Commodity Exposure and the Severity of Sudden Stops

Fernando Letelier
Rutgers University

September 2026

Abstract

How does the composition of tradable income shape financial fragility during sudden stops? Using the Bianchi and Mendoza (2020) systemic sudden stop chronology, I show that economies entering an episode with greater commodity dependence undergo larger absolute adjustment: moving from the 25th to the 75th percentile of pre-crisis exposure is associated with a 2.93 percentage point larger current account reversal one year after onset, alongside larger consumption and equity price losses. I then extend the Fisherian collateral model of Bianchi (2011) by allowing commodity exposure to change the sensitivity of tradable income to a common world price. Holding the inherited balance sheet and shock fixed, moving across the empirically mapped exposure interquartile range raises the current account reversal by 4.88 percentage points. When exposure is anticipated, households borrow less ex ante: on a fixed set of crisis dates, the severity differential peaks at intermediate exposure and falls by 46 percent by $\alpha = 0.40$. A borrowing rule counterfactual shows that this endogenous adjustment substantially attenuates the exposure gradient. Greater exposure also raises the marginal external cost of leverage at a common balance sheet. Financial fragility therefore depends not only on leverage, but also on the composition of the income backing it and on endogenous balance sheet adjustment.

1 Introduction

How does the composition of tradable income shape financial fragility during sudden stops? Standard accounts emphasize how much an economy owes, the terms on which it borrows, and the financial frictions that constrain adjustment. This paper emphasizes a complementary margin: two economies with similar financial structures can respond very differently to the same adverse external price movement when the income backing their liabilities has a different composition.

Commodity exposure provides a natural setting in which to study this idea. Commodity exporters are disproportionately emerging and developing economies, but exposure varies substantially within development groups. I begin with the broad sudden stop chronology, systemic and non-systemic episodes, to document the unconditional severity pattern with the widest episode coverage. Among developed economies, the average current account reversal is 5.78 percent of GDP for commodity exporters and 2.70 percent for non-commodity exporters; among developing economies, the corresponding values are 5.02 and 3.19 percent. The same within group ordering appears for GDP, consumption, investment, and the real exchange rate, and it is preserved when the event paths are summarized by medians rather than means (Appendix Table A2). Development status therefore does not monotonically order absolute severity in the raw data: developed commodity exporters have larger current account reversals than developing non-commodity exporters, 5.78 versus 3.19 percent of GDP, and the same cross group ordering holds for the median event paths, 4.22 versus 2.56 percent. The econometric analysis then narrows the event class to systemic sudden stops, which coincide with broader external financial stress, and replaces the static exporter indicator with episode specific pre-crisis exposure. In that sample, moving from the 25th to the 75th percentile of commodity dependence is associated with a 2.93 percentage point larger current account reversal one year after onset and a 1.89 percentage point larger reversal at two years. Consumption and equity prices also fall significantly more at higher exposure, while investment and the real exchange rate move in the same direction with less uniform precision across horizons.

The empirical analysis focuses on the intensive margin of systemic sudden stops: whether economies that enter the same class of external financing event with greater pre-crisis commodity dependence undergo larger absolute adjustment. The baseline allows severity to vary with pre-crisis real GDP per capita and with the realized country specific commodity export price movement, and includes country and year fixed effects. The current account gradient is stable when I add pre-crisis private credit, capital account openness, the exchange rate regime, global real rate tightening, a clean comparison sample, and overall trade intensity.

Commodity exposure therefore provides a distinct structural margin for organizing crisis severity, rather than simply relabeling development or the scale of international trade.

The mechanism is simple. Let a share α of tradable income be linked to an exogenous world commodity price. At the normalized price, economies with different α have the same tradable income. After the same adverse price movement, however, the more exposed economy suffers the larger income loss. If borrowing capacity depends on current income and on the endogenous relative price of nontradables, lower tradable absorption depresses that price, reduces collateral values, and forces further deleveraging. Exposure changes the shock hitting the balance sheet, not the financial friction itself.

I formalize this idea in a minimal extension of Bianchi (2011). I preserve the benchmark preferences, collateral parameter, world interest rate, and joint domestic endowment process, and introduce only a world commodity price and the exposure parameter

$$y_t^T(\alpha) = z_t\big[(1-\alpha) + \alpha p_t^C\big].$$

This structure makes the central comparison transparent: at $p_t^C = 1$, changing α leaves tradable income unchanged; away from that price, it changes the sensitivity of income to the same commodity price realization.

The first quantitative experiment holds the inherited balance sheet, domestic state, pre-crisis commodity price, and adverse commodity price innovation fixed. I map the empirical exposure interquartile range into $\alpha_L = 0.30$ and $\alpha_H = 0.44$ using pre-crisis national accounts value added. Under the same one standard deviation commodity price decline, the high exposure economy experiences a 4.88 percentage point larger current account reversal, 2.60 percentage points more forced deleveraging relative to common pre crisis income, a 4.28 percentage point larger aggregate consumption contraction, and an 8.74 percentage point larger real depreciation. The direct difference in tradable income losses is only 1.41 percentage points, while the endogenous consumption, relative price, collateral, and current account responses are several times larger. This gap between the initial income loss and the equilibrium adjustment is the quantitative signature of the Fisherian amplification mechanism.

The second quantitative experiment lets households adjust their balance sheets before crises occur. In the full stochastic equilibrium, greater exposure induces households to borrow less and changes which dates become sudden stops. This creates a tension between two forces: exposure makes a given balance sheet more fragile, while anticipated exposure induces households to choose a safer balance sheet. Pairwise common event comparisons show that the current account severity differential continues to rise with exposure, reaching 2.47 percentage points at $\alpha = 0.40$. When instead I hold the crisis set fixed across the numerically reliable exposure grid, the differential is hump shaped: it peaks at 4.56 percentage points at $\alpha = 0.25$ and declines to 2.46 percentage points at $\alpha = 0.40$, a 46 percent decline from its peak. Thus, endogenous balance sheet adjustment eventually offsets part, but not all, of the additional conditional fragility generated by greater exposure.

To quantify the role of that adjustment more directly, I conduct a borrowing rule counterfactual at the empirically mapped exposures $\alpha_L = 0.30$ and $\alpha_H = 0.44$. The high exposure economy is assigned the low exposure borrowing rule wherever feasible while retaining its own income process, collateral constraint, and viability limit. Because the model concentrates Fisherian deleveraging at onset whereas the empirical current account differential peaks one year later, I compare the average response over the onset and first post-onset years. The current account exposure differential is essentially zero in equilibrium and rises to 2.63 percentage points under the counterfactual, compared with 1.89 percentage points in the data over the same two year window. Across alternative event definitions, the equilibrium differential remains near zero while the counterfactual differential remains economically large. The small equilibrium gradient therefore reflects endogenous self insurance rather than a weak structural exposure mechanism.

The same distinction matters for macroprudential policy. At a common leverage position, the marginal pecuniary externality of debt is about 6.4 percent larger in the empirically mapped high exposure economy. Strikingly, the local Fisherian multiplier $1/D$ falls from 5.43 to 4.80 across the two post-shock states. Greater exposure therefore does not raise the externality by strengthening the local collateral feedback. It pushes the economy deeper into the crisis region: the larger consumption loss raises post-crisis marginal utility enough to dominate the weaker local multiplier. Once debt is chosen endogenously, average policy wedges need not rise with exposure because private self insurance changes the states in which leverage is carried into crises.

Contributions. The paper makes three contributions. First, it identifies the composition of tradable income as a structural determinant of conditional financial fragility. The empirical evidence documents a systematic exposure, severity ordering during systemic sudden stops, while the matched state model isolates the same leverage mechanism. Second, it shows that endogenous private adjustment is quantitatively central: the same exposure that increases fragility at a common balance sheet also induces precautionary deleveraging, producing a hump shaped severity response on a fixed set of crisis dates. A borrowing rule decomposition shows directly that removing the differential precautionary response generates much larger crisis severity differences. Third, it separates crisis depth from the local collateral multiplier in the policy analysis. Exposure can raise the marginal social cost of leverage even when the local multiplier falls, because it changes how far the economy is pushed into the high marginal utility crisis region. Together, these results imply that vulnerability depends jointly on leverage and on the composition of the income backing that leverage.

Related literature. The paper relates first to work on emerging market business cycles and global financial conditions. Neumeyer and Perri (2005) and Uribe and Yue (2006) emphasize world interest rates, country spreads, and external financial conditions, while Aguiar and Gopinath (2007), García-Cicco et al. (2010), and Chang and Fernández (2013) study trend shocks and financial frictions. I complement this literature by isolating a structural exposure margin that varies within development groups and affects the severity of external adjustment even when the financial friction is unchanged.

A second literature studies terms of trade and commodity price shocks in small open economies. Mendoza (1995), Fernández et al. (2017), and Schmitt-Grohé and Uribe (2018) study their role in business cycle fluctuations, while Drechsel and Tenreyro (2018) and Fornero and Kirchner (2018) focus on commodity exporting economies. Recent work brings commodity shocks closer to financial crises: Gazzani et al. (2025) emphasize asymmetric financial effects of negative commodity price shocks, and Faris and Granados (2026) study terms of trade and collateral capacity in an environment with self fulfilling sudden stops. My question is different: rather than asking whether commodity price shocks trigger crises, I ask how a predetermined exposure margin changes crisis severity, and then hold the shock fixed in the model to identify the amplification mechanism.

Finally, the paper builds on the sudden stop and Fisherian amplification literature. Calvo (1998), Calvo et al. (2004), Edwards (2004), Cavallo and Frankel (2008), and Forbes and Warnock (2012) study the incidence and consequences of abrupt external adjustment. Mendoza (2010) and Bianchi (2011) show how occasionally binding collateral constraints generate nonlinear crisis dynamics and pecuniary externalities. I use the Bianchi (2011) environment because it provides exactly the two objects needed here: a transparent collateral price amplification mechanism and a planner benchmark for the social cost of leverage. The model remains intentionally parsimonious so that the role of commodity exposure can be isolated without changing the underlying financial friction.

The remainder of the paper is organized as follows. Section 2 presents the motivating evidence. Section 3 estimates the continuous commodity exposure gradient in systemic sudden stops. Section 4 develops the Fisherian framework. Section 5 describes the calibration and exposure mapping. Sections 6 and 7 present the controlled and stochastic quantitative exercises. Section 8 develops the macroprudential implications, and Section 9 concludes.

2 Motivating Evidence

The paper begins with a simple descriptive comparison that motivates the continuous analysis. I classify economies by development status and by a binary commodity exporter indicator following the World Bank's *Global Economic Prospects* (January 2022). The descriptive exercise uses the broad sudden stop chronology, including both systemic and non-systemic episodes, because its purpose is to summarize how crisis severity varies with export structure across abrupt external adjustment episodes. The main econometric exercise then restricts attention to the systemic subset and uses exact, episode specific commodity dependence measured in the year before onset. The two samples therefore serve different roles by design: broad coverage for the descriptive fact, and a tighter external financial stress event class for the conditional exposure gradient.

2.1 Sudden Stops and the Commodity Exporter Severity Premium

I take the country risk set and event chronology from Bianchi and Mendoza (2020). Their candidate event filter identifies unusually large annual increases in the current account to GDP ratio relative to each country's own historical distribution, and the systemic filter additionally requires elevated market wide financial stress, measured with the EMBI for emerging economies and the VIX for advanced economies. Table 1 uses the broad candidate event chronology. Section 3 uses only the systemic subset, so the regression asks how pre-crisis exposure orders severity within episodes that occur under a common class of external financial stress rather than pooling all large current account adjustments.

Table 1 reports peak to trough changes computed from the mean event path in each group over the five year window $t-2,\dots,t+2$, using 135 broad sudden stop episodes. Commodity exporters experience more severe average adjustment within both development groups. Among developed economies, the current account reversal is 5.78 percent of GDP for commodity exporters and 2.70 percent for non-commodity exporters; among developing economies, the corresponding values are 5.02 and 3.19 percent. GDP, consumption, investment, and the real exchange rate display the same within group ordering. The cross group comparison is also informative: developed commodity exporters experience larger average current account reversals than developing non commodity exporters, 5.78 versus 3.19 percent of GDP, so development status alone does not monotonically order absolute crisis severity. Appendix Table A2 shows the same comparison in median event paths, 4.22 versus 2.56 percent, and preserves the commodity exporter ordering for every outcome shown in Table 1. The developed commodity exporter cell contains nine episodes, so the binary comparison serves as descriptive motivation rather than the paper's main source of inference. The continuous systemic event analysis below then asks a sharper question with a different estimand: whether episode specific pre-crisis commodity dependence orders severity within systemic crises.

Table 1: Peak to Trough Changes from Mean Event Paths by Commodity Exporter and Development Status

Commodity exporter	Developing	Episodes	GDP	Consumption	Investment	RER	CA/GDP
0	0	45	-2.08	-3.18	-11.73	-1.86	2.70
1	0	9	-3.94	-5.65	-19.76	-6.42	5.78
0	1	40	-3.25	-3.35	-12.58	-13.10	3.19
1	1	41	-4.36	-5.20	-25.23	-19.87	5.02

Notes: The table uses the 135 episode broad sudden stop chronology, including systemic and non-systemic episodes, after the sample reconciliation described in Appendix A. For each group and each relative year $t-2,\ldots,t+2$, I first average the available episode observations and then compute the peak to trough movement of that mean event path. GDP, consumption, and investment are reported in percent. The RER is normalized within episode at $t-1$ before aggregation; negative values denote real depreciations. CA/GDP is reported in percentage points, with positive values denoting current account reversals. "Episodes" is the number of sudden stop episodes in the classification cell. Missing observations are handled outcome by outcome and relative year by relative year, so an episode with a missing RER observation, for example, can still contribute to the GDP or CA/GDP path. Appendix Table A2 reports the corresponding median path statistics. The main local projections use only the systemic subset and a continuous pre-crisis exposure measure.

The descriptive comparison is intentionally simple. Its role is to establish a broad cross episode fact, not to reproduce the local projection coefficient. Section 3 then asks whether the same ordering survives under a more disciplined design: systemic events only, continuous exposure measured immediately before onset, and conditioning on pre-crisis income and the realized commodity price environment. Appendix Table A2 shows that the descriptive ordering is unchanged when the group paths are summarized by medians rather than means.

3 Commodity Exposure and Sudden Stop Severity

The main empirical exercise replaces the static binary classification with continuous, episode specific pre-crisis exposure. I take the systemic sudden stop dates from Bianchi and Mendoza (2020) as given and ask whether the subsequent macroeconomic adjustment is larger in economies whose merchandise exports were more commodity intensive immediately before onset. The estimand is therefore an exposure gradient in the intensive margin of systemic crisis severity.

3.1 Econometric Framework

Local projection outcome. I estimate cumulative responses using the local projection method of Jordà (2005). For each outcome y_{it} and horizon $h \in \{-3, -2, -1, 0, 1, 2, 3\}$, I define

$$\Delta^h y_{it} \equiv y_{i,t+h} - y_{i,t-1}, \tag{1}$$

and normalize $h = -1$ to zero. The use of $t-1$ as the common reference point makes the pre- and post-onset paths directly interpretable as cumulative movements relative to the last pre-crisis year.

Baseline specification and estimand. Let $CD_{i,t-1}$ denote commodity exports as a share of merchandise exports in the calendar year before t, and let $G_{i,t-1} \equiv \log GDPpc_{i,t-1}$. I estimate

$$\begin{aligned}\Delta^h y_{it} = \alpha_i + \gamma_t + \beta_h SS_{it} + a_h CD_{i,t-1} + b_h G_{i,t-1} + \rho_h \Delta \log x_{it} \\ + \delta_h \left(SS_{it} \times CD_{i,t-1}\right) + \eta_h \left(SS_{it} \times G_{i,t-1}\right) + \psi_h \left(SS_{it} \times \Delta \log x_{it}\right) + \varepsilon_{i,t+h}.\end{aligned} \tag{2}$$

The regression uses the full country-year panel, not only crisis years. Country fixed effects α_i absorb time invariant differences across economies, and year fixed effects γ_t absorb shocks common to all countries in a given year. The indicator SS_{it} equals one only in systemic sudden stop onset years. The interaction $SS_{it} \times CD_{i,t-1}$ is therefore the key heterogeneous response term.

I include exact $t-1$ real GDP per capita and its sudden stop interaction because commodity dependence is correlated with development and income. Allowing sudden stop severity to vary with pre-crisis income prevents the commodity coefficient from simply reproducing an advanced versus emerging severity difference. Commodity dependence and log GDP per capita are standardized for estimation using their distribution among systemic episodes, but all reported effects are converted back into economically meaningful changes in the raw commodity export share.

The headline statistic is the implied difference between economies at the 75th and 25th percentiles of pre-crisis commodity dependence,

$$\Delta_h^{75-25} \equiv \hat{\delta}_h \left(CD_{75} - CD_{25}\right). \tag{3}$$

In the baseline systemic sample,

$$CD_{25} = 22.5\%, \qquad CD_{75} = 72.9\%,$$

so the interquartile comparison changes the commodity share by 50.4 percentage points. I also report the local effect of a 10 percentage point increase in commodity dependence.

Commodity price conditioning. The variable x_{it} is the IMF country specific commodity export price index constructed from fixed historical commodity export basket weights. I include both $\Delta \log x_{it}$ and $SS_{it} \times \Delta \log x_{it}$, so the baseline exposure gradient is estimated after conditioning on the realized commodity price environment. The baseline coefficient therefore summarizes the average exposure, severity ordering across systemic events at a given realized price environment; Appendix B allows that gradient to vary flexibly with the price realization.

Appendix B shows that the one year current account gradient remains similar when the price terms are removed and when the exposure gradient is allowed to vary flexibly with the realized price movement. I therefore retain the parsimonious specification in equation (2).

