalue (Cragg–Donald) statistic, $h = 4$		85.4	
N, $h = 4$		421	

Notes: Fixed pre-sample trade-weighted IMF Commodity Terms of Trade indices (Gruss and Kebhaj, 2019) instrument the bilateral $\Delta \ln P_x / \Delta \ln P_m$ pair; same shift-share construction, FE-2SLS-after-FWL estimator, and country-clustered inference as Table A2. Sample is 1990–2018 (CToT data through 2018). Shares no weights or price source with the DJP or BACI/Pink Sheet instruments; correlates 0.87/0.83 with the DJP weight-tensor instrument. Year FE collapse the import-leg first stage (F from 107.9 to 5.9); reported specification uses country FE only. *, **, *** denote significance at the 10, 5, and 1 percent levels, two-tailed, using $t(G-1)$ critical values.

Table A5: Oil exporters: WTI LP-IV versus DJP structural export leg

Horizon	DJP structural (oil-6) (pp of GDP per one-s.d. u_x)	WTI LP-IV (oil-6) (pp of GDP / log-point $\Delta \ln P_x$)
$h = 0$	−2.80 (2.30)	−18.87*** (6.49)
$h = 1$	−4.68 (3.06)	−35.97*** (7.85)
$h = 2$	−4.46 (2.48)	−26.39*** (4.65)
$h = 3$	−5.48 (2.74)	−32.69*** (9.46)
$h = 4$	−6.25** (2.18)	−39.92*** (11.13)
First-stage F ($z \to \Delta \ln P_x$)	—	210.1
Corr($z, \Delta \ln P_x$)	—	0.91
N / G	136–148 / 6	148 / 6

Notes: Oil exporters only: Algeria, Gabon, Equatorial Guinea, Nigeria, Sudan, and Chad. Long-difference local projections with country fixed effects; year effects are excluded because they would absorb the WTI instrument. Column 1 uses DJP structural shocks jointly (import shock controlled); Column 2 instruments $\Delta \ln P_x$ with the annual log change in WTI (FRED WTISPLC). Columns use different units; the comparison is directional. Parentheses are country-clustered standard errors. *, **, *** denote significance at the 10, 5, and 1 percent levels.

Sovereign restructuring. Section 4.2.5 reports the restructuring results. Table A6 reports the conditional logit and the era-interaction tests at every horizon.

Table A6: Structural export-price shocks and sovereign restructuring: conditional logit and era interaction

	Horizon h (years)				
	0	1	2	3	4
Panel A: Fixed-effects conditional logit, restructuring outcome					
$\hat{\gamma}_{u_x}$	0.047	0.457***	0.296**	0.284*	0.100
	(0.172)	(0.155)	(0.149)	(0.152)	(0.157)
$\hat{\gamma}_{u_m}$	−0.112	−0.017	−0.155	−0.138	−0.124
	(0.160)	(0.141)	(0.141)	(0.147)	(0.153)
N	337	337	337	337	337
Countries with outcome variation	16	16	16	16	16
Panel B: Linear probability model with a post-2007 interaction					
p, joint test of the 2007-onward interactions	0.095	0.907	0.709	0.436	0.227

Notes: Outcome and controls as in Panel A of Table 5. Panel A: fixed-effects conditional logit coefficients (log-odds per one-standard-deviation shock) with observed-information standard errors in parentheses; countries without variation in the outcome drop out, so the sample has fewer countries than the linear probability model. Stars use normal critical values. Panel B: the linear probability model of Table 5 with both shocks interacted with an indicator for 2007 and later years; the p-value is from the country-clustered $F(2, 19)$ test that both interaction coefficients are zero. *, **, *** denote significance at the 10, 5, and 1 percent levels.

