Monetary Shocks, Currency Exposures, and the Cost of Sovereign Default

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Abstract:
How costly is sovereign default? I develop a probabilistic sovereign default model that features (i) foreign monetary shocks that induce self-fulfilling default equilibria; (ii) multiple equilibria that imply a local average treatment effect; and (iii) under Fréchet heterogeneity in nominal exchange rates, default probability admits a shift-share representation. Guided by these insights, I exploit aggregate variation in developing countries' currency denomination of external debt (endogenous shares) and advanced economies' quasi-random interest rate movements (exogenous shifts) to construct a shift-share instrumental variable (SSIV) for sovereign default decisions. Using a local projection–instrumental variable (LP-IV) approach, I causally estimate that sovereign defaults on average result in an 8% decline in real GDP per capita in the first year. The impact peaks at 18.5% around the second year, persists until the fourth year, and then fades toward zero by the sixth year. Moreover, I find that floating exchange rate regimes and lower external debt levels, especially short-term debt, effectively attenuate the output loss. Narrative monetary shocks and difference-in-difference analyses yield similar results, further confirming that sovereign default is indeed costly.

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1. Introduction

How costly is sovereign default? The seminal Eaton and Gersovitz (1981) model, which posits that defaulting countries lose access to cheaper foreign capital, cannot explain the high levels of external debt observed prior to an actual default. To address this discrepancy, subsequent research has incorporated direct output losses following default episodes to better align theoretical models with empirical data (Arellano 2008; Yue 2010; Mendoza and Yue 2012; Uribe and Schmitt-Grohe 2017). In reality, however, if a country can default without facing enforceable penalties, it might actually benefit from a debt burden being written off overnight. Therefore, the literature still lacks a credible causal estimate that identifies both the direction and the magnitude of default-induced output losses.

In this paper, I employ a novel empirical strategy—local projection with shift-share instrumental variable (LP-SSIV)—to estimate the causal effect of sovereign default on output loss. To be specific, I leverage developing countries' currency denomination of external debt as endogenous shares and the quasi-random interest rate movements in advanced economies as exogenous shifts to construct a shift-share instrument for sovereign default decisions. Building on the methodology of Jordà et al. (2020), I apply this shift-share instrument within a local projection–instrumental variable (LP-IV) framework to causally estimate the cost of default. The key mechanism works as follows: when an advanced economy raises its interest rates, its currency appreciates, making it more difficult for developing countries to repay their external debts denominated in that currency. This, in turn, increases the default probability for countries with relatively more external debt denominated in that appreciating foreign currency.

The 1980s Latin American Debt Crisis illustrates this mechanism vividly. Following the recycling of petrodollars during the 1970s oil crisis, many Latin American countries accumulated significant debts denominated in U.S. dollars. In 1979, U.S. interest rates unexpectedly surged under Paul Volcker's tightened monetary policy, causing the dollar to appreciate sharply, which then substantially increased the debt service cost for these countries. This led to widespread defaults, starting with Mexico in 1982, together with a "lost decade" of both sluggish economic growth and prolonged debt negotiations.

To preview the main results, the baseline LP-SSIV regressions show that on average, defaulting on external debt leads to an 8% output loss in the first year. The cumulative output loss peaks at around 18% in the second year and persists until the fourth year before gradually diminishing to zero by the sixth year, suggesting non-persistent impact. After accounting for extensive margin, traditional binary default indicators and continuous arrears-based measures of partial default deliver similar causal estimates. These results are robust to a variety of macroeconomic control variables commonly used in the literature. Furthermore, to mitigate spurious regressions in panel IV settings, I use long-difference and/or first-difference transformations in all of the specifications (Christian and Barrett 2024). The results are also robust to concerns about lead-lag exogeneity and incomplete shares highlighted by recent applied econometric literature (Borusyak et al. 2022; Stock and Watson 2018). Notably, the above findings are consistent across both narrative monetary shocks and/or changes in base countries' interest rates, further validating the significant cost of sovereign default.