Baseline conditioning set. The baseline conditions on exact $t-1$ real GDP per capita and the realized commodity price movement, in addition to country and year fixed effects. It deliberately does not condition on pre-crisis external balance sheet positions. The current account position is mechanically linked to both the event construction and the subsequent reversal, while external debt, net foreign assets, and reserves are state variables through which structural exposure can shape crisis vulnerability. Adding those variables to the baseline would therefore change the estimand from the overall exposure, severity gradient to a comparison that holds part of the balance sheet channel fixed. The matched state model performs that comparison directly by holding the inherited balance sheet exactly constant. I then add pre-crisis private credit, capital account openness, the exchange rate regime, and overall trade intensity in separate robustness exercises. For each added control block, the baseline is re-estimated on the same observations so that coefficient changes can be distinguished from sample changes.

Commodity dependence measures the *composition* of merchandise exports, whereas the trade to GDP ratio measures the *scale* of cross border trade. The trade intensity robustness therefore distinguishes the paper's exposure margin from overall openness.

Global financial conditions. World interest rates provide an important alternative external financing channel. I construct a U.S. real short rate as the 3 month Treasury bill rate minus U.S. CPI inflation and use its annual change. Because this variable is common across countries, its level is absorbed by year fixed effects. The robustness specification interacts sudden stop onset with the change in the world real rate, allowing episodes that coincide with tighter global financial conditions to have a different average severity.

Inference and pre-trends. Standard errors are clustered at the country level with the finite sample correction. Figure 1 reports 68% and 95% confidence bands. Outcomes are cumulative changes relative to $t-1$, so the response paths are normalized at the last pre-crisis observation. I jointly test the exposure differential at $h=-3$ and $h=-2$ using a country cluster bootstrap with 999 replications; the full response paths and joint pre-onset tests are reported together.

3.2 Data

The main outcome panel covers the sudden stop sample period using data assembled from the Global Macro Database, World Development Indicators, and IMF International Financial Statistics, depending on series availability. The construction of the key exposure and control variables is as follows. Appendix Table A3 provides a consolidated source, timing, and transformation table for all variables used in the empirical analysis.

Sudden stop episodes. The risk set contains 58 countries. The reconciled broad chronology contains 135 sudden stop episodes, of which 51 are classified as systemic. The baseline econometric sample uses the systemic subset described in Section 2.1; 50 of those 51 episodes have the exact $t-1$ commodity dependence and income observations required by the baseline. The number used for a given outcome can be smaller because outcome availability differs across countries and years.

Commodity dependence. Commodity dependence is the share of commodity exports in total merchandise exports. I construct the numerator from four World Development Indicators categories: agricultural raw materials, food, fuels, and ores and metals. The main measure is the exact calendar $t-1$ observation. Using the immediately preceding year is important because the paper asks whether exposure measured before onset predicts subsequent severity; it avoids using export composition realized after the crisis begins.

Commodity export prices. The contemporaneous price control is the IMF country specific Commodity Export Price Index, based on historical fixed commodity export weights. I use the log change during the onset year. This is a shock measure, not the structural exposure variable: two countries can have different values of CD_{t-1} even when they face similar price changes, and vice versa.

Income and robustness controls. Real GDP per capita comes from the World Development Indicators and is measured at $t-1$. The financial robustness adds domestic credit to the private sector by banks as a share of GDP, also at $t-1$. The institutional robustness adds the Chinn and Ito (2008) capital account openness index and the Ilzetzki et al. (2022) exchange rate regime, lagged one year. The global financial conditions robustness uses the 3 month U.S. Treasury bill rate and U.S. CPI inflation from FRED to construct the real short rate.

Outcomes and transformations. The six outcomes are real GDP per capita, real consumption per capita, real investment, the current account to GDP ratio, the real effective exchange rate, and real stock prices. I follow the event study convention in Bianchi and Mendoza (2020) and work with HP filtered cyclical components for the macroeconomic paths. For GDP, consumption, and investment I extract the cyclical component of the logarithm with $\lambda = 100$; for CA/GDP I apply the same annual frequency filter to the ratio in levels, so movements remain in percentage points. This transformation is separate from the construction of the sudden stop chronology: Bianchi and Mendoza (2020) identify candidate events from year on year changes in the raw CA/GDP ratio and then study HP filtered macroeconomic paths around those dates. I take their systemic dates as given and preserve that event study transformation.

The real effective exchange rate and stock price series are indexes whose arithmetic level changes are not comparable across countries with different index bases. I therefore use log differences,

$$\Delta^h y_{it}^{\text{idx}} \equiv \log y_{i,t+h} - \log y_{i,t-1}. \tag{4}$$

Under the convention used in the data, a negative real exchange rate log change is a real depreciation; a negative stock price log change is an equity loss.

3.3 Results

Figure 1 reports the $CD_{75} - CD_{25}$ differential from $h = -3$ through $h = 3$, and Table 2 reports the corresponding post-onset estimates. The most precise and persistent result is the current account response.

Current account. At $h = 1$, moving from the 25th to the 75th percentile of pre-crisis commodity dependence is associated with a 2.93 percentage point larger current account reversal, with a standard error of 1.21 and $p = 0.015$. The differential remains positive and significant at $h = 2$, at 1.89 percentage points ($p = 0.041$). Expressed locally, a 10 percentage point increase in commodity dependence is associated with a 0.58 percentage point larger reversal at one year.

Other outcomes. The current account is the most precise and persistent response, but the broader pattern is consistent with a deeper crisis at higher exposure. Consumption falls 2.17 percentage points more at onset and 2.45 percentage points more at $h = 1$ ($p = 0.029$). Equity prices also fall significantly more at onset and one year later, with the $h = 1$ log differential equal to -0.413 ($p = 0.017$). The real exchange rate depreciates more along the post-onset path, reaching a -0.120 log point differential at $h = 3$ ($p < 0.05$). Investment moves sharply in the same direction at onset, although the one year estimate is less precise. GDP shows no statistically precise exposure differential. This pattern, larger external adjustment, weaker absorption, real depreciation, and equity losses, is the empirical object the model is designed to interpret.

Pre-onset dynamics. The joint bootstrap tests at $h = -3$ and $h = -2$ do not reject zero pre-onset exposure differentials for any of the six outcomes. The bootstrap p-values are 0.588 for GDP, 0.359 for consumption, 0.342 for investment, 0.413 for CA/GDP, 0.127 for the real exchange rate, and 0.467 for equity prices. These tests support reading the post-onset pattern as a distinct crisis severity ordering rather than the continuation of a visible differential trend.

Table 2: Pre-crisis Commodity Exposure and Systemic Sudden Stop Severity

Variable	$h = 0$	$h = 1$	$h = 2$	$h = 3$
GDP	-0.12	-0.15	$+0.88$	$+0.85$
	(1.03)	(0.93)	(1.13)	(0.80)
Consumption	-2.17^{*}	-2.45^{**}	-1.61	-0.62
	(1.16)	(1.12)	(1.27)	(0.98)
Investment	-5.95^{*}	-6.14	-0.20	$+1.97$
	(3.61)	(3.97)	(3.99)	(2.96)
CA/GDP (pp)	$+0.84$	$+2.93^{**}$	$+1.89^{**}$	$+0.36$
	(1.07)	(1.21)	(0.93)	(0.68)
RER (log-diff)	-0.010	-0.061	-0.045	-0.120^{**}
	(0.063)	(0.057)	(0.044)	(0.061)
Stock prices (log-diff)	-0.394^{**}	-0.413^{**}	-0.278	-0.242
	(0.193)	(0.172)	(0.191)	(0.189)

Notes: Each coefficient is the implied $CD_{75} - CD_{25}$ differential from equation (2). In the baseline systemic sample, $CD_{25} = 22.5\%$ and $CD_{75} = 72.9\%$. Standard errors clustered by country are in parentheses. $h = -1$ is the normalization year. Joint pre-trend tests for $h = -3, -2$ are reported in Appendix B. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

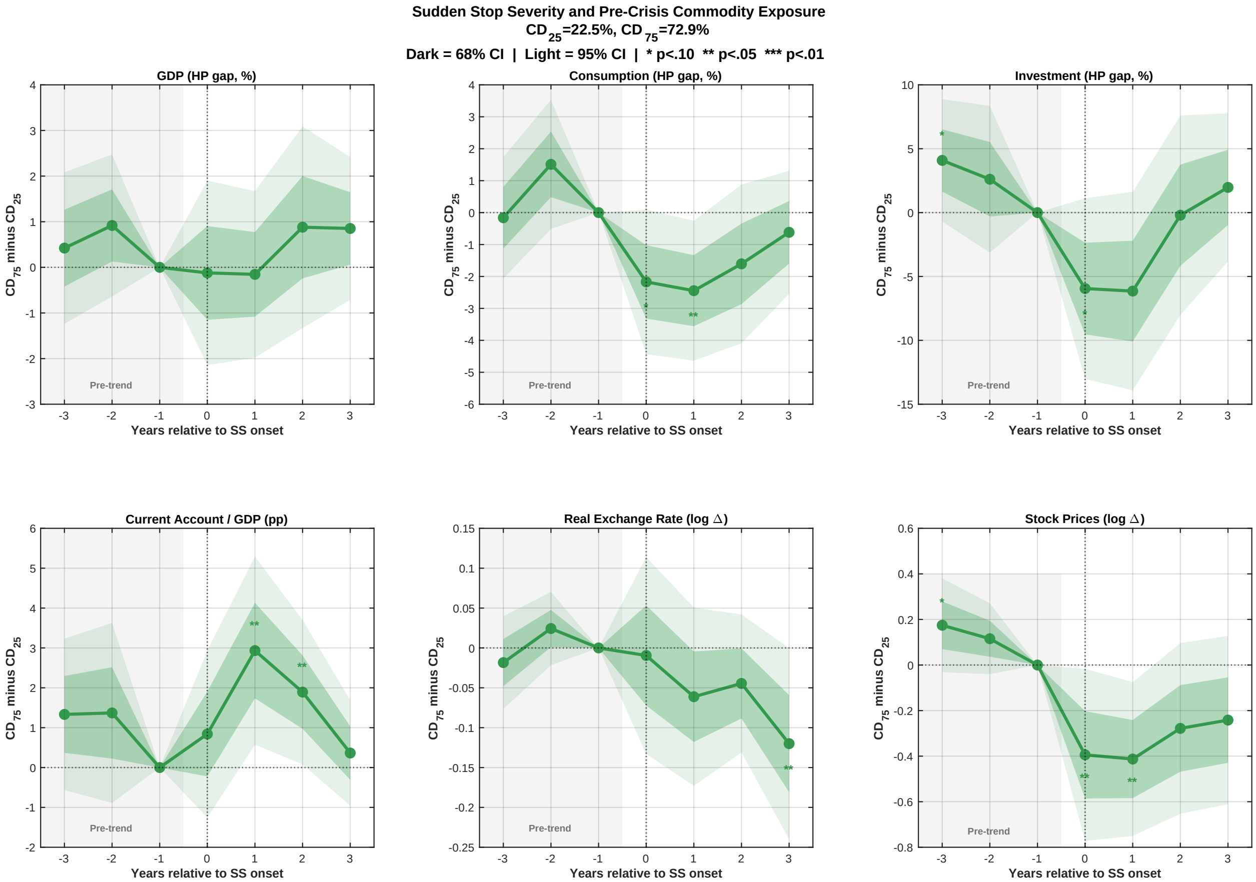

Figure 1: Commodity Exposure and Sudden Stop Severity

Notes: Each panel reports the estimated difference in the cumulative response to a systemic sudden stop between the 75th and 25th percentiles of pre-crisis commodity dependence. The baseline conditions on exact $t-1$ real GDP per capita and the contemporaneous country specific commodity export price movement, with country and year fixed effects. All paths are normalized to zero at $h=-1$. Dark bands are 68 % confidence intervals and light bands are 95 % confidence intervals. Standard errors are clustered by country. Stars denote $^{*}p<0.10$, $^{**}p<0.05$, and $^{***}p<0.01$.

3.4 Robustness

The robustness exercises are designed to answer specific alternative explanations rather than to search over arbitrary control sets. Table 3 summarizes the one year exposure differential.

Domestic balance sheets and institutions. Adding pre-crisis private credit leaves the main result essentially unchanged. On the common Credit/GDP sample, the CA/GDP differential is 3.19 percentage points without credit and 3.38 with credit. Adding capital account openness and the exchange rate regime yields 3.34 percentage points; on that exact sample, the baseline is 3.22 and the credit specification 3.40. These comparisons matter because they show that the augmented coefficients are not driven merely by losing countries with missing controls.

Global real interest rates. The world rate specification adds $SS_{it} \times \Delta R_t^{world}$. The one year current account differential remains 2.93 percentage points, compared with 2.93 in the baseline on the same sample; consumption changes from -2.45 to -2.40. The exposure gradient is therefore unchanged when sudden stops occurring during tighter global monetary conditions are allowed to have different average severity. This empirical robustness complements the model experiment, which fixes the world interest rate to isolate the commodity exposure mechanism.

Trade intensity. Commodity dependence is a composition measure, so a natural concern is that it may simply proxy for the overall size of the external sector. I therefore add exact $t-1$ Trade/GDP and $SS_{it} \times Trade/GDP_{i,t-1}$. On the common sample, the one year CA/GDP differential is 2.76 percentage points without the trade intensity terms and 3.59 percentage points with them. Consumption, investment, and equity price differentials also remain negative. The commodity exposure ordering is therefore not attenuated by allowing crisis severity to vary with overall trade intensity.

Flexible commodity price interaction. Allowing the gradient to vary with the realized commodity price movement leaves the average ordering essentially unchanged; the centered interaction results are reported in Appendix B. This is consistent with commodity exposure operating as a persistent characteristic of tradable income rather than through any single price realization.

Clean comparison sample. The clean comparison specification excludes observations within three years of a future systemic sudden stop, reducing contamination of the comparison observations by economies already approaching another event. The one year CA/GDP differential remains 2.87 percentage points.

Static country characteristics. As an additional timing check, Appendix Table B5 replaces episode specific $t-1$ commodity dependence and real GDP per capita with static country averages over 2021–2023. On the same observations, the one year current account differential is 3.31 percentage points under the static specification, compared with 2.93 percentage points in the preferred exact-$t-1$ specification. Because the 2021–2023 characteristics postdate many of the historical sudden stops, I treat this exercise only as a sensitivity check rather than as an alternative preferred specification.

Table 3: Robustness of the Commodity Exposure Gradient at $h = 1$

Specification	GDP	Consumption	Investment	CA/GDP	RER	Stock prices
Baseline	−0.15	−2.45**	−6.14	+2.93**	−0.061	−0.413**
+ Credit/GDP	+0.04	−2.37*	−6.89*	+3.38***	−0.031	−0.324**
+ Credit + KA openness + ER regime	−0.26	−2.73**	−7.76**	+3.34***	−0.026	−0.356***
+ World real rate	−0.15	−2.40**	−6.10	+2.93**	−0.062	−0.452**
Clean comparison sample	−0.29	−2.74**	−7.34*	+2.87**	−0.071	−0.467**
+ Trade/GDP	−1.00	−2.58**	−7.98**	+3.59***	−0.081*	−0.443***

Notes: Entries report the $CD_{75} - CD_{25}$ differential at $h = 1$. The baseline controls for exact $t-1$ real GDP per capita and the contemporaneous commodity export price movement. Credit/GDP, Chinn and Ito (2008) capital account openness, the exchange rate regime, and Trade/GDP are measured at $t-1$ when included. The Trade/GDP specification also includes $SS_{it} \times Trade/GDP_{i,t-1}$ and is estimated on its available common sample. The world rate specification adds $SS_{it} \times \Delta R_t^{world}$; the main effect of the world rate is absorbed by year fixed effects. The clean comparison specification excludes observations within three years of a future systemic sudden stop. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

3.5 Interpretation

The local projections establish a conditional ordering in the intensive margin of systemic sudden stops: crisis severity varies systematically with a predetermined structural exposure measure after allowing for development, common time effects, the realized commodity price environment, and several leading observable alternatives. The most robust individual responses are the current account, consumption, and equity prices, with the real exchange rate adjusting more gradually.

The historical data necessarily combine episodes generated by different shocks and inherited balance sheets. The model therefore performs a different but complementary task: it holds the balance sheet and commodity price shock fixed to isolate the exposure mechanism, and then reintroduces endogenous debt choice in the stochastic equilibrium. The empirical gradient motivates the mechanism; the quantitative experiments identify how exposure changes fragility and how private balance sheet adjustment modifies that fragility.

4 Theoretical Framework

The empirical analysis documents a conditional exposure gradient. The model isolates a mechanism that can generate that ordering by changing only the commodity sensitivity of tradable income in an otherwise unchanged Fisherian small open economy.

The starting point is the decentralized equilibrium of Bianchi (2011). The environment combines the three ingredients needed for the question: sudden stops emerge from inherited debt and adverse shocks; the collateral limit depends on current income valued at an endogenous relative price, creating a Fisherian depreciation collateral feedback; and the decentralized equilibrium has a social planner counterpart that isolates the associated pecuniary externality. I preserve the preferences, collateral structure, constant world interest rate, debt benchmark, and original joint domestic endowment process, and introduce only a world commodity price and a parameter governing how strongly tradable income loads on that price. This parsimonious design keeps the financial friction fixed across exposure levels, so cross exposure differences can be attributed directly to the composition of tradable income.

4.1 Households, Goods, and Financial Markets

A representative household consumes tradable and nontradable goods. Preferences are

$$E_0 \sum_{t=0}^{\infty} \beta^t u(c_t), \qquad u(c_t) = \frac{c_t^{1-\sigma} - 1}{1-\sigma}, \tag{5}$$

where aggregate consumption is a CES composite,

$$c_t = \left[\omega (c_t^T)^{-\eta} + (1-\omega)(c_t^N)^{-\eta}\right]^{-1/\eta}. \tag{6}$$

The economy trades a one period bond at the constant gross world interest rate $R = 1+r$. Bond holdings $b_{t+1} < 0$ denote debt. The period budget constraint is

$$c_t^T + p_t^N c_t^N + b_{t+1} = y_t^T + p_t^N y_t^N + R b_t, \tag{7}$$

where p_t^N is the relative price of nontradables in units of tradables. Since the nontradable good is consumed domestically,

$$c_t^N = y_t^N, \tag{8}$$

and the resource constraint reduces to

$$c_t^T + b_{t+1} = y_t^T + Rb_t. \tag{9}$$

Borrowing is limited by the market value of contemporaneous income:

$$b_{t+1} \geq -\kappa \left(p_t^N y_t^N + y_t^T \right). \tag{10}$$

The dependence of the borrowing limit on p_t^N is the source of Fisherian amplification.