Persistence and the placebo test. Gross debt is persistent, with a within-country AR(2) coefficient sum near 0.83; the median half-life is 10.3 years for P_x and 14.3 for P_m, against 4.6 for the terms of trade, so the ratio can mask the persistence of either price series. A reverse-causality placebo replaces the outcome with a lagged long difference of debt centered k years before the price change. The placebo is insignificant at $k = 1$. At $k = 2$ the import coefficient is significant at the 1 percent level but disappears with year fixed effects, so the pre-trend reflects common time variation; the same placebo on net lending, a fiscal-flow variable that should not share the stock trend, is insignificant at both leads, confirming the pre-trend is a stock feature. Year fixed effects are not used in the baseline because they absorb the cross-sectional variation in common-price episodes that the common-bust identification exploits.

C Additional Results

C.1 Transmission

This section details the transmission results in Section 4.4: counterfactual revenue and exchange-rate shutdowns, the reduced-form budget–remainder decomposition, the numerator–denominator split with an alternative price-level series, the foreign-currency valuation test, and the domestic issuance response.

Revenue and exchange-rate counterfactuals. Figure A1 shows the structural debt response to an export-price shock with and without the estimated government-revenue path (left panel) and the reduced-form debt response to an import-price change with and without the estimated exchange-rate path (right panel). Section 4.4 discusses the revenue shutdown.

The reduced-form exchange-rate shutdown in the right panel moves the bootstrap-mean $h = 2$ import coefficient from about -68 to about -7, which would suggest a strong exchange-rate valuation mechanism. That conclusion does not survive structural identification. In the Jordà–Taylor stacked-path counterfactual (10,000 bootstrap draws; as in the original implementation, each draw retains only the countries sampled exactly once, about seven of the twenty) for the separately identified import-price shock u_m, shutting down the estimated exchange-rate path leaves the debt response nearly unchanged at short horizons and larger at medium horizons (bootstrap means of -1.37 and -1.87 at $h = 4$, against the point estimate of -1.24 in Table 2). The exchange-rate path is only marginally jointly significant for the structural shock (stacked path test across $h = 0$–4), so the reduced-form attenuation reflects the endogeneity of the exchange rate to the import-price shock, not a genuine valuation channel. The reduced-form shutdown is therefore not interpreted as evidence of exchange-rate valuation.

Numerator–denominator split. The debt-ratio response is decomposed into a numerator contribution (changes in the debt stock) and a denominator contribution (changes in nominal GDP), following equation (6) with Global Macro Database nominal GDP, real GDP, and the GDP deflator (Müller, Xu, Lehbib, and Chen, 2025). Replacing the deflator with WEO CPI inflation cumulated into a price-level series leaves the denominator contribution unchanged at every horizon, confirming it reflects price pass-through rather than a deflator construction choice. Table A7 splits the debt-ratio response exactly into numerator and denominator terms. The export-price response is numerator-dominated at every horizon; a substantial share of the response is left unattributed, so the split is descriptive accounting.