Defending the exclusion restriction assumption is a central challenge for instrumental variable design. To address potential violations of this assumption, I adopt a control function approach within the LP-IV framework (Jordà et al. 2020; Wooldridge 2015; Conley et al. 2012). To be specific, I use never-defaulting countries as a control subsample to estimate the indirect effects of the shift-share instrument on cumulative output growth. This method delivers informative bounds on the causal estimates. In addition, assuming that the negative impacts of defaults dissipate over a six years window, I apply a local projection—difference-in-difference (LP-DiD) method, which is robust to spillovers and address negative-weight issues under staggered treatment—some countries have repeatedly defaulted multiple times (e.g., Argentina)—to provide an alternative set of causal estimates (Dube et al. 2025). Both approaches indicate that positive spillovers substantially attenuate the true default costs, consistent with the economic intuition that confounding factors, such as export boom following post-default depreciation (i.e., Twin Ds), can help reduce measured default cost.

To rigorously justify the SSIV approach, I build upon the seminal Cole and Kehoe (2000) multiple equilibria model and recast it as a probabilistic one. The corner solutions in the multiple equilibrium setting naturally implies a local average treatment effect (LATE): the estimated causal effect on output loss applies only to compliers at the margin of default. In addition, I follow the seminal Eaton and Kortum (2002) probabilistic trade model to examine frictions in the sovereign debt market that hinder rapid rebalancing of debt portfolios. These assumptions yield a shift-share representation of default probability, which underpins the empirical design.

Last but not least, within a state-dependent local projection framework, I find that deeply indebted countries—especially those owing a significant proportion of short-term debt (i.e., with maturities of less than one year)—face disproportionately larger default costs. Countries with pegged exchange rate regimes also experience more severe consequences, consistent with the limited ability to benefit from post-default depreciation. Interestingly, defaulting despite having adequate foreign exchange reserves—sufficient to cover three months of imports—is associated with higher default cost, possibly because it signals unwillingness to repay despite capacity.

1.1 Related Literature

This paper builds on previous efforts to estimate the cost of sovereign default on output loss. While there have been some structural attempts, such as trend-deviation approaches (Reinhart and Rogoff 2011; Tomz and Wright 2007), the computational complexity involved in solving sovereign default models has limited their prevalence. Much of the existing literature relies on panel fixed effects regressions to assess the static impact of default on output loss, though these approaches often face endogeneity issues (De Paoli et al. 2009; Borensztein and Panizza 2008; Levy-Yeyati and Panizza 2001). More recently, local projection (LP) methods have become increasingly popular for estimating dynamic impulse responses due to their flexibility and robustness, especially when combined with other techniques such as the generalized method of moments (LP-GMM) and inverse propensity score weighting (LP-IPSWRA) (Jordà 2005, 2024; Furceri and Zdzienicka 2012; Kuvshinov and Zimmerman 2019). While the literature generally estimates the first-year output loss from a default episode to range from 2% to 10% of real GDP per capita, it has not yet established a clear direction of causality.

The recent paper by Farah-Yacoub et al. (2024) is the closest counterpart to my LP-SSIV approach. They use a combination of local projection and synthetic controls to estimate the long-term effects of sovereign default on output loss and other social variables based on comprehensive historical data. Notably, my first-year estimate (-8%) is closely aligned with what they have found (-8.50% within three years). At first glance, there is a notable discrepancy regarding the persistence of the output cost: their results suggest a persistent negative impact of around 20% even after a decade, whereas my impulse response shows that the effects dissipate after approximately six years. This difference closes after accounting for spillover effects: the spillover-corrected IV estimate derived from the control function approach and/or LP-DiD estimates, which approximate a −20% to −30% output loss, align almost perfectly with their results.

This paper also builds on the expanding applied literature that leverages the shift-share instrumental variable and difference-in-difference approaches to estimate causal effects in fields such as international trade and immigration (Autor et al., 2013; Peri and Sparber 2009). While these methods have a long history in economic research, their application to sovereign default—especially with evolving currency shares and repeated defaults—illustrates how recent advances in their theoretical foundations can be applied in practice (Adão et al. 2019; Borusyak et al. 2022; Goldsmith-Pinkham 2020; Borusyak and Hull 2024; de Chaisemartin and D'Haultfœuille 2020; Roth et al. 2023; Sun and Abraham 2021; Borusyak et al. 2024; Borusyak et al. 2025). Given the increasing focus on causal inference in macroeconomic research, this paper is, to the best of my knowledge, the first to integrate these recent developments to answer a macroeconomic question.