The intratemporal optimality condition and nontradable market clearing imply

$$p_t^N = \frac{1-\omega}{\omega} \left(\frac{c_t^T}{y_t^N} \right)^{1+\eta}. \tag{11}$$

Thus, when a binding constraint forces tradable consumption downward, the relative price of nontradables falls. This depreciation reduces the collateral value of nontradable income and tightens the borrowing constraint further.

4.2 Commodity Exposure

Let z_t denote the tradable endowment shock in the original Bianchi economy and p_t^C the world commodity price. Tradable income is

$$y_t^T(\alpha) = z_t \left[(1-\alpha) + \alpha p_t^C \right] = z_t \left[1 + \alpha (p_t^C - 1) \right], \qquad \alpha \geq 0. \tag{12}$$

The parameter $\alpha \in [0, 1]$ measures the commodity share of tradable income. It is not the same object as the empirical commodity share CD, whose denominator is merchandise exports. The distinction matters throughout the quantitative analysis: empirical CD disciplines low and high model exposures through a separate mapping rather than being inserted directly as α.

The multiplicative formulation in equation (12) is chosen to preserve the original Bianchi tradable endowment shock for every exposure economy. An additive specification that replaced part of z_t with an independent commodity component would mechanically alter the domestic shock process as α changed. Here every economy faces the same underlying tradable quantity realization z_t; exposure changes only how the value of that tradable income responds to the common commodity price.

Equation (12) differs importantly from an additive decomposition of tradable income. The domestic tradable quantity shock z_t remains present in the entire tradable endowment for every α. At $p_t^C = 1$, $y_t^T(\alpha) = z_t$ for all exposure levels, so changing α does not mechanically change tradable income at the normalized commodity price. Instead,

$$\frac{\partial y_t^T}{\partial p_t^C} = \alpha z_t, \tag{13}$$

so higher α makes the same commodity price movement more consequential for tradable income.

The domestic endowment pair (z_t, y_t^N) follows the original joint Markov process used by Bianchi (2011), preserving the benchmark comovement between tradable and nontradable endowments. The commodity price is independent of this domestic Markov process. This separation allows the quantitative experiment to vary commodity exposure while holding the Bianchi domestic shock structure fixed. The commodity price follows

$$\log p_t^C = \phi_1 \log p_{t-1}^C + \phi_2 \log p_{t-2}^C + \varepsilon_t^C, \qquad \varepsilon_t^C \sim N(0, \sigma_C^2). \tag{14}$$

The quantitative calibration of $(\phi_1, \phi_2, \sigma_C)$ is described in Section 5.

Because the commodity price process is independent of the domestic process and is normalized so that $E[p_t^C] = 1$, the stationary mean of tradable income is one for every α:

$$E[y_t^T(\alpha)] = 1. \tag{15}$$

Commodity exposure therefore changes the composition and risk of tradable income rather than its unconditional mean.

4.3 Recursive Equilibrium

The endogenous state consists of inherited bond holdings and the exogenous state,

$$s_t = (b_t, z_t, y_t^N, p_t^C, p_{t-1}^C).$$

The household solves

$$V(s_t) = \max_{b_{t+1}} \{u(c_t) + \beta E_t[V(s_{t+1})]\} \tag{16}$$

subject to (9), (10), (11), and (12). Let $\mu_t \geq 0$ be the multiplier on the collateral constraint. The Euler equation can be written as

$$u_T(c_t^T, y_t^N) = \beta R E_t\big[u_T(c_{t+1}^T, y_{t+1}^N)\big] + \mu_t, \tag{17}$$

where

$$u_T(c_t^T, y_t^N) = \omega c_t^{1+\eta-\sigma}(c_t^T)^{-1-\eta}.$$

When $\mu_t = 0$, the household smooths consumption intertemporally. When the collateral constraint binds, the household is forced to reduce debt, tradable consumption falls, and the current account improves sharply.

I define the current account as

$$CA_t = b_{t+1} - b_t. \tag{18}$$

A sudden stop therefore corresponds to a large positive current account adjustment together with a binding collateral constraint.

4.4 Commodity Prices and Fisherian Amplification

The commodity price affects borrowing capacity directly through tradable income and indirectly through the endogenous price of nontradables. Let

$$\underline{b}_{t+1} = -\kappa \left(p_t^N y_t^N + y_t^T\right) \tag{19}$$

denote the collateral lower bound. Holding p_t^N fixed,

$$\frac{\partial \underline{b}_{t+1}}{\partial p_t^C} = -\kappa\alpha z_t. \tag{20}$$

Hence, a fall in the commodity price raises the debt lower bound, reducing borrowing capacity, by more when α is larger.

When the collateral constraint binds, substituting the borrowing limit into the resource constraint gives

$$c_t^T = (1+\kappa)y_t^T + Rb_t + \kappa p_t^N y_t^N. \tag{21}$$

Using (11) and differentiating locally with respect to p_t^C,

$$\frac{dc_t^T}{dp_t^C} = \frac{(1+\kappa)\alpha z_t}{1-\kappa(1+\eta)\dfrac{p_t^N y_t^N}{c_t^T}}, \tag{22}$$

provided

$$1-\kappa(1+\eta)\frac{p_t^N y_t^N}{c_t^T} > 0.$$

The denominator is the Fisherian feedback. A commodity price decline lowers tradable income and tradable consumption; lower tradable consumption reduces p_t^N; and the fall in p_t^N further reduces collateral capacity.

Combining the direct and price effects,

$$\frac{d\underline{b}_{t+1}}{dp_t^C} = -\kappa\left[\alpha z_t + y_t^N \frac{dp_t^N}{dp_t^C}\right]. \tag{23}$$

Equations (22)–(23) establish the local amplification channel. They do not, by themselves, imply that every endogenous crisis statistic must be globally monotone in α, because inherited debt and the set of crisis dates also respond endogenously. The quantitative exercise below therefore asks whether the full nonlinear economy preserves the ordering suggested by this local mechanism.

Proposition 1 (Commodity exposure and local Fisherian amplification)**.** *Consider two economies with $\alpha_H > \alpha_L$ that enter a binding regular branch state with the same inherited debt b_t, domestic endowments z_t and y_t^N, and normalized commodity price $p_t^C = 1$. For the same marginal adverse commodity price innovation, the high exposure economy experiences a larger decline in tradable income, a larger decline in tradable consumption, a larger fall in the relative price of nontradables, and a larger tightening of borrowing capacity. Hence, the required current account reversal is larger at higher α.*

The proposition is a local statement around a common binding state. Its purpose is to isolate the channel, not to characterize unconditional crisis frequency. At $\alpha = 0$, the Fisherian mechanism remains fully operative for domestic shocks; commodity price innovations simply do not affect tradable income. The finite shock matched state experiment in Section 6 solves the nonlinear binding equilibrium and asks whether the ordering in Proposition 1 survives for an economically meaningful adverse price innovation. The stochastic equilibrium exercise then allows inherited debt and crisis dates to adjust endogenously.

The mechanism can be summarized in five steps:

$$p_t^C \downarrow \Rightarrow\; y_t^T \downarrow \Rightarrow\; c_t^T \downarrow \Rightarrow\; p_t^N \downarrow \Rightarrow\; \text{collateral capacity} \downarrow \Rightarrow\; \text{forced deleveraging} \uparrow.$$

A higher α strengthens the first link and, when the collateral constraint binds, propagates through every subsequent link.

4.5 Current Price Income and Real Output

The model contains two distinct income concepts that are useful for interpreting the empirical results. The collateral constraint depends on income valued at current relative prices,

$$Y_t^{\text{value}} = p_t^N y_t^N + y_t^T(\alpha). \tag{24}$$

Commodity prices affect this object directly through $y_t^T(\alpha)$ and indirectly through p_t^N.

By contrast, the model has no endogenous production response. I therefore construct constant price real GDP as

$$Y_t^{\text{real}} = z_t + \bar{p}^N y_t^N, \qquad \bar{p}^N = \frac{1-\omega}{\omega}. \tag{25}$$

This is a quantity measure valued at the steady state relative price. Since all α, economies face the same realizations of z_t and y_t^N, Y_t^{real} is identical across exposure levels on the same dates. This distinction is intentional: the model isolates a valuation, collateral, and absorption channel rather than a production channel.

5 Calibration

The calibration follows one organizing principle: preserve the Bianchi sudden stop environment and discipline only the new objects introduced by commodity exposure. Preferences, the collateral parameter, the world interest rate, the domestic endowment process, and the zero-exposure debt benchmark are taken from Bianchi (2011). The commodity price process is disciplined externally, and commodity exposure is mapped from the empirical interquartile comparison using national accounts information. The financial friction is therefore held fixed rather than chosen to reproduce the empirical exposure gradient.

A second principle is to keep the quantitative roles distinct. The empirically mapped values $\alpha_L = 0.30$ and $\alpha_H = 0.44$ discipline the controlled matched state comparison and the borrowing rule decomposition. The baseline stochastic equilibrium is reported over the rectangular grid range $\alpha \in [0, 0.40]$, where the pre-specified numerical reliability criterion is satisfied, and a state dependent feasible debt implementation extends the stochastic robustness check to $\alpha = 0.44$. The social planner is solved over the same clean baseline range through $\alpha = 0.40$, beginning at $\alpha = 0$ where the model reproduces the Bianchi benchmark. These exercises serve different purposes and should not be read as competing estimates of the same country parameter.

5.1 Structural Parameters and Domestic Endowments

Table 4 summarizes the structural calibration. The coefficient of relative risk aversion is $\sigma = 2$, the intratemporal elasticity of substitution is $1/(1+\eta) = 0.83$, the discount factor is $\beta = 0.906$, the tradable weight is $\omega = 0.307$, and the world interest rate is $r = 0.04$. Most importantly, the collateral parameter remains at

$$\kappa = 0.3235,$$

the value in Bianchi (2011). Thus, differences across exposure economies are not generated by assigning tighter borrowing constraints to more commodity dependent economies.

Rather than approximating the domestic shocks with separate AR(1) processes, I use the original 16 state joint Markov chain for tradable and nontradable endowments from Bianchi's `proc_shock.mat`. I relabel the tradable endowment in that process as z_t. Under its stationary distribution,

$$E[z_t] = E[y_t^N] = 1,$$

$$sd(\log z_t) = 0.0571, \qquad sd(\log y_t^N) = 0.0556,$$

$$\rho_1(\log z_t) = 0.526, \qquad \rho_1(\log y_t^N) = 0.621,$$

and

$$corr(\log z_t, \log y_t^N) = 0.833.$$

Preserving this joint process keeps both benchmark domestic risk and its positive cross endowment correlation unchanged.

Table 4: Structural Calibration

Parameter	Description	Value	Source
Preferences and financial structure			
σ	Risk aversion	2	Bianchi (2011)
$1/(1+\eta)$	Intratemporal elasticity	0.83	Bianchi (2011)
β	Discount factor	0.906	Bianchi (2011)
ω	Tradable weight	0.307	Bianchi (2011)
r	World interest rate	0.04	Bianchi (2011)
κ	Collateral parameter	0.3235	Bianchi (2011)
Commodity price			
ϕ_1	First AR coefficient	0.95	Drechsel and Tenreyro (2018)
ϕ_2	Second AR coefficient	-0.13	Drechsel and Tenreyro (2018)
σ_C	Innovation std. dev.	0.1064	Drechsel and Tenreyro (2018)
Numerical, stochastic equilibrium			
N_B	Benchmark debt grid points	100	Bianchi (2011)
$[\underline{b}, \bar{b}]$	Benchmark debt grid, $\alpha = 0$	$[-1.02, -0.20]$	Bianchi (2011)
N_B^{exo}	Domestic Markov states	16	Bianchi (2011)
N_C	Commodity price nodes	5	This paper
N_C^2	Commodity AR(2) pair states	25	This paper
N_S	Joint exogenous states	400	This paper

Notes: The domestic endowment process is the original 16 state Bianchi joint Markov chain. The commodity price process is independent of the domestic process and is discretized as a Markov chain in the pair (p_t^C, p_{t-1}^C). Structural parameters are common across all exposure levels.

5.2 Commodity Price Process

Following Drechsel and Tenreyro (2018), the annual log commodity price follows

$$\log p_t^C = 0.95 \log p_{t-1}^C - 0.13 \log p_{t-2}^C + \varepsilon_t^C, \qquad sd(\varepsilon_t^C) = 0.1064. \tag{26}$$

The implied first order autocorrelation is approximately 0.84, and the stationary standard deviation of $\log p_t^C$ is 0.198.

I discretize the AR(2) in companion form. Five nodes for the current log price imply $5^2 = 25$ current lag pair states. The grid width is chosen so that the discrete chain reproduces the 0.198 stationary standard deviation of the continuous process; the discretized first order autocorrelation is 0.816, compared with 0.841 in the continuous AR(2). After exponentiation, the discrete price process is normalized in levels so that

$$E[p_t^C] = 1. \tag{27}$$

This normalization is economically useful rather than cosmetic. Because tradable income is $y_t^T = z_t[(1-\alpha) + \alpha p_t^C]$, it guarantees that changing α changes exposure to commodity price risk without mechanically changing mean tradable income. Together with $E[z_t] = 1$ and independence between the domestic and commodity processes, it implies $E[y_t^T(\alpha)] = 1$ for every α.

The baseline stochastic law is not pruned. The joint exogenous transition matrix is the exact Kronecker product of the 16 state Bianchi transition matrix and the 25 state commodity price transition matrix, producing 400 exogenous states and preserving independence between the domestic and commodity shocks.

5.3 Empirical Discipline for Commodity Exposure

The main empirical exposure variable and the model parameter have different denominators. Empirically,

$$CD = \frac{\text{commodity exports}}{\text{merchandise exports}},$$

whereas α is the commodity share of *tradable income*. Setting $\alpha = CD$ would therefore equate two conceptually different ratios.

I construct a value added proxy for the commodity share of tradable income,

$$\alpha_{i,t-1}^{VA} = \frac{VA_{i,t-1}^{Agr} + VA_{i,t-1}^{Mining}}{VA_{i,t-1}^{Agr} + VA_{i,t-1}^{Mining} + VA_{i,t-1}^{Manufacturing}}. \tag{28}$$

The denominator excludes services because the model's α splits tradable income between commodity and non commodity components. Agriculture (including hunting, forestry, and fishing where jointly reported) and mining proxy for the commodity component, while manufacturing proxies for non-commodity tradable activity.

The sectoral inputs come from the United Nations National Accounts Official Country Data, Table 2.1. I use the calendar year immediately preceding each systemic sudden stop, matching the timing of $CD_{i,t-1}$ in the empirical analysis. To avoid combining internally inconsistent national accounts observations, agriculture, mining, and manufacturing must form a homogeneous triplet within the same country year, Series, currency, SNA system, and fiscal year convention. When more than one homogeneous triplet is available, the baseline selection gives priority to the newest SNA system and then to the largest numeric Series code. This procedure yields 46 usable systemic episodes in 34 countries.

Rather than selecting a small number of episodes around the empirical quartiles, I use all usable episodes to estimate a descriptive bridge between the export based exposure measure and the model object:

$$\alpha_{i,t-1}^{VA} = a + b\left(\frac{CD_{i,t-1}}{100}\right) + u_{i,t}. \tag{29}$$

Standard errors are clustered by country because several countries contribute more than one systemic episode. The estimated relationship is

$$\widehat{\alpha}^{VA} = 0.2360 + 0.2775\left(\frac{CD}{100}\right), \tag{30}$$

with cluster robust standard errors of 0.0431 and 0.0723, respectively, and $R^2 = 0.270$.

Evaluating equation (30) at the same empirical quartiles used in the local projections,

$$CD_{25} = 22.5\%, \qquad CD_{75} = 72.9\%,$$

gives

$$\widehat{\alpha}_L = 0.2985\ (0.0307), \qquad \widehat{\alpha}_H = 0.4383\ (0.0285),$$

and a high minus low difference of 0.1399 with standard error 0.0364. For transparency in the quantitative exercises, I round these regression implied endpoints to

$$\alpha_L = 0.30, \qquad \alpha_H = 0.44. \tag{31}$$

Using the unrounded fitted values produces virtually identical matched state results.

The regression provides a bridge from the empirical export share measure to the model's tradable income exposure parameter. Agriculture can include production that is not exported, and manufacturing can include processing of primary commodities, so the two measures are not numerically identical. The bridge places the empirical $CD_{75} - CD_{25}$ comparison in model units using a homogeneous national accounts source and the full set of usable systemic episodes. The mapping is disciplined independently of the empirical 2.93 percentage point current account coefficient. Appendix C reports the regression details and sensitivity to alternative selections among internally homogeneous UN national accounts series.

5.4 Exposure Grid and the Feasible Debt Domain in the Stochastic Equilibrium

The candidate stochastic grid is

$$\alpha \in \{0, 0.05, 0.10, \ldots, 0.40, 0.44\}. \tag{32}$$

The rectangular grid baseline is reported over the range that satisfies a pre-specified numerical reliability criterion; the mapped high exposure endpoint $\alpha = 0.44$ is used as an additional state dependent robustness check.

As $c_t^T \to 0$, $p_t^N \to 0$, so a necessary condition for positive tradable consumption is

$$Rb_t + (1+\kappa)y_t^T > 0. \tag{33}$$

Thus, for a rectangular debt grid to be feasible in every exogenous state,

$$b_t > -\frac{(1+\kappa)\min_s y^T(s;\alpha)}{R}. \tag{34}$$

As α rises, adverse commodity price realizations reduce the minimum value of tradable income, causing this worst state feasibility bound to move inward.