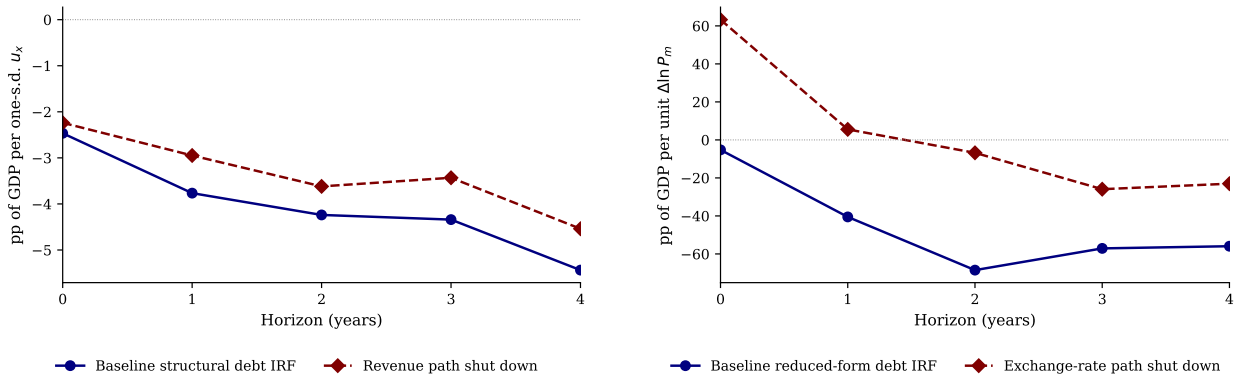


Figure A1: *Left*: structural gross-debt response to a one-standard-deviation export-price shock, with (solid) and without (dashed) the government-revenue path, from the Jordà–Taylor stacked-path counterfactual. Units: percentage points of GDP per one-standard-deviation shock. *Right*: reduced-form gross-debt responses to an import-price change, with and without the exchange-rate path. Units: percentage points of GDP per unit $\Delta \ln P$.

Table A7: Exact numerator/denominator split: export leg and common bust

h	Actual	Numerator	Budget	Stock-net	Denominator
Export leg, 10 percent export-price decline					
0	2.06	1.58	0.79	0.78	0.49
1	3.68	2.58	1.21	1.37	1.10
2	3.53	2.82	−0.08	2.91	0.71
3	3.98	3.22	1.51	1.71	0.76
4	4.25	3.53	1.71	1.82	0.72
Common bust, 10 percent decline in both prices					
0	0.86	−0.42	−2.53	2.11	1.28
1	4.37	1.87	−1.77	3.64	2.50
2	6.52	3.32	2.79	0.53	3.20
3	5.29	2.40	0.76	1.64	2.89
4	6.03	2.97	0.97	2.00	3.07

Notes: Percentage points of GDP. Same panel, sample, and high-inflation exclusion as Table **??**. Each regression adds two lags of the change in log nominal GDP and of the primary balance to the controls of Table 2. The export column scales the export-shock coefficient by the estimated impact response $\hat{a}_{xx}$, and the common-bust column uses the estimated impact matrix $\hat{\mathbf{A}}$ (Section 3.2). Numerator + Denominator = Actual and Budget + Stock-net = Numerator by construction. Numerator is the nominal-debt-stock change; Budget is minus the cumulative primary balance; Stock-net bundles currency revaluation, debt relief, stock-flow adjustments, and arrears (not a clean valuation estimate). The exclusion and the added controls give a smaller sample than Table 3, so Actual differs from the scenarios there; the impact matrix is the same.

Valuation. Section 4.4 notes that the import-price debt response falls largely outside the primary balance, with the GDP deflator as the leading candidate for the residual channel. A foreign-currency revaluation channel is tested by interacting the structural import shock with each country's predetermined foreign-currency debt share. The interaction coefficient is small and statistically indistinguishable from zero at every horizon (−0.046 to 0.008 across $h = 0$–4), providing no evidence of a revaluation channel. With only nineteen countries, the test has limited power, but the import response is larger in the eight CFA franc countries, where no member can devalue unilaterally against the peg (the one joint devaluation, in January 1994, enters the estimation sample only for Gabon and Equatorial Guinea), consistent with the absence of a revaluation effect.

Financing. A further candidate for the residual channel is gross financing: the import-price shock may reduce domestic debt issuance, contributing to the debt-ratio decline through the numerator. A one-standard-deviation import shock lowers gross domestic issuance by 1.47 percentage points of GDP on impact, with no offsetting rise in external disbursements, while the export shock has no significant effect (Figure A2). The domestic issuance panel covers $G = 13$ countries at $h = 0$–2 and

$G = 8$ at $h = 4$.