Lastly, this paper closely relates to the "original sin" literature—why do developing countries mostly borrow in foreign currencies? (Eichengreen and Hausmann 1999). On the theoretical side, Coppola et al. (2025) and Eren and Malamud (2022) develop models to explain currency choice of external debt, highlighting determinants (e.g., search frictions) of endogenous currency exposure. In a related historical setting, Bordo and Meissner (2023) study how gold-clause debt shaped the staggered exit from the gold standard during the Great Depression. They use the September 1931 sterling devaluation as a natural experiment and find that staying on the gold standard reduced borrowing costs (i.e., lowered bond yields) in the short run, particularly for countries with a larger share of gold-denominated debt. On the other hand, Hébert and Schreger (2017) exploit the legal ruling in NML Capital, Ltd. v. Republic of Argentina in 2001 as a natural experiment to causally identify equity value declines around Argentina's default. Whereas these papers emphasize financial costs, my analysis focuses on real outcomes (i.e., real GDP per capita). Quantifying the contribution of the financial channel to the overall cost of sovereign default—and comparing it to potential gains from trade through currency depreciation or debt relief—remains an important avenue for future research.

The rest of the paper is organized as follows: Section 2 develops a novel probabilistic sovereign default model that motivates the key mechanism underpinning the empirical strategy. Section 3 describes the data. Section 4 outlines the empirical strategy. Section 5 presents the baseline LP-OLS and LP-SSIV results, together with the control function approach to address spillovers. Section 6 presents robustness checks, including narrative monetary shocks and difference-in-difference estimates. Section 7 discusses the baseline results and benchmarks them against historical crises. Section 8 examines the state-dependent heterogeneous output cost of sovereign default, and Section 9 concludes.

2. A Probabilistic Sovereign Default Model

This section develops a simple probabilistic sovereign default model based on Romer (2019) and Cole and Kehoe (2000). A sovereign debtor i borrows in dollar-denominated debt, faces stochastic repayment obligations as exogenous foreign monetary policy influences exchange rate movements, and must service this debt out of a stochastic domestic fiscal capacity earned in pesos. Not only does a contractionary foreign monetary shock raise the borrowing cost, it also leads to a higher perceived default probability because of currency mismatch (i.e., the debtor i earns pesos but must repay in appreciated dollars). Assuming creditors have adaptive expectations, these two mechanisms feed into the rollover rates: a higher perceived default risk for the next period translates into a higher rollover rate, which in turn raises the subsequent default likelihood. This recursive feedback between the default probability and risk premium—linked by adaptive expectations—can drive the economy into a self-fulfilling equilibrium.

How is this probabilistic default model useful for the empirical strategy? First, Section 2.2 shows that the multiple equilibria property in this model is directly related to the local average treatment effect (LATE) interpretation—the instrumental variable approach only estimates the causal effects on countries that have higher default likelihood because of foreign interest rate hikes (i.e., the "complier" effects). In other words, these causal estimates cannot speak to the output cost of the 2012 Greek debt crisis, as this episode occurred during the monetary easing in advanced economies after the global financial crisis. Second, motivated by extreme depreciation episodes observed in developing countries, Section 2.3 shows that, by imposing a Fréchet distribution of nominal exchange rate deprecation, the default probability admits a shift-share representation. In particular, the increase in default likelihood can be expressed as a linear approximation of the latent default probability with respect to foreign monetary shocks weighted by endogenous shares, providing a theoretical justification for using currency denomination shares to measure the exposure to foreign monetary shocks.

2.1 Environment and Timing

The model admits discrete time with three periods $t,\ t+1,\ t+2$, and a steady state period t_0. Normalizing the initial funding needed in an arbitrary period t to one peso, a representative debtor country i can borrow $\frac{1}{E_t^{ik}}$ dollar-equivalent of this 1-peso loan, where E_t^{ik} is the known nominal exchange rate (pesos per dollar) at t, in currency k and at an initial gross interest factor R_t^{ik} (determined at t and due at $t+1$). [2] When the stochastic counterpart E_{t+1}^{ik} is realized in period $t+1$, the peso-equivalent repayment for this dollar-denominated debt due becomes $\frac{1}{E_t^{ik}} R_t^{ik} E_{t+1}^{ik} = R_t^{ik} \mathcal{E}_{t+1}^{ik}$, where $\mathcal{E}_{t+1}^{ik} \equiv \frac{E_{t+1}^{ik}}{E_t^{ik}}$ is the gross depreciation factor with $\mathcal{E}_{t+1}^{ik} > 1$ indicating a peso depreciation. The debtor may negotiate a debt rollover from period $t+1$ to $t+2$, but any renewed contract is repriced at R_{t+1}^{ik} (determined at $t+1$ and due $t+2$) to reflect the default probability in period $t+2$ revealed at $t+1$.