For $\alpha > 0$, I place the lower endpoint of the debt grid 0.0035 above the worst state bound in equation (34), while preserving approximately the Bianchi grid spacing and the common upper endpoint -0.20. Table 5 reports the resulting domains and the simulated mass at the lower boundary.

Table 5: Commodity Exposure and the Numerical Debt Domain

α	$\min y^T$	Feasibility bound	Grid lower bound	Grid points	Mass at lower bound (%)
0.00	0.891	-1.134	-1.020	100	0.000
0.05	0.877	-1.116	-1.112	111	0.000
0.10	0.863	-1.098	-1.094	109	0.000
0.15	0.849	-1.080	-1.077	107	0.000
0.20	0.835	-1.062	-1.059	105	0.000
0.25	0.821	-1.044	-1.041	103	0.000
0.30	0.807	-1.027	-1.023	100	0.090
0.35	0.793	-1.009	-1.005	98	0.403
0.40	0.779	-0.991	-0.987	96	0.790
0.44	0.767	-0.977	-0.973	94	1.121

Notes: The feasibility bound is $-(1+\kappa)\min_s y^T(s;\alpha)/R$. The rectangular grid stochastic equilibrium is considered numerically clean when simulated debt at the lower boundary is below the pre-specified 1 percent threshold and tradable consumption remains strictly positive. The $\alpha = 0.44$ rectangular solution is shown for transparency but lies just outside this criterion; the state dependent feasible debt robustness is used at that endpoint.

The rectangular grid comparative statics are therefore reported through $\alpha = 0.40$. At that endpoint, simulated lower boundary mass is 0.79% and tradable consumption remains strictly positive. The $\alpha = 0.44$ rectangular solution has boundary mass of 1.12%, just above the strict criterion. Section 7.6 verifies the mapped high exposure endpoint with a state dependent feasible debt implementation.

5.5 Solution Method and Benchmark Validation

I first solve the model at $\alpha = 0$ using the original Bianchi parameters, debt grid, and 16 state domestic process. Since $y_t^T(0) = z_t$, the commodity price state is irrelevant for decisions at zero exposure. The Euler equation algorithm converges to a sup norm distance below 10^{-5}. Simulating the resulting policy with Bianchi's sudden stop definition yields a crisis frequency of approximately 5.2%, close to the roughly 5.5% benchmark reported by Bianchi (2011). This provides a direct check that the zero exposure implementation remains close to the benchmark before commodity exposure is introduced.

For $\alpha > 0$, the policy is solved sequentially over the exposure grid using the previous solution as the initial guess. At each debt and exogenous state, the algorithm evaluates the collateral lower bound. If the constraint binds, debt is set at that bound; otherwise, tradable consumption is obtained from the Euler equation using a safeguarded Newton inversion of marginal utility. Policy updates are damped for numerical stability, and convergence requires a sup norm distance below 10^{-5}. The structural parameters, including $\kappa = 0.3235$, remain fixed at every exposure level.

This validation anchors the new exposure mechanism in the original sudden stop environment: the financial friction and zero exposure benchmark are inherited from Bianchi before commodity exposure is varied.

6 Quantitative Strategy

The quantitative analysis is organized around two exercises that answer different questions. The first is a controlled matched state experiment. It holds inherited leverage, domestic endowments, the pre-crisis commodity price, and the adverse commodity price innovation fixed, and varies only the commodity share of tradable income, α. This exercise isolates the intensive margin mechanism developed in Proposition 1: conditional on entering the same vulnerable state, does a larger exposure to commodity prices generate a larger Fisherian adjustment?

The second exercise solves and simulates the full stochastic decentralized equilibrium.

In that environment households know how exposed their income is to commodity price risk and choose debt accordingly. Commodity exposure can therefore change precautionary saving, the stationary debt distribution, the frequency with which the collateral constraint binds, and the dates that qualify as sudden stops. This exercise asks how the controlled amplification mechanism interacts with endogenous private balance sheet adjustment and endogenous crisis selection. Within this stochastic environment, I also use a borrowing rule counterfactual at the mapped exposure endpoints to isolate how much of the equilibrium attenuation comes from the differential precautionary borrowing response.

The two exercises have a deliberate division of labor. The matched state experiment identifies the conditional mechanism at a common balance sheet and shock. The stochastic equilibrium then shows how endogenous private adjustment and crisis selection reshape that mechanism in equilibrium, while the borrowing rule decomposition isolates the contribution of that adjustment. This distinction links the empirical exposure gradient to the policy analysis in Section 8 without conflating a controlled structural experiment with the historical local projection.

6.1 Controlled Matched State Crisis Experiment

Common crisis entry state. I normalize the domestic state and the pre-crisis commodity price to

$$z_{pre} = 1, \qquad y^N_{pre} = 1, \qquad p^C_{pre} = 1. \tag{35}$$

At $p^C_{pre} = 1$, equation (12) implies $y^T_{pre} = 1$ for every α. Hence differences in the pre-crisis allocation do not arise mechanically from different commodity shares.

The benchmark crisis entry state is constructed on the regular binding branch. I set the common pre-crisis borrowing limit to

$$\underline{b}_{pre} = -1.01, \tag{36}$$

which lies inside the original Bianchi debt domain and provides a convenient binding crisis state in which the Fisherian channel is operative. The value is not estimated from the exposure regression and is not chosen to reproduce the 2.93 percentage point empirical coefficient. I use the binding collateral constraint to recover the associated relative price of nontradables. Equation (11) then determines tradable consumption, and the resource constraint determines inherited bond holdings. This procedure implies a common inherited position of approximately $b_t = -1.0192$ in the benchmark state. The low and high exposure economies therefore enter the experiment with the same inherited debt, the same domestic endowments, the same relative price state, and the same pre-crisis commodity price.

The identifying comparison is

$$\text{same } (b_t, z_t, y_t^N, p_t^C) + \text{ same commodity price innovation} + \text{ different } \alpha. \tag{37}$$

Any difference in the post-shock allocation is therefore generated by the exposure parameter rather than by differences in inherited leverage, domestic shocks, or the financial friction.

Exposure mapping. The empirical commodity dependence measure and the model parameter α are conceptually related but not numerically identical. The bridge regression in Section 5.3 maps the empirical 25th and 75th percentiles,

$$CD_{25} = 22.5\%, \qquad CD_{75} = 72.9\%,$$

into fitted values $\widehat{\alpha}_L = 0.2985$ and $\widehat{\alpha}_H = 0.4383$. I use the rounded values

$$\alpha_L = 0.30, \qquad \alpha_H = 0.44$$

in the matched state experiment. The mapping also defines the endpoints used in the borrowing rule decomposition below; no parameter is chosen to reproduce the empirical current account coefficient. The rectangular grid stochastic equilibrium is reported through the numerically clean endpoint $\alpha = 0.40$, while the mapped high exposure value $\alpha = 0.44$ is additionally validated with the state dependent feasible debt implementation.

Common commodity price shock. Both economies receive the same one standard deviation adverse innovation to the calibrated commodity price process. A one standard deviation innovation provides a standardized shock whose size is determined by the external commodity price calibration rather than by the desired model response. Shock size sensitivity is reported below. The benchmark shock is

$$p^C_{post} = p^C_{pre} \exp(-0.1064), \tag{38}$$

which is a 10.09% decline from the normalized pre-crisis price. Tradable income after the shock is

$$y^T_{post}(\alpha) = z_{pre} \left[(1 - \alpha) + \alpha p^C_{post}\right]. \tag{39}$$

For each α, I solve the nonlinear binding equilibrium on the regular branch,

$$c^T_{post} = (1+\kappa)y^T_{post} + Rb_t + \kappa p^N_{post} y^N_{pre}, \tag{40}$$

together with equation (11). Thus, the initial commodity income loss is mechanical conditional on α, but the responses of tradable consumption, p^N, collateral values, and external adjustment are endogenous.

Outcome measurement and sign convention. I report two measures of external adjustment. The first is the change in the current account ratio,

$$\Delta(CA/Y) = 100 \left(\frac{CA_{post}}{Y^{value}_{post}} - \frac{CA_{pre}}{Y^{value}_{pre}} \right), \tag{41}$$

which is the model object closest to the empirical current account outcome. The second is forced deleveraging normalized by common pre-crisis collateral relevant income,

$$100 \times \frac{CA_{post} - CA_{pre}}{Y^{value}_{pre}}, \qquad Y^{value}_{pre} = p^N_{pre} y^N_{pre} + y^T_{pre}. \tag{42}$$

Because the denominator in equation (42) is identical across exposure levels, this measure isolates the balance sheet adjustment from the endogenous fall in the contemporaneous denominator.

For the remaining variables I report signed percentage changes relative to the pre-crisis state. A negative number therefore denotes a decline in aggregate consumption, tradable consumption, the relative price of nontradables, collateral relevant income, or the model real exchange rate index. Under the model's convention, a fall in the real exchange rate index is a real depreciation. Constant price real GDP is given by equation (25); because z and y^N are held fixed, its cross exposure differential is zero in this controlled experiment.

Inherited debt sensitivity. The strength of the Fisherian feedback depends on the initial balance sheet. To verify that the exposure ordering is not an artifact of the benchmark value in equation (36), I repeat the controlled experiment over a grid of inherited debt positions. For each value of b_t, I hold $z_{pre} = y^N_{pre} = p^C_{pre} = 1$ fixed and solve for the pre-crisis binding equilibrium. I then subject the low and high exposure economies to the same commodity price shock. This is a pure inherited debt sensitivity: within each point of the grid the two economies have the same debt and the same pre-crisis domestic state. I also record

$$D_{pre} = 1 - \kappa(1+\eta)\frac{p^N_{pre} y^N_{pre}}{c^T_{pre}} \tag{43}$$

to verify that every reported state remains on the regular branch.

6.2 Full Stochastic Decentralized Equilibrium

The matched state exercise deliberately suppresses precautionary saving. I therefore also solve the full stochastic equilibrium over the clean rectangular grid range

$$\alpha \in \{0, 0.05, 0.10, \ldots, 0.40\}. \tag{44}$$

At $\alpha = 0.40$, simulated mass at the lower debt boundary remains below the pre-specified one percent diagnostic threshold. Structural parameters are unchanged across exposure levels.

Simulation design. For each α, I simulate eight independent histories of 79,000 annual observations. Within a replication, all exposure economies face exactly the same primitive sequence (z_t, y^N_t, p^C_t). These common random numbers substantially reduce simulation noise in cross exposure comparisons because differences in outcomes are not driven by different realizations of the exogenous processes. The Monte Carlo dispersion reported below is therefore simulation uncertainty across independent shock histories, not parameter uncertainty.

The simulations record debt, aggregate consumption, current price collateral relevant income,

$$Y^{value}_t = p^N_t y^N_t + y^T_t,$$

constant price real GDP,

$$Y^{real}_t = z_t + \bar{p}^N y^N_t,$$

the current account, and the real exchange rate index

$$RER_t = \left[\omega^{1/(1+\eta)} + (1-\omega)^{1/(1+\eta)} (p^N_t)^{\eta/(1+\eta)}\right]^{(1+\eta)/\eta}. \tag{45}$$

Endogenous sudden stops. I use the executable sudden stop definition of the Bianchi benchmark. A model sudden stop occurs when the current account exceeds its own simulated mean by more than one standard deviation and the collateral constraint binds. Formally,

$$SS_t(\alpha) = 1\left\{CA_t(\alpha) > \overline{CA}(\alpha) + sd[CA(\alpha)]\right\} \times 1\{\text{constraint binds}\}. \tag{46}$$

This definition is not the empirical episode rule. Its role is to identify endogenous Fisherian crises in the model and to study how exposure changes both crisis incidence and crisis severity.

Own events, pairwise common events, and a fixed event set. Commodity exposure can change which dates satisfy equation (46). I therefore distinguish three objects. First, own event statistics summarize crises selected separately in each exposure economy. They describe the economy's endogenous crisis experience, but combine amplification with changes in the debt distribution and event selection.

Second, for each $\alpha > 0$ I define pairwise common events as

$$\mathcal{S}(0, \alpha) = \mathcal{S}(0) \cap \mathcal{S}(\alpha). \tag{47}$$

The $\alpha = 0$ and $\alpha > 0$ economies are evaluated on exactly these calendar dates and face identical primitive shocks. This comparison removes differences in realized shocks and dates within each pair while allowing debt policies to differ across exposure levels. Because the shared event set changes with α, the fixed event comparison below provides the complementary object with event composition held constant.

Third, I construct the more restrictive fixed event set

$$\mathcal{S}^{all} = \bigcap_{\alpha \in \{0, 0.05, \ldots, 0.40\}} \mathcal{S}(\alpha). \tag{48}$$

Across the eight simulated histories, this intersection contains 10,939 sudden stop events. Every exposure economy is then evaluated on exactly the same crisis dates. This exercise removes the changing conditioning set across exposure values and makes the role of endogenous balance sheet adjustment particularly transparent.

For each event I use a five period window $t-2, \ldots, t+2$. The current account reversal is the difference between the maximum post onset CA/Y in $t, t+1, t+2$ and the minimum pre-onset value in $t-2, t-1$. The consumption and real exchange rate severity measures are the corresponding pre-event peak to post-event trough. Current price income is measured from $t-1$ to t. For presentation, contractions and depreciations are plotted as negative differences, even though the underlying severity statistics are stored as positive magnitudes.

6.3 Relationship to the Empirical Exposure Gradient

The empirical and quantitative exercises answer complementary questions. The data establish a conditional exposure severity ordering across systemic sudden stops. The matched state model isolates the same balance sheet comparative static:

$$\left.\frac{\partial\, \text{crisis severity}}{\partial \alpha}\right|_{b,z,y^N,p^C,\varepsilon^C} > 0, \tag{49}$$

while the stochastic equilibrium asks how this partial effect changes once households choose debt in response to the additional risk.

The three objects are deliberately distinct. The empirical local projection averages across Bianchi–Mendoza systemic episodes generated by different shocks and inherited balance sheets. The matched state experiment fixes a binding state and a standardized commodity price shock, while the stochastic model lets debt policies and crisis incidence adjust endogenously under the model's crisis rule. Together they connect the reduced form ordering to a structural mechanism and then quantify how private balance sheet adjustment reshapes that mechanism.

The endowment structure also clarifies which empirical outcomes the model speaks to directly. It generates external adjustment, consumption, relative price, and collateral income responses; investment is outside the model because there is no capital accumulation, and constant price real GDP is identical across exposure levels on common primitive shock dates. The structural comparison therefore centers on the margins generated by the Fisherian mechanism itself.

7 Quantitative Results

The results reveal a sharp distinction between conditional fragility and endogenous equilibrium adjustment. Holding leverage fixed, higher commodity exposure substantially amplifies the response to the same commodity price shock. When leverage is chosen endogenously, households partially self-insure against that exposure, reducing debt and changing the set of crisis dates. Both margins are quantitatively relevant.

7.1 Matched State Amplification

Table 6 reports the benchmark low versus high exposure comparison. The two economies enter with the same inherited debt and domestic state and receive the same 10.09% commodity price decline. Only α differs.

Table 6: Commodity Exposure and Crisis Severity at a Common Leverage Position

	Low exposure	High exposure	High – Low
Empirical commodity dependence (%)	22.5	72.9	50.4
Mapped α	0.300	0.440	0.140
Current account reversal, $\Delta(CA/Y)$ (pp)	9.48	14.37	4.88
Forced deleveraging / pre-crisis income (pp)	7.28	9.89	2.60
Aggregate consumption change (%)	-9.52	-13.80	-4.28
Real exchange rate change (%)	-22.94	-31.68	-8.74
Relative price of nontradables change (%)	-31.69	-42.87	-11.18
Collateral relevant income change (%)	-22.51	-30.56	-8.05
Constant price real GDP change (%)	0.00	0.00	0.00

Notes: Both economies enter the experiment with identical inherited debt, domestic endowments, pre-crisis relative prices, and commodity prices and receive the same one standard deviation adverse commodity price innovation. Only α differs. The mapped values 0.30 and 0.44 are rounded from the regression implied endpoints 0.2985 and 0.4383. Negative entries denote declines relative to the common pre-crisis state. Under the model's convention, a negative real exchange rate change is a real depreciation. The exposure mapping and commodity price shock are disciplined independently of the empirical local projection coefficient. Across the inherited debt sensitivity in Figure 4, all reported matched state comparisons remain binding regular branch states, D_{pre} ranges from approximately 0.09 to 0.19, and the high minus low forced deleveraging gap remains between 2.49 and 2.68 percentage points.

The external adjustment difference is economically large. Forced deleveraging rises from 7.28 to 9.89 percentage points of common pre-crisis income, and the difference in $\Delta(CA/Y)$ is 4.88 percentage points. The magnitude is generated by the externally calibrated price shock and the exposure mapping, with all other state variables held common. The historical local projection averages across heterogeneous systemic episodes, while this experiment isolates the conditional mechanism at a common state and balance sheet.

The mechanism figure makes clear that the quantitative result is not simply the direct income arithmetic associated with a larger commodity share. In the low and high exposure economies, the same price shock lowers tradable income by 3.03 and 4.44 percent, respectively, a direct high minus low difference of only 1.41 percentage points. Endogenous responses are much larger: tradable consumption falls by 27.12 versus 37.17 percent, the relative price of nontradables falls by 31.69 versus 42.87 percent, and collateral relevant income falls by 22.51 versus 30.56 percent. The Fisherian feedback therefore transforms a relatively modest difference in the initial income loss into substantially larger differences in absorption, relative prices, and collateral values.