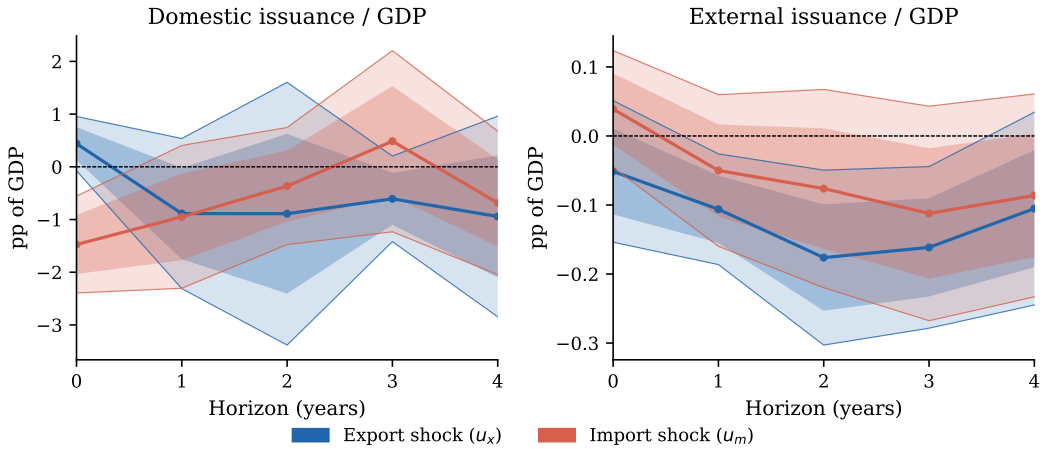


Figure A2: Structural shocks and gross financing. Blue: export-price shock; red: import-price shock. Left: African Debt Database domestic issuance. Right: World Bank IDS public and publicly guaranteed disbursements ($G = 19$). Shaded areas as in Figure 1. Units: percentage points of GDP per one-standard-deviation shock.

C.2 Tax Composition

Table A8 reports structural local projections for four tax outcomes from the Bachas et al. (2024) effective-tax-rate dataset: total tax revenue, corporate income tax, the effective capital tax rate, and the spread between the effective capital and labor tax rates. The ETR spread directly measures whether a shock tilts the tax mix toward capital or labor income. Both shocks widen the spread significantly, with the effect on capital taxation rising and persistent through $h = 4$ for the import shock. This suggests that price shocks transiently reverse the secular compression of capital taxation that Bachas et al. document for the same countries over the long run.

Table A8: Tax-composition responses to export- and import-price shocks

Outcome	Leg	$h=0$	$h=1$	$h=2$	$h=3$	$h=4$	N
Total tax/NDP	Export	0.08	0.19	0.31**	0.17	0.01	510
		(0.11)	(0.14)	(0.13)	(0.11)	(0.19)	
	Import	0.01	0.16	0.11	0.19	0.45***	510
		(0.13)	(0.14)	(0.21)	(0.21)	(0.14)	
Corporate income tax/NDP	Export	0.10*	0.18**	0.17**	0.11	0.08	510
		(0.05)	(0.08)	(0.08)	(0.06)	(0.07)	
	Import	0.11**	0.18**	0.20*	0.21**	0.24***	510
		(0.05)	(0.08)	(0.10)	(0.09)	(0.07)	
Effective capital tax rate	Export	0.20	0.38*	0.37*	0.30	0.21	510
		(0.16)	(0.20)	(0.20)	(0.18)	(0.22)	
	Import	0.25*	0.45**	0.45	0.48*	0.65***	510
		(0.13)	(0.20)	(0.28)	(0.26)	(0.18)	
ETR spread (capital minus labor)	Export	0.29*	0.48**	0.40**	0.30*	0.22	510
		(0.14)	(0.18)	(0.18)	(0.17)	(0.22)	
	Import	0.27*	0.42**	0.49*	0.49*	0.52***	510
		(0.13)	(0.20)	(0.27)	(0.24)	(0.17)	

Notes: Structural local projections using DJP export- and import-price shocks (u_x, u_m). Long-difference specification with country fixed effects and baseline structural controls (Table 2). Tax outcomes are from Bachas, Brockmeyer, Continentino, and Moretti (2024); units are percentage points of NDP per one standard deviation of the shock, except the ETR spread (percentage points). The ETR spread is the effective capital tax rate minus the effective labor tax rate. Standard errors clustered at the country level in parentheses ($G = 20$). *, **, *** denote significance at the 10, 5, and 1 percent levels using $t(G-1)$ critical values.