2.2 A Simple Multiple Equilibria Model and Its Connection to LATE

2.2.1 Exchange Rate Block

Following the standard monetary approach to exchange rates, I assume the Purchasing Power Parity (PPP) and constant but country-specific real money demands. Under these conditions, the level of the bilateral nominal exchange rate satisfies:

$$E_t^{ik} = \frac{P_t^i}{P_t^k} = \frac{M_t^i}{M_t^k} \frac{\overline{L}^k}{\overline{L}^i}$$

Hence, the gross depreciation of peso against currency k from t to $t+1$ is:

$$\mathcal{E}_{t+1}^{ik} \equiv \frac{E_{t+1}^{ik}}{E_t^{ik}} = \frac{M_{t+1}^i \big/ M_{t+1}^k}{M_t^i \big/ M_t^k} \cdot \underbrace{\frac{\overline{L}^k \big/ \overline{L}^i}{\overline{L}^k \big/ \overline{L}^i}}_{=1} = \frac{\mathcal{M}_{t+1}^i}{\mathcal{M}_{t+1}^k} \propto \frac{1}{\mathcal{M}_{t+1}^k} \tag{2.1}$$

[2] While there is growing interest on local-currency denominated debt among emerging markets, this paper primarily examines foreign-currency external debt and therefore abstract from domestic-currency borrowing. The "original sin" literature—developing countries cannot borrow internationally in their own currency—provides justification for temporarily setting aside analysis on local currency denominated debts (Eichengreen and Hausmann 1999).

where $\mathcal{M}_{t+1}^{k} \equiv \frac{M_{t+1}^{k}}{M_{t}^{k}}$ is country k's gross money supply factor from t to $t+1$. For tractability, I shut down domestic monetary response and set $\mathcal{M}_{t+1}^{i} = 1$. The key mechanism is that a foreign contractionary monetary shock (i.e., $\mathcal{M}_{t+1}^{k} < 1$, equivalent to $\mathcal{M}_{t+1}^{k} - 1 = \frac{M_{t+1}^{k} - M_{t}^{k}}{M_{t}^{k}} = \%\Delta M_{t+1}^{k} < 0$) raises $\mathcal{E}_{t+1}^{ik}$, which in turn *depreciates* the peso, ceteris paribus.

2.2.2 Taxation/Revenue Block

Define $\mathcal{T}_{t+1}^{i} = \frac{T_{t+1}^{i}}{D_{t}^{i}}$ as the realized revenue-to-debt ratio for debtor i at $t+1$ and assume it will default whenever the realized revenue falls short of the *peso-equivalent* foreign currency obligation $R_{t}^{ik}\mathcal{E}_{t+1}^{ik}$:

$$\pi_{t+1}^{iG} \equiv \mathbb{P}\left(\mathcal{T}_{t+1}^{i} \leq R_{t}^{ik}\mathcal{E}_{t+1}^{ik}\right) \tag{2.2}$$

Therefore, given gross interest factor R_{t}^{ik}, a contractionary foreign monetary policy depreciates the nominal exchange rate, raising $\mathcal{E}_{t+1}^{ik}$ and therefore government default probability π_{t+1}^{iG} (exchange rate channel).