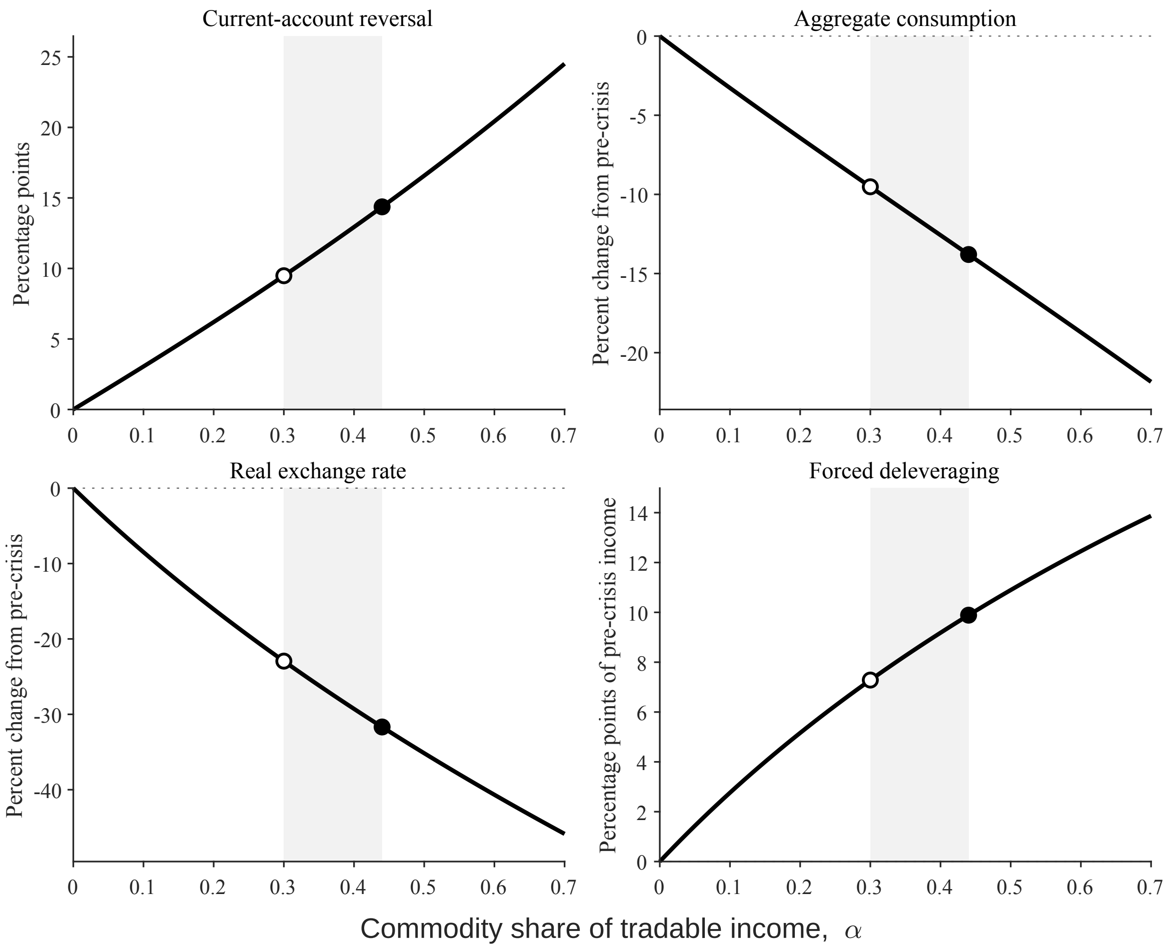

Figure 2: **Commodity exposure and crisis outcomes.** The figure reports the response to the same adverse commodity price shock over the controlled matched state experiment. The shaded region is the model exposure interval associated with the empirical CD_{25}–CD_{75} comparison. Open and filled circles denote the mapped low and high exposure economies. Consumption and the real exchange rate are shown as signed changes from the pre-crisis state; negative values denote a contraction and a real depreciation, respectively.

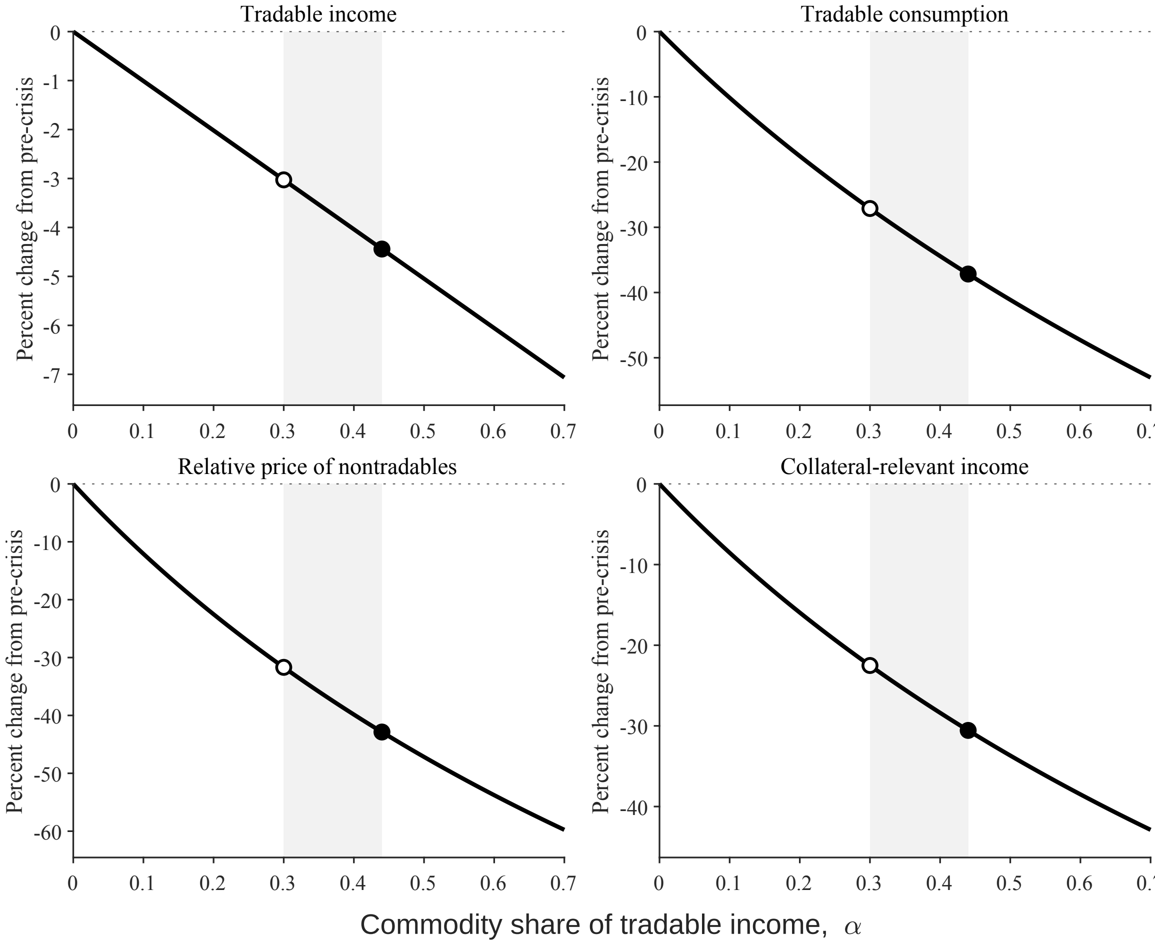

Figure 3: **Fisherian amplification in the controlled experiment.** The figure decomposes the common shock into the direct change in tradable income and the endogenous changes in tradable consumption, the relative price of nontradables, and collateral relevant income. All panels report signed percentage changes from the common pre-crisis state.

The comparison is also robust to details of the empirical mapping. Using the exact regression implied endpoints, 0.2985 and 0.4383, instead of the rounded 0.30 and 0.44 changes the high minus low difference in $\Delta(CA/Y)$ only from 4.884 to 4.875 percentage points. Alternative selections among internally homogeneous UN national accounts series imply low exposure endpoints around 0.29–0.30 and high exposure endpoints around 0.44–0.46. Reevaluating the matched state experiment at those endpoints yields a current account gap of roughly 4.9–5.6 percentage points, with the same ordering for consumption, the real exchange rate, the relative price of nontradables, and collateral relevant income. Appendix C reports these mapping diagnostics.

7.2 Robustness to Inherited Leverage and Shock Size

The exposure ordering does not depend on the single benchmark debt position. Figure 4 varies inherited external debt from approximately 26.5 to 43.6 percent of pre-crisis collateral relevant income. At every point, the high exposure economy experiences more forced deleveraging than the low exposure economy. The high minus low forced deleveraging gap remains between 2.49 and 2.68 percentage points. The high exposure economy also has a larger marginal external cost of debt throughout the grid, a result used in Section 8. The regularity condition remains strictly positive: D_{pre} ranges from approximately 0.09 to 0.19 across the reported states.

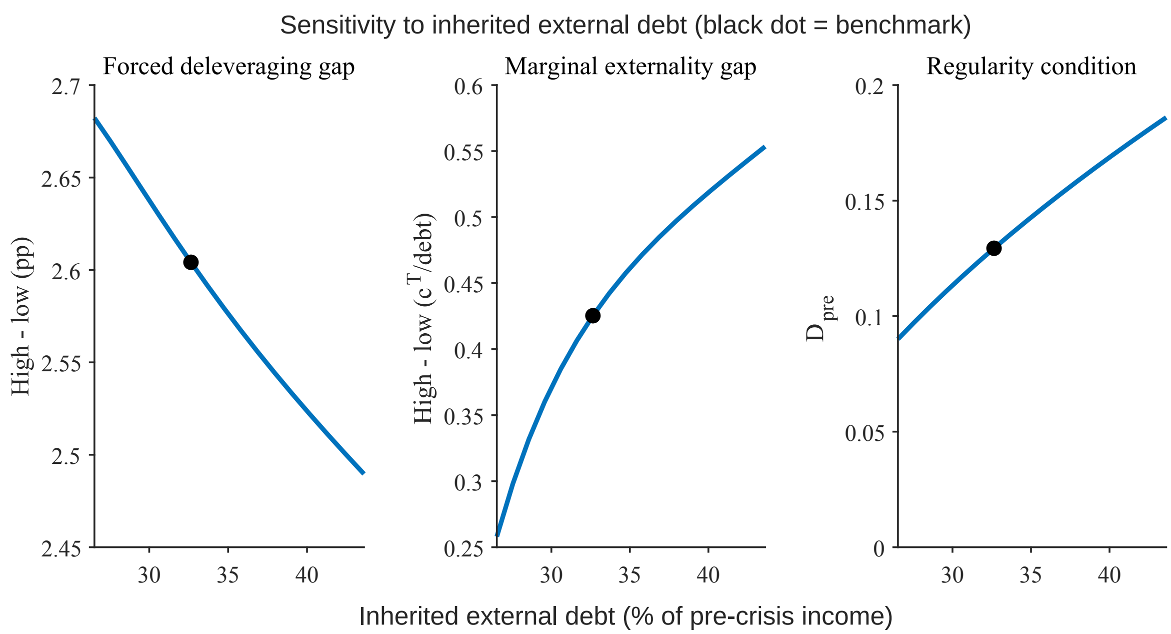

Figure 4: **Sensitivity to inherited external debt.** For each debt position, the low and high exposure economies share the same inherited debt, domestic state, and commodity price shock. The first panel reports the high minus low forced deleveraging gap; the second reports the high minus low marginal externality of debt; the third reports the regular branch denominator D_{pre}. The black dot denotes the benchmark state used in Table 6.

The same ordering also survives substantial variation in shock size. For adverse commodity price innovations of 0.5, 1, 1.5, and 2 standard deviations, the high minus low forced deleveraging gap is 1.64, 2.60, 3.22, and 3.62 percentage points, respectively. The corresponding difference in $\Delta(CA/Y)$ is 2.30, 4.88, 7.85, and 11.22 percentage points. These results are reported as robustness rather than as additional calibration targets.

7.3 Private Precautionary Adjustment in the Stochastic Equilibrium

The full stochastic equilibrium changes the interpretation because households can react before a crisis occurs. As commodity exposure rises from $\alpha = 0$ to $\alpha = 0.40$, decentralized external debt falls from 29.52 to 28.01 percent of collateral relevant income. Thus, the same structural feature that makes a given balance sheet more fragile also induces private agents to choose a safer balance sheet ex ante.

This precautionary response is visible in crisis incidence. The collateral constraint binds in 6.52 percent of periods at $\alpha = 0$ and 5.10 percent at $\alpha = 0.40$. Sudden stop frequency falls from 5.27 to 4.36 percent over the same range. The incidence pattern reflects endogenous self-insurance: conditional fragility rises with exposure at a common balance sheet, but the equilibrium balance sheet becomes safer as households anticipate that risk.

Event overlap also changes materially with α. Among the sudden stop dates of the $\alpha = 0$ benchmark, 89.3 percent remain sudden stops at $\alpha = 0.05$, but only 34.3 percent do so at $\alpha = 0.40$. This changing overlap motivates the fixed event comparison below; Appendix Figure C1 reports the incidence and overlap diagnostics.

7.4 Common Event Severity and the Hump Shaped Equilibrium Effect

Endogenous exposure changes both debt policies and the set of dates classified as sudden stops. I therefore report two common event comparisons. The first is pairwise: for each $\alpha > 0$, the exposed economy is compared with $\alpha = 0$ on dates that are sudden stops in both economies, $\mathcal{S}(0) \cap \mathcal{S}(\alpha)$. Primitive shocks and calendar dates are identical within each pair, but the conditioning set changes with α. The pairwise current account differential rises monotonically from 0.02 percentage points at $\alpha = 0.05$ to 2.47 at $\alpha = 0.40$. At the upper endpoint, aggregate consumption contracts 2.68 percentage points more, the real exchange rate depreciates 5.25 percentage points more, and collateral relevant income falls 4.15 percentage points more than at zero exposure.

The second comparison fixes the event set itself. I evaluate every exposure economy on the 10,939 crisis dates, pooled across the eight simulated histories, that are common to the clean exposure grid through $\alpha = 0.40$, while allowing each economy to retain its endogenous debt policy. This removes changes in event composition from the comparison. The current account differential rises from 1.67 percentage points at $\alpha = 0.05$ to a peak of 4.56 at $\alpha = 0.25$, then declines to 2.46 at $\alpha = 0.40$, a 46 percent decline from its peak. Aggregate consumption displays the same hump shaped pattern: its severity differential peaks at 3.94

percentage points at $\alpha = 0.25$ and falls to 2.64 at $\alpha = 0.40$.

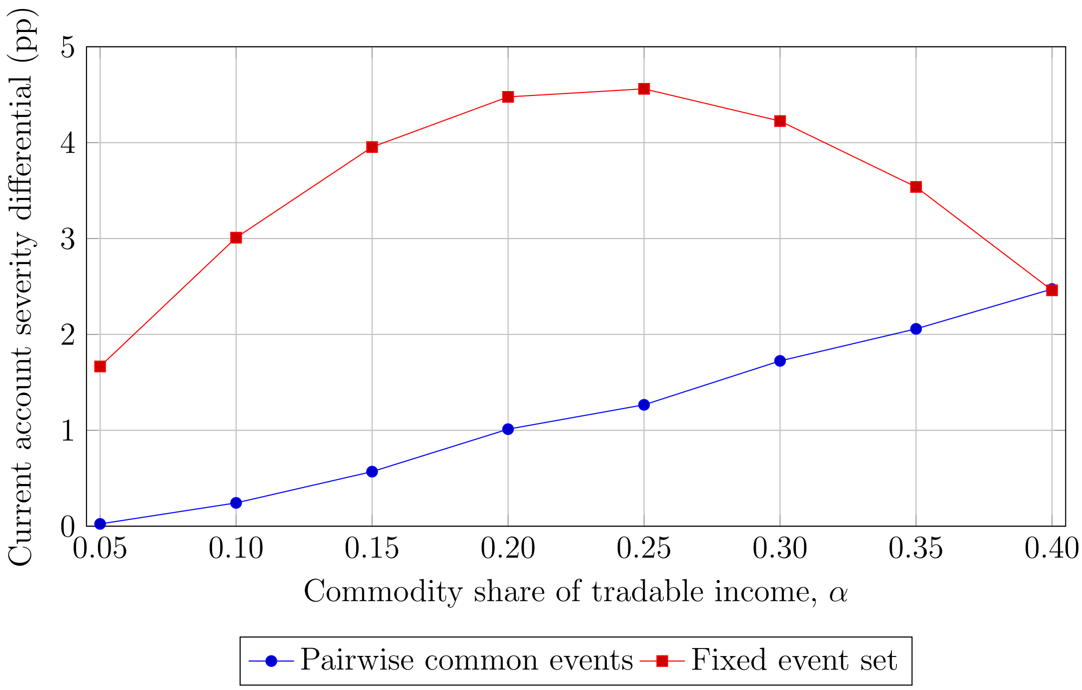

Figure 5: **Common event severity under endogenous balance sheet adjustment.** Pairwise comparisons condition on $\mathcal{S}(0) \cap \mathcal{S}(\alpha)$, so the event set varies with exposure. Fixed event comparisons use the 10,939 sudden stop events, pooled across eight simulated histories, that belong to the intersection of crisis dates common to the clean stochastic exposure grid through $\alpha = 0.40$. Both series use identical primitive shock histories and endogenous debt policies. The two lines condition on different event sets and are therefore not comparable in levels; within each series, the shape across exposure is the object of interest.

The contrast between the two lines is informative rather than contradictory. Pairwise overlap with the zero exposure benchmark falls from 89.3 percent at $\alpha = 0.05$ to 34.3 percent at $\alpha = 0.40$. The corresponding pairwise event counts are 29,711 and 11,413, compared with 10,939 events in the fixed intersection. The fixed set is therefore much more restrictive at low exposure but nearly coincides with the pairwise set at the upper clean endpoint. On that fixed set, exposure initially raises crisis severity, but beyond intermediate exposure endogenous balance sheet adjustment offsets an increasing share of the additional conditional fragility. The severity differential nevertheless remains positive relative to $\alpha = 0$. This is the central equilibrium counterpart to the matched state result: more exposure makes the same balance sheet more fragile, but forward looking households do not choose the same balance sheet.