C.3 Inference and Robustness

This section documents the robustness of the symmetry rejection and the individual price coefficients to alternative inference procedures, fixed-effect choices, lag orders, sample restrictions, heterogeneity splits, and influential observations.

Wild cluster bootstrap p-values (Rademacher weights, 9,999 draws) are consistent with the asymptotic ones throughout: the export-price bootstrap p rises from 0.035 to 0.082 at $h = 0$ but stays at or below 0.02 from $h = 2$ onward, and the import-price bootstrap p is 0.001 or below at $h = 2$–6. Replacing country fixed effects with year fixed effects moves the $h = 2$ import coefficient from -66.7 to 16.5 (standard error 34.0) and the export coefficient from -9.7 to 2.9, confirming that identification requires within-country variation. With the global-macro controls of Section 4.3, the import coefficient is 28.0 and the export coefficient 4.7, with symmetry not rejected; the sample restriction alone moves the import coefficient from -66.7 to -39.4 and leaves the rejection intact. Restricting outcomes to $t + h \leq 2019$ to exclude pandemic-era debt movements leaves the symmetry rejection intact at every horizon from $h = 0$ to $h = 5$. The $h = 4$ scenario estimates in this sample are 4.8 points for the export bust, -3.6 for the import-cost shock, and 5.7 points for the common bust (95 percent delta-method interval $[1.7, 9.6]$; Table A9), so the main estimate does not depend on post-2019 outcomes.

Cross-sectional dependence and few clusters. Table A10 re-tests the structural symmetry restriction with Driscoll–Kraay standard errors and with a restricted wild cluster bootstrap using Webb weights. Both reject at the 5 percent level from $h = 0$, through $h = 6$ and $h = 7$ respectively. Driscoll–Kraay standard errors approximately double the uncertainty around the common bust: its interval at $h = 4$ is $[-0.6, 12.5]$, insignificant at the 10 percent level, against $[2.7, 9.2]$ with country clustering. In the twelve LIC-DSF economies, Webb weights and exact enumeration of all

Table A9: The $h = 4$ scenarios under alternative samples

	Export bust	Import cost	Common bust	95% interval	N
Baseline (Table 3)	5.71	−3.40	5.96	[2.73, 9.20]	441
Dropping extreme debt values	5.73	−3.58	6.15	[2.85, 9.46]	438
Outcomes up to 2019 only	4.77	−3.58	5.66	[1.67, 9.64]	361

Notes: Debt responses at $h = 4$, in percentage points of GDP, to a 10 percent export-price decline, a 10 percent import-price increase, and a 10 percent decline in both prices, computed as in Table 3. "Dropping extreme debt values" removes Equatorial Guinea through 1995 and Sudan through 1993 from both the price and the debt regressions. "Outcomes up to 2019 only" keeps observations with $t + 4 \leq 2019$ in the debt regression. Intervals are delta-method country-clustered intervals with the impact matrix held fixed; N is the number of observations in the debt regression.

2^{12} Rademacher sign vectors both reject structural symmetry at the 5 percent level at $h = 0$ and at the 1 percent level at $h = 4$.