2.2.3 No-arbitrage Pricing by Risk-Neutral Foreign Investors

Given the government default condition in Equation (2.2), I close the model with a pricing equation that maps perceived default risk into the rollover rate. Suppose at $t+1$, the debtor i wants to roll over the $R_{t}^{ik}\mathcal{E}_{t+1}^{ik}$ peso-equivalent of dollar-denominated debt service. Let the model-consistent expectations for $t+2$ be $\overline{\pi}_{t+1}^{iG} \equiv \mathbb{E}_{t+1}\left[\pi_{t+2}^{iG}\right] = \mathbb{E}\left[\pi_{t+2}^{iG} \middle| \mathcal{F}_{t+1}\right]$

. If creditors have anchored adaptive expectations, then after observing π^{iG}_{t+1} in period $t+1$, their perceived default probability—formed at $t+1$—for $t+2$ is given as:

$$\pi^{iR}_{t+2|t+1} = \sigma\pi^{iG}_{t+1} + (1-\sigma)\overline{\pi}^{iG}_{t+1} \qquad \text{with} \quad \sigma \in [0,1] \tag{2.3}$$

Thus, when $\sigma \to 1$, the model features purely adaptive expectations; when $\sigma \to 0$, the model nests the rational expectations case.[3]

Then risk-neutral creditors choose the gross rollover factor R^{ik}_{t+1}—set at $t+1$ and due at $t+2$—such that expected return at $t+2$ equals the currency k's risk-free benchmark R^{fk}_{t+1}:

$$\pi^{iR}_{t+2|t+1}(0) + (1-\pi^{iR}_{t+2|t+1})R^{ik}_{t+1} = R^{fk}_{t+1} \tag{2.4}$$

where $\pi^{iR}_{t+2|t+1}$ is an exogenous variable (input) given by Equation (2.3), and R^{ik}_{t+1} is an endogenous variable (output). The above equation implies a risk premium $RP\left(\pi^{iR}_{t+2|t+1}\right)$ as $R^{ik}_{t+1} = \dfrac{R^{fk}_{t+1}}{1-\pi^{iR}_{t+2|t+1}} = R^{fk}_{t+1}\left(1 + RP(\pi^{iR}_{t+2|t+1})\right) > R^{fk}_{t+1}$, and we can show that:

$$RP(\pi^{iR}_{t+2|t+1}) = \frac{\pi^{iR}_{t+2|t+1}}{1-\pi^{iR}_{t+2|t+1}} \qquad \text{and} \qquad RP'(\cdot) = \frac{1}{\left(1-\pi^{iR}_{t+2|t+1}\right)^2} > 0 \tag{2.5}$$

which implies higher perceived default risk raises the rollover rate accordingly.

2.2.4 Partial Default

The specification in Section 2.2.3 extends to the partial default case: instead of all-or-nothing, the debtor i pledges to liquidate all tradable assets and repay in dollars upon

[3] If random variables $\left(\mathcal{T}^i_{t+1}, \mathcal{M}^k_{t+1}\right)$ are independently drawn across time, then rational expectations imply a constant steady state anchor as $\mathbb{E}_{t+1}\left[\pi^{iG}_{t+2}\right] = \mathbb{E}\left[\pi^{iG}_{t+2}\right] \equiv \pi^i_0$. In that case, Equation (2.3) becomes $\pi^{iR}_{t+2|t+1} = \sigma\pi^{iG}_{t+1} + (1-\sigma)\pi^i_0$.

default. Reinterpret $\pi^{iR}_{t+2|t+1} \in [0,1]$ as the expected share of loss given default at $t+2$, creditors expect to recover $1-\pi^{iR}_{t+2|t+1}$ of disbursed debt. Since payoffs are denominated in dollars, dollar recovery from pledged assets makes spot exchange rate movements irrelevant to creditors conditional on repayment or recovery. Pricing then equates expected dollar repayment on the surviving fraction of rollover debt to the borrowing cost factor $(1-\pi^{iR}_{t+2|t+1})R^{ik}_{t+1} = R^{fk}_{t+1}$, the same functional form as Equation (2.4). Therefore, in both full and partial default, a higher perceived default probability $\pi^{iR}_{t+2|t+1}$ raises the rollover rate R^{ik}_{t+1} (pricing channel).

2.2.5 General Equilibrium: Joint Determination of Default Risk and Rollover Premium

To highlight the feedback mechanism between the default probability $\pi^{i}_{t+h|t+h-1}$ and the risk premium $RP(\pi^{i}_{t+h|t+h-1})$, consider a parsimonious setting where the initial steady-state equilibrium $A = \left(\pi^{i}_{0}, R^{ik}_{0}\right)$ lies in the interior of the state space at period t_0. Suppose there is a one-period temporary foreign contractionary monetary shock $\mu_k > 0$ at period t, $\%\Delta E^{ik}_{t+1} = \mu_k > 0$, so the peso *depreciates.*[4] Unlike the baseline model, the shock is modeled as a single realization rather than a random process, but the revenue-to-debt $\mathcal{T}^{i}_{t+1}$ remains stochastic.