7.5 Borrowing Rule Decomposition of Precautionary Adjustment

The preceding results show that endogenous borrowing attenuates the effect of exposure, but they do not isolate how much of that attenuation comes from the differential precautionary response. I therefore conduct a borrowing rule counterfactual at the empirically mapped values $\alpha_L = 0.30$ and $\alpha_H = 0.44$, using the state dependent feasible debt solution. In the counterfactual, the high exposure economy follows the low exposure borrowing rule wherever that rule is feasible, while retaining its own tradable income, collateral constraint, and state dependent viability limit. The exercise removes the differential borrowing response to exposure without giving the high exposure economy the borrowing capacity of the low-exposure economy. Because the assigned rule is not optimal for the high exposure economy, this is a decomposition rather than an alternative equilibrium.

To connect this decomposition to the empirical event study, I simulate 200 pseudo panels of 60 economies observed for 60 years, with 30 economies at each mapped exposure. Candidate events are defined separately within each pseudo country when the one year change in CA/Y exceeds that country's own 95th percentile, mirroring the country specific candidate event component of the empirical chronology. The model does not contain EMBI or VIX and therefore does not reproduce the empirical systemic filter. I estimate the exposure differential with country and time fixed effects and normalize outcomes at $h = -1$.

The model concentrates Fisherian deleveraging at onset, whereas the empirical current account differential peaks one year after the identified sudden stop date. I therefore do not interpret the model as matching the exact annual timing of the response. An adjustment that occurs within one model period can be distributed across adjacent calendar years in annual data, and real world sudden stops can persist longer than the model's acute crisis mechanism. To compare magnitudes without selecting the larger response ex post, I use the linear average of the onset and first post-onset responses,

$$\bar{\beta}_{0:1} \equiv \frac{\beta_0 + \beta_1}{2}. \tag{50}$$

Figure 6 summarizes the two year window comparison in terms of average exposure differentials over the onset and first post-onset years. For the current account, the high minus low exposure differential is 0.01 percentage points in equilibrium, 1.89 percentage points in the data, and 2.63 percentage points when the high exposure economy is assigned the low-exposure borrowing rule. Figure 7 reports the underlying horizon-specific paths together with empirical sampling uncertainty and simulated sample dispersion. The model adjustment is more front-loaded than the empirical response, so the comparison is not intended as an exact timing fit. Instead, the two figures show how endogenous borrowing compresses the exposure gradient and how removing that differential response generates substantially greater crisis severity.

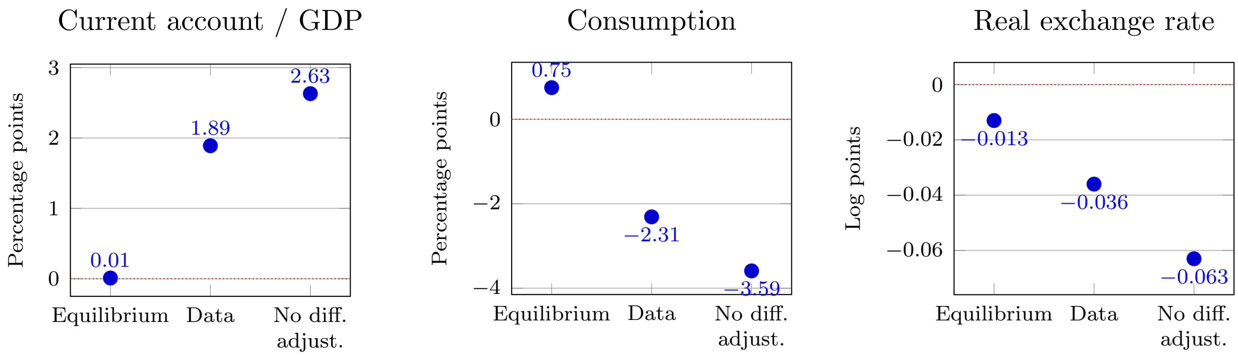

Figure 6: **Empirical and simulated exposure differentials.** Each point is the high minus low exposure differential averaged over the onset and first post-onset years, $\bar{\beta}_{0:1}$. The empirical points use the baseline local projections; the model points use the country specific P95 candidate event rule at $(\alpha_L, \alpha_H) = (0.30, 0.44)$. "No differential adjustment" assigns the high exposure economy the low exposure borrowing rule wherever feasible while preserving its own income process and borrowing constraints. The figure compares the ordering and magnitude of the exposure differential over the crisis window, not exact horizon by horizon timing. Figure 7 reports the underlying paths together with empirical confidence intervals and model Monte Carlo dispersion bands.

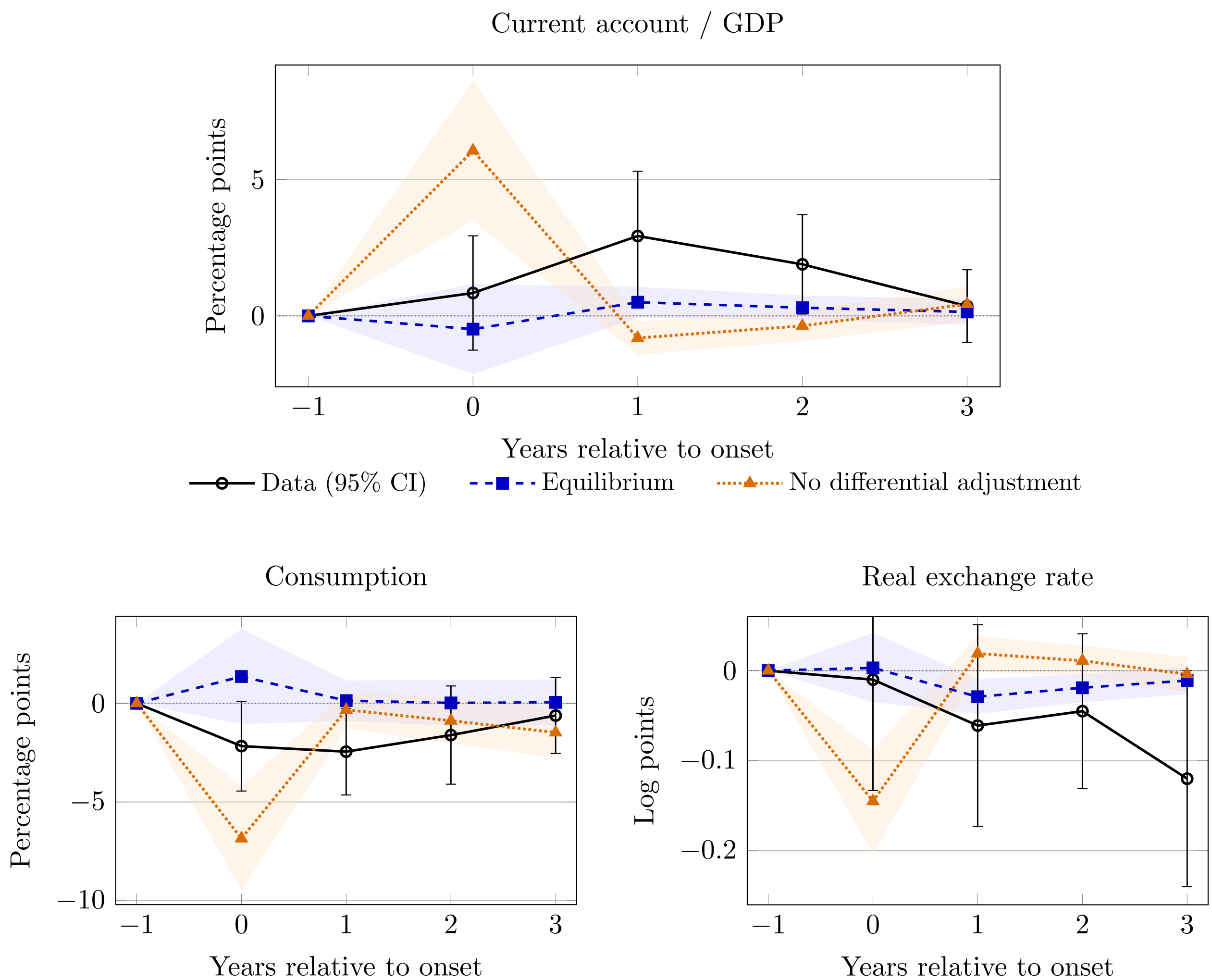

Figure 7: **Empirical and simulated exposure differentials over the crisis window.** Lines report horizon specific high minus low exposure differentials under the country specific P95 candidate event rule. Vertical intervals around the empirical estimates are 95 percent confidence intervals based on country clustered standard errors. Shaded bands around the two model paths report plus or minus one Monte Carlo standard deviation across the 200 simulated pseudo panels; they capture simulated sample dispersion rather than parameter or model uncertainty. The model's Fisherian adjustment is more front loaded than the empirical response, especially for the current account. The purpose of the comparison is to make the timing of the model data comparison transparent and to show how endogenous borrowing adjustment attenuates the exposure response, rather than to assess exact horizon by horizon fit. Figure 6 summarizes the onset and first post-onset years with the linear average in equation (50).

The current account result is the clearest summary. Removing the differential borrowing response raises the average exposure differential by 2.62 percentage points relative to equilibrium, with a paired Monte Carlo standard error of 0.06. Consumption and the real exchange rate move in the same direction under the counterfactual. Under the country specific event rule, equilibrium consumption does not reproduce the empirical high minus low ordering, consistent with strong endogenous balance sheet adjustment and event selection; removing the differential borrowing response reverses that sign and generates a substantially larger contraction.

The quantitative magnitude varies with event selection, but the attenuation result does not. Across six event definitions, the average current account differential lies between -0.01 and 0.17 percentage points in equilibrium and between 1.3 and 3.1 percentage points when differential precautionary adjustment is removed. Appendix Table C3 reports these alternatives. The small equilibrium gradient therefore does not imply a weak structural exposure channel; it reflects endogenous self-insurance in response to that channel.

7.6 Numerical Reliability

The rectangular grid solution satisfies the pre-specified one percent lower boundary criterion through $\alpha = 0.40$, where simulated boundary mass is 0.79% and tradable consumption remains strictly positive. At $\alpha = 0.44$, rectangular grid boundary mass is 1.12%, just above the criterion. A state dependent feasible debt implementation solves through $\alpha = 0.44$ with no simulated feasibility violations; its viability bound actually binds in only 0.95% of simulated observations at that endpoint. The resulting pairwise current account differential is 2.71 percentage points, compared with 2.77 in the rectangular implementation.

The state dependent robustness defines recursive viability using successor transitions with probability above 10^{-8}. For policy choices that would be infeasible following an excluded tail transition, that successor does not contribute to the Euler expectation; the maximum omitted probability in an Euler expectation is 1.2×10^{-8}. Simulation uses the full transition law, and no excluded-tail transition is realized in 631,992 simulated transitions. The planner also converges throughout the clean range through $\alpha = 0.40$; after excluding numerical debt grid corner choices, the maximum discrepancy in the interior tax-identity check is 1.6×10^{-6}, and simulated planner boundary mass is zero. All controlled binding state comparisons use the regular branch $D > 0$.

8 Macroprudential Implications

The policy analysis follows the same distinction between conditional fragility and endogenous balance sheet adjustment. At a common inherited balance sheet, higher commodity exposure makes the same adverse shock more costly because the induced relative price decline tightens collateral more strongly. In stochastic equilibrium, private agents anticipate this risk and reduce leverage. I therefore distinguish the *conditional marginal externality at common leverage* from *unconditional policy and welfare outcomes after private balance sheets adjust.*

8.1 Planner Validation and the Marginal Externality of Leverage

Before using the planner for the exposure comparison, I validate the implementation at $\alpha = 0$. With the original Bianchi calibration, the simulated consumption equivalent welfare gain from the planner is 0.1328 percent, compared with the 0.13 percent benchmark. This validation is performed before introducing the low versus high exposure policy comparison.

The economic externality is standard in the Fisherian collateral environment. A decentralized household takes the aggregate relative price of nontradables as given when choosing debt. In a future binding state, one additional unit of net foreign assets raises tradable consumption privately by

$$\left.\frac{dc^T}{db}\right|_{private} = R. \tag{51}$$

The planner internalizes that aggregate deleveraging affects p^N, which changes collateral capacity. Differentiating the binding equilibrium gives

$$\left.\frac{dc^T}{db}\right|_{planner} = \frac{R}{D}, \qquad D = 1 - \kappa(1+\eta)\frac{p^N y^N}{c^T}. \tag{52}$$

On the regular branch, $0 < D < 1$, so the social marginal value of an additional unit of net foreign assets exceeds the private value.

Let u_T denote marginal utility with respect to tradable consumption. The one step marginal pecuniary externality of inherited debt in the post-shock binding state is

$$\mathcal{E}^{MU} = \beta R u_T(c^T_{post}, y^N)\left(\frac{1}{D_{post}} - 1\right). \tag{53}$$

To express this object in current tradable consumption units, I divide by pre-shock marginal utility,

$$\widetilde{\mathcal{E}} = \frac{\beta R u_T(c^T_{post}, y^N)\left(D^{-1}_{post} - 1\right)}{u_T(c^T_{pre}, y^N)}. \tag{54}$$

A larger $\widetilde{\mathcal{E}}$ means that entering the same crisis with one additional unit of debt has a larger social marginal cost. I use this as the state contingent policy object at common leverage; unconditional policy and welfare outcomes after private balance sheets adjust are reported separately below.

8.2 Commodity Exposure at Common Leverage

Table 7 evaluates equation (54) in the same low and high exposure states used in Table 6. Inherited debt, domestic endowments, the pre-crisis commodity price, and the adverse shock are all common.

Table 7: Commodity Exposure and the Marginal External Cost of Leverage

	Low exposure	High exposure	High – Low
Mapped α	0.300	0.440	0.140
Forced deleveraging / pre-crisis income (pp)	7.28	9.89	2.60
Aggregate consumption change (%)	-9.52	-13.80	-4.28
Relative price of nontradables change (%)	-31.69	-42.87	-11.18
D_{post}	0.184	0.208	0.024
Fisherian multiplier, $1/D_{post}$	5.43	4.80	-0.64
Marginal externality, $\widetilde{\mathcal{E}}$	6.62	7.05	0.43

Notes: The two economies are evaluated at the same inherited debt position and under the same adverse commodity price innovation. The marginal externality is measured in current tradable consumption units per unit of additional debt. It is a conditional marginal object, not an unconditional optimal tax estimate.

The marginal externality rises from 6.62 to 7.05, or approximately 6.4 percent. The source of that increase is revealing. The local Fisherian multiplier $1/D$ actually falls from 5.43 to 4.80 across the two post-shock states. Greater exposure therefore does not strengthen the local collateral feedback; it pushes the economy further into the crisis region. The resulting consumption loss raises post-shock marginal utility enough to dominate the weaker local multiplier, so the social marginal cost of leverage rises even as $1/D$ falls.

The inherited debt sensitivity in Figure 4 reinforces this point. Across the full reported debt range, $\widetilde{\mathcal{E}}_H - \widetilde{\mathcal{E}}_L$ remains positive, ranging from approximately 0.26 to 0.55. The benchmark 0.43 difference therefore does not rely on the single leverage state used in Table 7. Larger adverse commodity price shocks also increase the externality gap: it rises from 0.10 under a half standard deviation price innovation to 1.79 under a two standard deviation innovation.

8.3 Endogenous Adjustment and Macroprudential Policy

Once households choose debt before the shock, higher exposure induces private precautionary deleveraging, so unconditional policy objects incorporate both the exposure effect and endogenous self insurance. Figure 8 compares the decentralized and planner balance sheets and policy wedges across the clean stochastic exposure grid.

In the decentralized equilibrium, external debt falls from 29.52 percent of collateral relevant income at $\alpha = 0$ to 28.01 percent at $\alpha = 0.40$. The planner chooses still less debt: 28.66 percent at zero exposure and 26.81 percent at $\alpha = 0.40$. Thus, the planner-private debt gap widens from about 0.86 to 1.20 percentage points as exposure rises. Private agents self insure, but the planner chooses an even safer balance sheet.

Consistent with this endogenous self insurance, unconditional policy objects need not move in the same direction as the common leverage marginal externality. The mean implemented planner tax falls from 5.04 to 4.39 percent over $\alpha \in [0, 0.40]$, and the simulated consumption equivalent welfare gain from correcting the externality falls from 0.133 to 0.100 percent. The tax equivalent externality evaluated at endogenous decentralized states similarly falls from 6.12 to 5.44 percent. These unconditional declines reflect that more exposed households enter highly leveraged states less frequently; they do not overturn the conditional result in Table 7.

The externality is also strongly state dependent. One period before a decentralized sudden stop, the tax equivalent wedge rises from 10.22 percent at $\alpha = 0$ to a peak of 10.96 percent at $\alpha = 0.25$, and remains elevated at 10.76 percent at $\alpha = 0.40$. Crisis prone balance sheets therefore carry substantially larger external costs than the unconditional average.