Table A10: Inference under cross-sectional dependence and few clusters

Panel A: Symmetry test, $H_0 : \beta_{u_x} + \beta_{u_m} = 0$ (Table 2 specification), p-values

	Horizon h (years)								
	0	1	2	3	4	5	6	7	8
Country-clustered, $t(19)$	0.009	<0.001	<0.001	<0.001	<0.001	0.005	<0.001	0.012	0.094
Driscoll–Kraay, $t(T-1)$	0.028	0.003	0.001	<0.001	<0.001	0.016	0.040	0.120	0.410
Wild cluster bootstrap, Webb	0.007	<0.001	<0.001	<0.001	<0.001	<0.001	0.001	0.010	0.101

Panel B: Common bust, 10 percent decline in both prices (Table 3 specification)

	Horizon h (years)				
	0	1	2	3	4
Debt response (pp of GDP)	1.67	4.39	5.83	5.10	5.96
Country-clustered 95% interval	[−1.65, 5.00]	[0.88, 7.90]	[2.36, 9.30]	[1.91, 8.28]	[2.73, 9.20]
Driscoll–Kraay 95% interval	[−2.46, 5.80]	[−0.82, 9.60]	[0.35, 11.31]	[−0.49, 10.68]	[−0.58, 12.50]
Driscoll–Kraay p-value	0.413	0.095	0.038	0.072	0.072

Panel C: Symmetry test in the twelve LIC-DSF economies ($G = 12$), p-values

	$h = 0$			$h = 4$		
	Clustered	Webb	Rademacher	Clustered	Webb	Rademacher
Structural (u_x, u_m)	0.025	0.030	0.038	0.003	0.001	0.002
Reduced form ($\Delta \ln P_x$, $\Delta \ln P_m$)	0.232	0.259	0.258	0.005	0.001	<0.001

Notes: Panel A re-tests the symmetry restriction of Table 2 on the same regressions. Driscoll–Kraay standard errors use a Bartlett kernel with bandwidth $\lfloor 4(T/100)^{2/9} \rfloor$ and $t(T-1)$ critical values, with T the number of years in each regression. The wild cluster bootstrap imposes the null, multiplies the restricted residuals by one Webb six-point weight per country, and uses 9,999 draws. Panel B reports the common-bust scenario of Table 3 with its country-clustered interval (normal critical value, as in that table) and with Driscoll–Kraay standard errors ($t(T-1)$ critical value); the impact matrix is held fixed. Panel C re-tests the restriction in the LIC-DSF-12 subsample of Table A12: "Rademacher" enumerates all $2^{12} = 4{,}096$ sign vectors, so it is exact; "Webb" uses 9,999 draws.

Heterogeneity and influence. When the sample is split by commodity export structure (Table A11), the symmetry rejection is strongest for non-oil exporters (significant at the 1 percent level at $h = 2$ and $h = 4$), weaker for agricultural and mineral exporters (significant at the 5 percent level), and not significant for oil exporters or diversified economies. Dropping one country at a time (Section 4.3) leaves the $h = 2$ export coefficient between −11.4 and −7.5. Country-by-country structural projections give a negative export coefficient in 14 of 20 countries at $h = 2$ and 12 of 20 at $h = 4$, and a negative import coefficient in 16 and 13. The individual estimates are imprecise, with 16 to 27 observations and 14 controls, but the sign distribution is inconsistent with a few

countries driving the pooled result.[16] A Hodrick–Prescott-filter commodity-boom interaction finds no evidence that boom or bust regimes alter pass-through. The result also holds with Global Macro Database public debt ($\hat{\beta}_m^{(2)} = -66.8$) and on a post-MDRI subsample from 2006 ($\hat{\beta}_m^{(2)} = -36.9$), both significant at the 1 percent level.

Table A11: Commodity-Architecture Heterogeneity: Gross Debt Responses

	$h = 0$	$h = 2$	$h = 4$
Panel A. Import-price coefficient $\hat{\beta}_m^{(h)}$ by subsample			
Oil exporters ($N = 6$)	62.79	−136.92	−86.54
Non-oil ($N = 14$)	−12.14	−59.33	−52.15
Agricultural ($N = 5$)	14.55	−21.48	−31.80
Mineral ($N = 5$)	−27.68	−106.72	−89.34
Diversified ($N = 4$)	−22.67	−18.09	−19.08
Panel B. Symmetry test p-value ($H_0 : \beta_x + \beta_m = 0$)			
Oil exporters ($N = 6$)	0.684	0.137	0.095
Non-oil ($N = 14$)	0.090	0.001	0.001
Agricultural ($N = 5$)	0.922	0.030	0.047
Mineral ($N = 5$)	0.174	0.051	0.039
Diversified ($N = 4$)	0.105	0.271	0.101

Notes: Reduced-form gross-debt local projections by commodity architecture for African economies, 1990–2019. Long-difference local projections with country fixed effects and the baseline controls. Units: percentage points of GDP per unit $\Delta \ln P$. Subsamples follow Table A1. Oil samples are underpowered ($T \approx 27$ per country).

Table A12: Gross-debt local projections: full sample versus LIC-DSF subsample

	$h = 0$			$h = 4$		
Panel A: Structural shocks (u_x, u_m)						
	$\hat{\beta}_{u_x}$	$\hat{\beta}_{u_m}$	p(sym)	$\hat{\beta}_{u_x}$	$\hat{\beta}_{u_m}$	p(sym)
Full Africa-20 ($G = 20$, $N = 441$)	−2.56***	−0.14	0.009	−5.49***	−1.24*	<0.001
	(0.85)	(0.71)		(1.19)	(0.70)	
LIC-DSF-12 ($G = 12$, $N = 255$)	−2.77**	−0.23	0.025	−5.92**	−1.14	0.003
	(1.01)	(0.91)		(1.99)	(1.10)	
Panel B: Reduced form ($\Delta \ln P_x$, $\Delta \ln P_m$)						
	$\hat{\beta}_x$	$\hat{\beta}_m$	p(sym)	$\hat{\beta}_x$	$\hat{\beta}_m$	p(sym)
Full Africa-20 ($G = 20$, $N = 441$)	−12.57**	−2.51	0.368	−26.25***	−55.72***	<0.001
	(5.72)	(19.63)		(9.04)	(17.76)	
LIC-DSF-12 ($G = 12$, $N = 255$)	−8.55	−15.61	0.232	−26.91	−56.53*	0.005
	(7.46)	(23.33)		(20.23)	(30.77)	

Notes: Reduced-form and DJP local projections of gross debt for the Africa-20 and the twelve LIC-DSF economies in the sample, 1990–2019. Long-difference local projections with country fixed effects, as in Tables 2 and 1. Units: percentage points of GDP per unit $\Delta \ln P$ or per one-standard-deviation shock. Parentheses are country-clustered standard errors. LIC-DSF-12 is BFA, CMR, TCD, CIV, COD, GHA, KEN, MDG, MWI, NER, SEN, SDN. With $G = 12$, inference is looser at long horizons; the structural export ranking survives. *, **, *** denote significance at the 10, 5, and 1 percent levels.

[16]The three largest deviations in the export-price coefficient (Malawi, Burkina Faso, Sudan) each contain structural breaks that a short series cannot absorb. Sudan's coefficient is driven by the 2011 secession of South Sudan, which removed most of its oil revenue base in a single year.

Table A13: Composition stress map: Africa-20 versus LIC-DSF-12 ($h = 4$)

	Composition map (pp of GDP)			ToT-only (pp of GDP)		
Sample	Export bust	Import cost	Common bust	Export bust	Import cost	Common bust
Africa-20	5.7	−3.4	6.0	3.9	3.9	0.0
LIC-DSF-12	7.3	−3.1	7.5	4.8	4.8	0.0

Notes: Same structural stress-map construction as Figure 3, re-estimated on the twelve LIC-DSF surveillance economies (BFA, CMR, TCD, CIV, COD, GHA, KEN, MDG, MWI, NER, SEN, SDN). With $G = 12$ inference is looser; the qualitative ToT-compression failure survives.