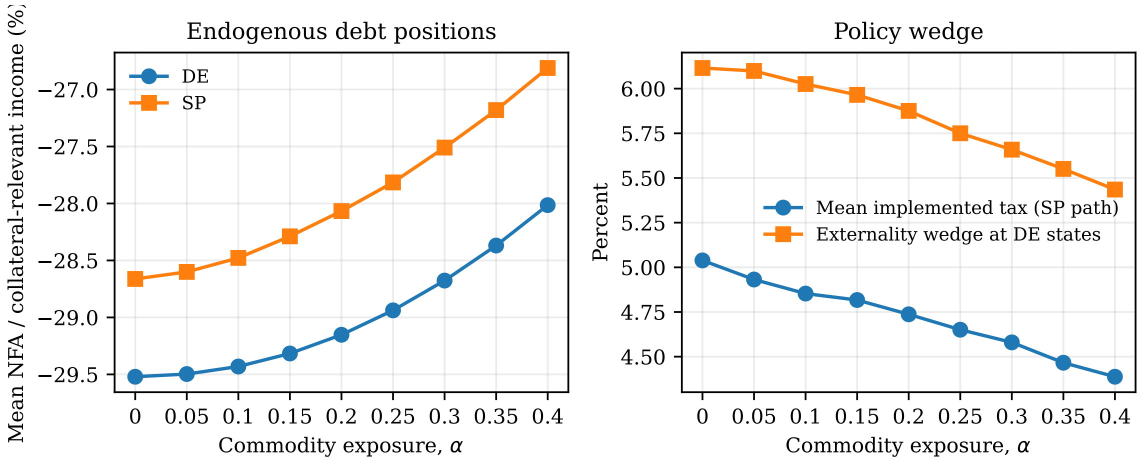

Figure 8: **Endogenous balance sheet adjustment and macroprudential wedges.** The left panel reports decentralized and planner net foreign asset positions as a share of collateral relevant income; more negative values denote more external debt. The right panel reports the mean implemented planner tax and the tax equivalent externality evaluated at decentralized equilibrium states.

The policy implication is state contingent. Commodity exposure raises the social marginal cost of carrying a given amount of leverage into an adverse commodity price state, while endogenous private adjustment can lower average leverage and therefore reduce unconditional wedges. Macroprudential vulnerability is consequently a joint function of structural exposure and the balance sheet position. The planner wedge is a tax equivalent representation of this pecuniary externality rather than a ranking of specific policy instruments.

9 Conclusion

This paper studies a simple question: how does the composition of tradable income affect financial fragility during sudden stops? The broad descriptive evidence shows the same within development group ordering in both mean and median event paths: commodity exporters undergo larger current account reversals and deeper macroeconomic adjustment than non-commodity exporters. Development status alone does not monotonically order absolute severity: developed commodity exporters also experience larger current account reversals than developing non-commodity exporters in both mean and median event paths. The conditional evidence then sharpens that fact within systemic episodes. Moving from the 25th to the 75th percentile of pre-crisis commodity dependence is associated with a 2.93 percentage point larger current account reversal one year after onset, together with larger consumption and equity price losses.

The model isolates why composition can matter even when leverage and financial frictions are initially identical. A larger commodity share makes the same world price decline a larger tradable income shock. In a Fisherian collateral environment, that initial difference is magnified through lower tradable absorption, a deeper fall in the relative price of nontradables, weaker collateral values, and forced deleveraging. At the empirically mapped exposure quartiles, the same adverse price shock generates a 4.88 percentage point larger current account reversal and 2.60 percentage points more forced deleveraging in the high exposure economy.

Endogenous balance sheet adjustment changes the equilibrium response. When households anticipate greater commodity risk, they borrow less. On the fixed set of crisis dates common to the clean stochastic exposure grid, the current account severity differential peaks at $\alpha = 0.25$ and remains positive but declines by 46 percent by $\alpha = 0.40$. A borrowing rule decomposition confirms that this self insurance is quantitatively important: assigning the high exposure economy the low exposure borrowing rule while retaining its own income and financial constraints generates much larger exposure differentials in current account adjustment, consumption, and the real exchange rate. The result reconciles two facts that would otherwise appear in tension. Greater exposure makes a given balance sheet more fragile, yet more exposed economies need not experience monotonically more severe crises because they adjust the balance sheet ex ante.

The policy analysis adds a second distinction. At common leverage, greater exposure raises the marginal external cost of debt by about 6.4 percent even though the local Fisherian multiplier falls. Exposure raises the externality by pushing the economy deeper into the crisis region, not by mechanically strengthening the local feedback at every point. Once private borrowing responds endogenously, average policy wedges need not move in the same direction as this conditional marginal externality.

The broader implication is that financial fragility is not determined by leverage alone. The composition of the income backing external liabilities affects how far an adverse shock moves the economy into the constrained region, while forward looking households partially insure against that structural risk through their borrowing choices. This interaction between exposure, leverage, and endogenous self insurance provides a compact way to understand why similar financial frictions can generate very different crisis outcomes across economies.

References

Aguiar, M. and Gopinath, G. (2007). Emerging market business cycles: The cycle is the trend. *Journal of Political Economy*, 115(1):69–102.

Bianchi, J. (2011). Overborrowing and systemic externalities in the business cycle. *American Economic Review*, 101(7):3400–3426.

Bianchi, J. and Mendoza, E. G. (2020). A fisherian approach to financial crises: Lessons from the sudden stops literature. *Review of Economic Dynamics*, 37:S254–S283.

Calvo, G. A. (1998). Capital flows and capital-market crises: the simple economics of sudden stops. *Journal of Applied Economics*, 1(1):35–54.

Calvo, G. A., Izquierdo, A., and Mejia, L.-F. (2004). On the empirics of sudden stops: the relevance of balance-sheet effects.

Cavallo, E. A. and Frankel, J. A. (2008). Does openness to trade make countries more vulnerable to sudden stops, or less? using gravity to establish causality. *Journal of International Money and Finance*, 27(8):1430–1452.

Chang, R. and Fernández, A. (2013). On the sources of aggregate fluctuations in emerging economies. *International Economic Review*, 54(4):1265–1293.

Chinn, M. D. and Ito, H. (2008). A new measure of financial openness. *Journal of comparative policy analysis*, 10(3):309–322.

Drechsel, T. and Tenreyro, S. (2018). Commodity booms and busts in emerging economies. *Journal of International Economics*, 112:200–218.

Edwards, S. (2004). Financial openness, sudden stops, and current-account reversals. *American Economic Review*, 94(2):59–64.

Faris, A. and Granados, C. (2026). Terms of trade fluctuations and sudden stops in a small open economy. *Review of International Economics*, 34(1):20–38.

Fernández, A., Schmitt-Grohé, S., and Uribe, M. (2017). World shocks, world prices, and business cycles: An empirical investigation. *Journal of International Economics*, 108:S2–S14.

Forbes, K. J. and Warnock, F. E. (2012). Capital flow waves: Surges, stops, flight, and retrenchment. *Journal of international economics*, 88(2):235–251.

Fornero, J. and Kirchner, M. (2018). Learning about commodity cycles and saving-investment dynamics in a commodity-exporting economy. *53rd issue (March 2018) of the International Journal of Central Banking*.

García-Cicco, J., Pancrazi, R., and Uribe, M. (2010). Real business cycles in emerging countries? *American Economic Review*, 100(5):2510–2531.

Gazzani, A., Herrera, V., and Vicondoa, A. (2025). The asymmetric effects of commodity price shocks in emerging economies. *European Economic Review*, page 105219.

Ilzetzki, E., Reinhart, C. M., and Rogoff, K. S. (2022). Rethinking exchange rate regimes. In *Handbook of international economics*, volume 6, pages 91–145. Elsevier.

Jordà, Ò. (2005). Estimation and inference of impulse responses by local projections. *American economic review*, 95(1):161–182.

Mendoza, E. G. (1995). The terms of trade, the real exchange rate, and economic fluctuations. *International Economic Review*, pages 101–137.

Mendoza, E. G. (2010). Sudden stops, financial crises, and leverage. *American Economic Review*, 100(5):1941–1966.

Müller, K., Xu, C., Lehbib, M., and Chen, Z. (2025). The global macro database: A new international macroeconomic dataset (version 2026-03). Working Paper 33714, National Bureau of Economic Research.

Neumeyer, P. A. and Perri, F. (2005). Business cycles in emerging economies: the role of interest rates. *Journal of Monetary Economics*, 52(2):345–380.

Schmitt-Grohé, S. and Uribe, M. (2018). How important are terms-of-trade shocks? *International Economic Review*, 59(1):85–111.

Uribe, M. and Yue, V. Z. (2006). Country spreads and emerging countries: Who drives whom? *Journal of International Economics*, 69(1):6–36.

A Sample and Sudden Stop Episodes

Table A1 distinguishes between the broad sudden stop episodes used for descriptive event evidence and the systemic episodes used in the main local projection analysis. I inherit these episode dates from the Bianchi–Mendoza chronology rather than recomputing the event filter in my sample. In that chronology, the 95th percentile current account filter is computed within country, and the systemic indicator additionally applies the global stress classification described in Section 2.1.

Descriptive sample construction. The source event workbook contains 137 candidate sudden stop episodes and manually flags Morocco 1981 and 1995 as removed.[1] I retain those two source file exclusions consistently, leaving 135 episodes in the broad descriptive chronology. Both excluded Moroccan episodes are non-systemic, so this cleaning rule affects Table 1 but not the systemic chronology used by the local projections. No other episode is dropped globally. When a particular macroeconomic series is unavailable, or when an individual cell was removed during the underlying `Macro_lookup` cleaning, primarily unusually large real exchange rate observations that would otherwise dominate a group path, that cell remains missing only for that outcome and relative year. The episode continues to contribute to every other available outcome. Means and medians are therefore computed outcome by outcome and relative year by relative year using the available observations rather than imposing a common balanced panel.

Table A1: Sudden Stop Episodes by Development Status, Commodity Exporter Status, and Systemic Classification

Country	Developing	Commodity exporter	Year	Systemic SS episode
ARG	1	1	1982	1
ARG	1	1	1990	0
ARG	1	1	1995	1
ARG	1	1	2002	1
AUS	0	1	2008	0
AUT	0	0	2014	0
BGR	0	0	1994	0
BGR	0	0	2009	1
BRA	1	1	1983	1

Continued on next page

[1] The workbook's ReadMe text refers to Morocco 1981 and 1983, whereas the episode level removal flags are attached to 1981 and 1995. I follow the episode level flags used in the workbook calculations.

Country	Developing	Commodity exporter	Year	Systemic SS episode
BRA	1	1	1984	0
BRA	1	1	1987	0
BRA	1	1	2003	0
CAN	0	1	1982	1
CAN	0	1	1996	0
CHE	0	0	2010	0
CHL	1	1	1982	0
CHL	1	1	1983	1
CHL	1	1	1986	0
CHL	1	1	1999	1
CHN	1	0	1992	0
CHN	1	0	2005	0
CIV	1	1	1984	1
COL	1	1	1989	0
COL	1	1	1999	1
DEU	0	0	1981	0
DEU	0	0	2004	0
DEU	0	0	2006	0
DEU	0	0	2009	0
DNK	0	0	1999	1
DNK	0	0	2010	0
DOM	1	0	1981	0
DOM	1	0	1987	0
DOM	1	0	1994	0
DOM	1	0	2011	0
DZA	1	1	1991	0
DZA	1	1	1999	0
ECU	1	1	1983	1
ECU	1	1	1988	0
ECU	1	1	1999	1
EGY	1	0	1991	0
ESP	0	0	1984	0
ESP	0	0	2009	0

Continued on next page

Country	Developing	Commodity exporter	Year	Systemic SS episode
ESP	0	0	2012	0
FIN	0	0	1986	0
FIN	0	0	1995	0
FIN	0	0	2000	1
FRA	0	0	1983	0
FRA	0	0	1997	1
FRA	0	0	2015	0
GBR	0	0	1981	0
GBR	0	0	1991	1
GRC	0	0	1993	0
GRC	0	0	2012	0
HRV	0	0	1998	0
HRV	0	0	2003	0
HRV	0	0	2015	0
HUN	0	0	1995	0
HUN	0	0	2009	1
IDN	1	1	1984	0
IDN	1	1	1998	1
ISL	0	1	2001	1
ISL	0	1	2009	1
ITA	0	0	1983	1
ITA	0	0	1993	0
ITA	0	0	2012	0
JPN	0	0	2015	0
KOR	1	0	1998	1
LBN	1	0	1990	0
LBN	1	0	1992	0
LBN	1	0	1997	0
LBN	1	0	2000	1
LBN	1	0	2004	0
MAR	1	0	1983	1
MAR	1	0	1997	1
MAR	1	0	2003	0

Continued on next page

Country	Developing	Commodity exporter	Year	Systemic SS episode
MEX	1	0	1983	1
MEX	1	0	1987	0
MEX	1	0	1995	1
MYS	1	0	1994	0
MYS	1	0	1998	1
NGA	1	1	1984	1
NGA	1	1	1992	0
NGA	1	1	1995	0
NGA	1	1	2000	0
NGA	1	1	2005	0
NLD	0	0	1981	0
NLD	0	0	2003	1
NLD	0	0	2009	0
NOR	0	1	1990	1
NOR	0	1	1999	1
NOR	0	1	2008	0
NZL	0	1	2009	0
PAK	1	0	1997	0
PAK	1	0	2002	0
PAK	1	0	2009	1
PAN	1	0	1995	0
PAN	1	0	1998	0
PAN	1	0	2009	0
PER	1	1	1983	1
PER	1	1	1985	0
PER	1	1	1998	1
PHL	1	0	1984	0
PHL	1	0	1998	1
PHL	1	0	2009	1
POL	0	0	1999	0
PRT	0	0	1982	1
PRT	0	0	2012	0
RUS	1	1	1998	1

Continued on next page

Country	Developing	Commodity exporter	Year	Systemic SS episode
RUS	1	1	1999	0
SLV	1	0	1982	1
SLV	1	0	1999	0
SWE	0	0	1983	0
SWE	0	0	1994	0
SWE	0	0	2001	1
SWE	0	0	2003	0
SWE	0	0	2006	0
THA	1	0	1998	1
TUN	1	0	1982	1
TUN	1	0	2007	0
TUN	1	0	2016	0
TUR	1	0	1988	0
TUR	1	0	1994	1
TUR	1	0	1999	1
TUR	1	0	2001	0
UKR	1	1	1999	0
URY	1	1	1982	0
URY	1	1	1984	1
URY	1	1	2009	1
USA	0	0	1991	1
USA	0	0	2009	1
VEN	1	1	1983	1
VEN	1	1	1990	0
VEN	1	1	1996	0
ZAF	1	1	1983	1
ZAF	1	1	1985	0
Total episode-years			135	51

Notes: This table reports the episode level classification used in the paper. Each row is a country-year sudden stop episode. *Developing* equals one for developing/emerging economies and zero for developed economies. *Commodity exporter* equals one for commodity exporting economies and zero otherwise. *Systemic SS episode* equals one when the episode also satisfies the Bianchi–Mendoza global financial stress filter. The reconciled chronology therefore contains 135 broad episodes and 51 systemic episodes; 50 of the latter have the exact pre-crisis commodity dependence and income observations required by the baseline local projection. Non-systemic episodes enter only the descriptive evidence in Table 1 and Appendix Table A2. After the final classification reconciliation, the descriptive cells contain 45 developed non-commodity exporter episodes, 9 developed commodity exporter episodes, 40 developing non-commodity exporter episodes, and 41 developing commodity exporter episodes.

A.1 Median Event Path Statistics

Table A2 repeats the descriptive exercise in Table 1 using the median, rather than the mean, across available episodes at each relative year. The within group ordering is unchanged for every reported outcome: commodity exporters experience larger current account reversals and deeper declines in GDP, consumption, investment, and the real exchange rate in both development groups. The median exercise therefore shows that the descriptive ordering is not specific to the mean event path.

Table A2: Peak to Trough Changes from Median Event Paths by Commodity Exporter and Development Status

Commodity exporter	Developing	Episodes	GDP	Consumption	Investment	RER	CA/GDP
0	0	45	-2.52	-2.69	-8.95	-1.67	2.22
1	0	9	-3.82	-3.56	-19.24	-3.09	4.22
0	1	40	-3.11	-2.59	-11.99	-3.08	2.56
1	1	41	-4.46	-5.66	-19.39	-19.17	4.19

Notes: The construction mirrors Table 1, except that the median across available episodes is computed at each event time before the peak to trough statistic is formed. The episode classification, full episode exclusions, and treatment of outcome specific missing values are identical to Table 1. Negative RER values denote real depreciations; CA/GDP is reported in percentage points.

A.2 Data Sources and Variable Construction

Table A3 consolidates the provenance and construction of the variables used in the main empirical analysis and its robustness exercises. The final empirical code constructs all pre-crisis characteristics using exact calendar year lags rather than row lags, so a missing calendar year does not silently substitute for $t-1$. The baseline uses commodity dependence and real GDP per capita measured in the exact year before onset; private credit and institutional variables enter only in robustness exercises. Credit is intentionally taken from one WDI definition throughout and is not backfilled with alternative credit measures from auxiliary files.

Table A3: Data Sources and Variable Construction

Variable / object	Source	Definition, timing, and use
Systemic sudden stop onset, SS_{it}	Bianchi and Mendoza (2020); empirical input file `SS episodes Bianchi.csv`.	The event chronology is inherited rather than reestimated. Bianchi–Mendoza identify candidate events from unusually large annual increases in CA/GDP relative to each country's own historical distribution and classify systemic episodes using EMBI/VIX financial stress filters. $SS_{it} = 1$ in the onset year used by the LP.
Macroeconomic outcomes	Clean macro panel assembled from the Global Macro Database (Müller et al. (2025)), World Bank WDI, and IMF IFS, according to country/series availability.	Real GDP per capita, real consumption per capita, real investment, CA/GDP, RER, and real stock prices; annual 1980–2016 for the sudden stop analysis. GDP, consumption, and investment are log HP cycle measures with $\lambda = 100$; CA/GDP is HP filtered in levels. The final LP code constructs exact calendar changes relative to $t-1$; RER and stock prices use log differences.
Commodity dependence, $CD_{i,t-1}$	WDI: `TX.VAL.FOOD.ZS.UN`, `TX.VAL.AGRI.ZS.UN`, `TX.VAL.FUEL.ZS.UN`, `TX.VAL.MMTL.ZS.UN`.	Sum of food, agricultural raw materials, fuel, and ores and metals exports, each as a share of merchandise exports. Measured in exact calendar $t-1$. The aggregate is left missing unless all four component shares are observed. This is the continuous baseline exposure measure.
Real GDP per capita	WDI, `NY.GDP.PCAP.KD`.	GDP per capita in constant 2015 U.S. dollars, measured in exact calendar $t-1$ and logged in the regression. It allows sudden stop severity to vary with pre-crisis development.

Continued on next page

Variable / object	Source	Definition, timing, and use
Commodity export price, x_{it}	IMF, *Commodity Export Price Index, Individual Commodities Weighted by Ratio of Exports to GDP, Historical Fixed Weights*.	Country specific export price index. The final code divides the index by 100, takes logs, and constructs the exact one year change $\Delta \log x_{it}$. Both $\Delta \log x_{it}$ and $SS_{it} \times \Delta \log x_{it}$ enter the baseline.
Private credit / GDP	WDI, `FD.AST.PRVT.GD.ZS`.	Domestic credit to the private sector by banks as a share of GDP, exact calendar $t-1$. Robustness only. The final code uses this definition exclusively and does not backfill missing observations with alternative credit measures.
Capital account openness	Chinn and Ito (2008).	Chinn–Ito capital account openness index, exact calendar $t-1$. Robustness only; its sudden stop interaction is added in the institutional control specification.
Exchange rate regime	Ilzetzki et al. (2022).	Exact calendar $t-1$. The final code treats regime codes 1–6 as valid and defines $ER^{fixed} = 1\{\text{regime code} = 1\}$; codes 2–6 are zero. Robustness only.
World real short rate	FRED: U.S. 3-Month Treasury Bill Secondary Market Rate, Discount Basis, and U.S. CPI inflation.	Constructed as the nominal three month Treasury bill rate minus U.S. CPI inflation. The robustness specification uses the exact annual change ΔR_t^{world} interacted with SS_{it}; the common level is absorbed by year fixed effects.

Continued on next page

Variable / object	Source	Definition, timing, and use
Binary commodity exporter status	World Bank, *Global Economic Prospects*, January 2022.	Static 2017–2019 classification: commodity exporter if commodities average at least 30% of total exports or a single commodity accounts for at least 20%. Used only for the descriptive CE/NCE comparisons in Table 1 and Appendix Table A2, not as the main historical exposure measure.
Developing / developed split	United Nations development classification in the country classification file.	The Bianchi 58 country risk set is partitioned into developing/emerging and developed economies using the same classification as the descriptive table. The pooled continuous specification allows severity to vary directly with pre-crisis real GDP per capita.

Notes: The empirical LP is implemented on the 58 country Bianchi risk set. Exact calendar construction is used throughout: a value at $t-1$ is available only when the immediately preceding calendar year is observed. The final code scales the commodity price change by 100 for numerical readability, which does not affect the exposure coefficient. Country and year fixed effects are included in the baseline; standard errors are clustered by country with the finite sample correction. The source hierarchy for CD and GDP per capita fills missing cells from configured annual files, while the bank credit measure is deliberately restricted to the single WDI bank credit series to preserve a homogeneous definition.

B Additional Local Projection Evidence

This appendix reports the diagnostics and robustness exercises underlying Section 3. All continuous exposure specifications use systemic sudden stops only. Unless stated otherwise, the specification is the pooled baseline in equation (2) and does not include Credit/GDP.

B.1 Pre-trend tests

Table B1 reports country cluster bootstrap tests of the joint null that the $CD_{75} - CD_{25}$ differential equals zero at $h = -3$ and $h = -2$. None of the six outcomes rejects the null at conventional levels.

Table B1: Joint Pre-trend Tests for the Continuous Exposure Gradient

Outcome	Wald statistic	Bootstrap p-value	Valid replications
GDP	1.063	0.588	999
Consumption	2.047	0.359	999
Investment	2.145	0.342	999
CA/GDP	1.769	0.413	999
RER	4.126	0.127	999
Stock prices	1.523	0.467	999

Notes: Country cluster bootstrap test of $H_0 : \Delta_{-3}^{75-25} = \Delta_{-2}^{75-25} = 0$, with $B = 999$ valid replications for each outcome.

B.2 Common sample robustness

Table B2 separates changes in controls from changes in sample composition. Within each block, the baseline and augmented regressions use exactly the same observations.

Table B2: Common Sample Robustness at $h = 1$

	Credit sample		Institutional sample		World rate sample	
Outcome	Baseline	+ Credit	Baseline	+ Credit + KA + ER	Baseline	+ World rate
GDP	+0.24	+0.04	+0.24	−0.26	−0.15	−0.15
Consumption	-2.39^{*}	-2.37^{*}	-2.41^{*}	-2.73^{**}	-2.45^{**}	-2.40^{**}
Investment	−6.21	-6.89^{*}	−6.18	-7.76^{**}	−6.14	−6.10
CA/GDP	$+3.19^{**}$	$+3.38^{***}$	$+3.22^{**}$	$+3.34^{***}$	$+2.93^{**}$	$+2.93^{**}$
RER	−0.044	−0.031	−0.044	−0.026	−0.061	−0.062
Stock prices	-0.446^{**}	-0.324^{**}	-0.449^{**}	-0.356^{***}	-0.413^{**}	-0.452^{**}

Notes: Each entry is the $CD_{75} - CD_{25}$ differential at $h = 1$. Within each block, the two columns use the same observations. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

B.3 Sensitivity to commodity price conditioning

Table B3 reestimates the continuous exposure regression on the exact baseline sample after removing the contemporaneous commodity price terms.

Table B3: Sensitivity to Commodity Price Conditioning at $h = 1$

Outcome	Baseline	No price conditioning
GDP	-0.15	-0.06
Consumption	-2.45^{**}	-1.90^{*}
Investment	-6.14	-5.08
CA/GDP	$+2.93^{**}$	$+2.60^{**}$
RER	-0.061	-0.051
Stock prices	-0.413^{**}	-0.413^{**}

B.4 Overall trade intensity

Commodity dependence measures the composition of merchandise exports rather than the size of the external sector. Table B4 therefore adds exact $t-1$ Trade/GDP and its interaction with sudden stop onset. The first column reestimates the baseline on exactly the same observations used by the augmented specification. The one year current account gradient rises from 2.76 to 3.59 percentage points after adding trade intensity, while the consumption, investment, and equity price differentials remain negative.

Table B4: Trade Intensity Robustness at $h = 1$

Outcome	Baseline, common sample	+ Trade/GDP
GDP	-0.69	-1.00
Consumption	-2.35^{**}	-2.58^{**}
Investment	-5.88	-7.98^{**}
CA/GDP	$+2.76^{**}$	$+3.59^{***}$
RER	-0.088^{*}	-0.081^{*}
Stock prices	-0.395^{**}	-0.443^{***}

Notes: Entries report the $CD_{75} - CD_{25}$ differential at $h = 1$. Both columns use exactly the same observations. The augmented specification adds exact $t-1$ Trade/GDP and $SS_{it} \times Trade/GDP_{i,t-1}$ to the pooled baseline. Trade/GDP is an overall trade intensity measure and is used here to distinguish export composition from the scale of cross border trade. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

B.5 Exact pre-crisis versus static country characteristics

The baseline measures commodity dependence and real GDP per capita in the calendar year immediately preceding each sudden stop. As a timing sensitivity exercise, I instead characterize each country using its average commodity dependence and real GDP per capita over 2021–2023 and assign those static country characteristics to all of its historical observations. Because these averages postdate many of the sudden stops in the sample, I do not interpret this specification as an alternative pre-crisis design. Its purpose is only to ask whether the exposure ordering depends on measuring country characteristics in the year immediately preceding each episode.

To separate the change in measurement from changes in sample composition, Table B5 estimates the exact $t-1$ and static specifications on the same observations. The results are very similar. At $h = 1$, the current-account differential is +2.93 percentage points in the preferred exact-$t-1$ specification and +3.31 percentage points using the static 2021–2023 characteristics. Consumption, investment, the real exchange rate, and stock prices also move in the same direction under both measurement schemes.

Table B5: Pre-crisis versus Static Characteristics on the Same Observations

Outcome	Exact $t-1$	Static 2021–23
GDP	-0.15	-0.58
Consumption	-2.45^{**}	-3.04^{***}
Investment	-6.14	-7.33^{*}
CA/GDP	$+2.93^{**}$	$+3.31^{***}$
RER	-0.061	-0.087
Stock prices	-0.413^{**}	-0.465^{**}

Notes: Entries report the $CD_{75} - CD_{25}$ differential at $h = 1$. Both specifications use exactly the same observations and retain the baseline commodity price conditioning. The first column uses exact calendar $t-1$ commodity dependence and real GDP per capita. The second replaces these variables with country averages over 2021–2023, replicated across that country's historical observations. Country fixed effects absorb the time invariant main effects of the static characteristics. The exact $t-1$ specification remains preferred because the static 2021–2023 characteristics postdate many historical crises. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

B.6 Centered commodity price interaction

The baseline conditions on the realized commodity price movement but does not require the commodity exposure gradient itself to vary with that movement. Table B6 relaxes this restriction. I center $100\Delta \log x_{it}$ at its mean among systemic sudden stops, -4.83, and include the full interaction hierarchy $SS_{it} \times CD_{i,t-1} \times \widetilde{\Delta \log x_{it}}$. The first numeric column therefore reports the $CD_{75} - CD_{25}$ differential at the mean systemic event price realization, while the second reports how that differential changes with a one percentage point higher commodity price movement.

Table B6: Centered Commodity Price Interaction at $h = 1$

Outcome	Exposure gradient at mean SS price move	Triple interaction slope
GDP	-0.20	-0.099
Consumption	-2.33^{**}	-0.021
Investment	-5.97	-0.028
CA/GDP	$+3.10^{**}$	$+0.014$
RER	-0.057	-0.006
Stock prices	-0.347^{*}	$+0.007$

Notes: The commodity price movement is centered at the mean among systemic sudden stop observations, $100\Delta \log x = -4.83$. The first column reports the $CD_{75} - CD_{25}$ differential at that price realization. The second column is the change in that differential associated with a one percentage point higher value of $100\Delta \log x$. For CA/GDP, the exposure gradient p-value is 0.012 and the triple interaction p-value is 0.905. The specification is a robustness diagnostic for whether the average exposure gradient is sensitive to the realized commodity price movement. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

B.7 Coverage of additional controls

The baseline continuous specification has 50 systemic sudden stop episodes with the required pre-crisis commodity dependence and income characteristics. Requiring Credit/GDP reduces this set to 45 episodes. After retaining all valid exchange rate regime codes, adding capital account openness and the exchange rate regime does not reduce the episode count further.

Table B7: Coverage of Additional Pre-crisis Controls

Specification requirement	Eligible systemic episodes
Commodity dependence + GDPpc + price conditioning	50
+ Credit/GDP $t-1$	45
+ Credit + KA openness + ER regime $t-1$	45

C Model Exposure Mapping and Additional Quantitative Details

C.1 Regression bridge from commodity dependence to α

The final mapping uses the full set of systemic sudden stop episodes for which the three sectoral value added components in equation (28) can be constructed from a homogeneous national accounts observation. All sectoral inputs come from a single source, the United Nations National Accounts Official Country Data, Table 2.1. For each country and pre-crisis year, agriculture, mining, and manufacturing must be observed under the same Series, currency, SNA system, and fiscal year convention. When multiple internally homogeneous triplets are available, the baseline selects the newest SNA system and then the largest numeric Series code. Of the 50 systemic episodes with the empirical exposure and income variables required by the baseline local projection, 46 episodes from 34 countries have a usable homogeneous Table 2.1 triplet.

I estimate the descriptive bridge in equation (29) on those 46 episodes. Table C1 reports the regression and the implied interquartile mapping.

Table C1: Regression Mapping from Commodity Dependence into Model Exposure

Object	Estimate	Country clustered SE
Intercept, a	0.236	0.043
Slope on $CD/100$, b	0.278	0.072
Mapped α at $CD_{25} = 22.5\%$	0.298	0.031
Mapped α at $CD_{75} = 72.9\%$	0.438	0.029
High minus low mapped α	0.140	0.036
Observations	46 episodes	
Countries	34	
R^2	0.270	
Pearson correlation	0.519	
Spearman correlation	0.520	

Notes: The dependent variable is $\alpha^{VA} = (VA^{Agr} + VA^{Mining})/(VA^{Agr} + VA^{Mining} + VA^{Manufacturing})$ in the calendar year preceding each systemic sudden stop. Commodity dependence CD is the pre-crisis commodity export share used in the empirical analysis. The regression is $\alpha^{VA} = a + b(CD/100) + u$. Standard errors are clustered by country. The main quantitative exercises use the rounded endpoints $\alpha_L = 0.30$ and $\alpha_H = 0.44$.

The fitted endpoints are insensitive to reasonable choices among alternative homogeneous UN series. Table C2 varies the rule used when more than one observation is available within the preferred SNA vintage.

Table C2: Sensitivity of the Commodity Exposure Mapping to Series Selection

Series selection rule	$\widehat{\alpha}_L$	$\widehat{\alpha}_H$	High – Low
Baseline: newest SNA, largest Series	0.298	0.438	0.140
Minimum α^{VA} within preferred SNA	0.293	0.437	0.144
Maximum α^{VA} within preferred SNA	0.296	0.455	0.159
Mean α^{VA} within preferred SNA	0.294	0.446	0.152

Notes: Each row reestimates the bridge regression on the same 46 episodes and 34 countries after changing only the rule for selecting among internally homogeneous Table 2.1 observations. The implied low endpoint remains approximately 0.29–0.30 and the high endpoint approximately 0.44–0.46.

Because the final mapping uses one harmonized national accounts source rather than country specific reports, a separate country by country source table is unnecessary. The replication files retain the episode level sectoral values, Series identifiers, currencies, SNA systems, fiscal year conventions, and the alternative selection diagnostics.

C.2 Quantitative sensitivity to the mapping

The rounded benchmark $(\alpha_L, \alpha_H) = (0.30, 0.44)$ is numerically indistinguishable from using the exact fitted endpoints $(0.2985, 0.4383)$: the matched state high minus low difference in $\Delta(CA/Y)$ is 4.884 percentage points under the rounded calibration and 4.875 percentage points under the exact fitted values. Across the alternative series selection rules in Table C2, the corresponding current account differential lies between approximately 5.01 and 5.58 percentage points. In every case, the high exposure economy also experiences a larger consumption contraction, real depreciation, fall in the relative price of nontradables, and decline in collateral relevant income. These exercises are mapping sensitivity checks rather than alternative calibrations selected to improve model fit.

C.3 Stochastic incidence and event overlap

The stochastic equilibrium changes both balance sheets and the set of dates classified as sudden stops. Figure C1 reports the resulting incidence and the overlap between the zero exposure crisis dates and those at higher exposure. These objects complement the fixed event comparison in Section 7.4: they show why event composition itself becomes endogenous when households adjust debt to structural exposure.

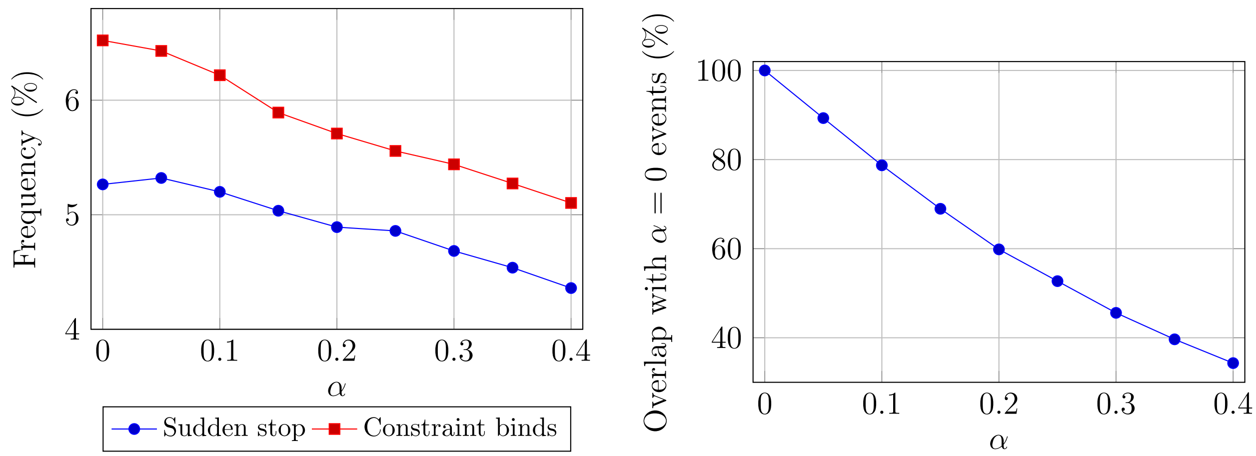

Figure C1: **Endogenous crisis incidence and event overlap.** The left panel reports sudden stop and collateral binding frequencies in the decentralized stochastic equilibrium. The right panel reports the share of benchmark $\alpha = 0$ sudden stop dates that are also sudden stops at each exposure level.

C.4 Borrowing rule decomposition: event definition robustness

Table C3 verifies that the central attenuation result does not depend on the country specific 95th percentile candidate event rule used in the main decomposition. The magnitude varies because different rules select different parts of the simulated crisis distribution, but the equilibrium current account gradient remains close to zero under every definition and rises substantially when the differential precautionary borrowing response is removed.

Table C3: Borrowing Rule Decomposition across Event Definitions

Event definition	Equilibrium	No differential adjustment
Country specific P95 (baseline)	+0.01	+2.63
Pooled P95	+0.11	+1.48
Country specific mean $+2sd$	+0.16	+3.11
Bianchi model rule	+0.17	+1.88
Pooled P95, random 50 event subsample	−0.01	+1.32
Country specific P95 + binding	+0.13	+2.54

Notes: Entries are high minus low current account differentials in percentage points using the linear average $\bar{\beta}_{0:1}$ from equation (50). The empirical counterpart is 1.885 percentage points. Results average 200 Monte Carlo pseudo panels. The country specific P95 rule is the preferred model to data candidate event analogue. The P95 + binding rule deliberately conditions on the model's collateral mechanism and is reported only as a structural diagnostic.