[4] From Equation (2.1) and $\mathcal{M}^{i}_{t+1}$, we can derive $\Delta\%E^{ik}_{t+1} \approx \ln \mathcal{E}^{ik}_{t+1} = -\ln \mathcal{M}^{k}_{t+1} \approx -\%\Delta M^{k}_{t+1}$. Define $\mu_k \equiv -\%\Delta M^{k}_{t+1}$, that is, $\mu_k > 0$ is a contractionary foreign monetary policy shock. For small shock, $\%\Delta E^{ik}_{t+1} = -\%\Delta M^{k}_{t+1} = \mu_k > 0$.

To illustrate the key mechanism, Panel (a) of Figure 1 shows the partial equilibrium effect of a temporary foreign contractionary monetary shock at period $t+1$. Since the peso-value of dollar-denominated obligations increases, $\pi^i > \pi_0^i$ for all R^{ik}, the government default curve shifts upward. Therefore, since the interest factor is predetermined, the equilibrium moves to B with $\pi_B^i > \pi_0^i$.

At the same time, suppose debtor i must roll over its entire debt service into $t+2$; the risk-neutral global investors, observing π_B^i, reprice the rollover debt by endogenously raising $R_C^{ik} > R_0^{ik}$. In other words, when investors observe higher default probability due to peso depreciation, they charge a higher risk premium; the next period's default probability further rises to $\pi_D^i > \pi_B^i > \pi_0^i$, and the risk premium subsequently increases to R_E^{ik}, further raising the peso-value obligation. The higher debt burden again elevates default probability to π_F^i, prompting an even larger premium on rollover debt R_G^{ik}. This feedback loop can continue and may drive the economy to a self-fulfilling default equilibrium.[5]

Not only does a contractionary foreign monetary policy inflate the peso value of dollar liabilities, which raise the default risk premium, but it also lifts the dollar borrowing cost—even holding π^i fixed—as tightened monetary policy reduces market liquidity and pushes up interest rate. Therefore, Panel (b) in Figure 1 shows the general equilibrium effect: the government default curve shifts leftward, and the investor return curve shifts rightward. As the government default probability rises with peso depreciation, the cost of rolling over debt also increases at the same time, resulting in a faster convergence from the initial equilibrium to the default equilibrium.

[5] The mechanism is the same under a permanent shock (path denoted by primes).

Formally we can also derive the comparative statics from the general equilibrium system: combining Equation (2.2), (2.3) and (2.4) we have

$$R_{t+1}^{ik} = \frac{R_{t+1}^{fk}}{1-\pi_{t+2|t+1}^{iR}} = \frac{\mathcal{R}\left(\mathcal{M}_{t+1}^{k}\right)}{1-\left(\sigma\pi_{t+1}^{iG}+(1-\sigma)\overline{\pi}_{t+1}^{iG}\right)} = \frac{\mathcal{R}\left(\mathcal{M}_{t+1}^{k}\right)}{1-\sigma F_T\left(R_t^{ik}\mathcal{E}\left(\mathcal{M}_{t+1}^{k}\right)\right)-(1-\sigma)\overline{\pi}_{t+1}^{iG}}$$

where $F_T(\cdot)$ is the cumulative distribution function of the random variable $\mathcal{T}_{t+1}^{i}$ and by definition, the probability density function $F_T'(\cdot) > 0$.

$$\frac{\partial R_{t+1}^{ik}}{\partial \mathcal{M}_{t+1}^{k}} = \left(\frac{1}{1-\sigma F_T(\cdot)-(1-\sigma)\overline{\pi}_{t+1}^{iG}}\right)\overbrace{\frac{\partial R_{t+1}^{fk}}{\partial \mathcal{M}_{t+1}^{k}}}^{<0} + \left(\frac{\sigma F_T'(\cdot) R_t^{ik} R_{t+1}^{fk}}{\left(1-\sigma F_T(\cdot)-(1-\sigma)\overline{\pi}_{t+1}^{iG}\right)^2}\right)\overbrace{\frac{\partial \mathcal{E}_{t+1}^{ik}}{\partial \mathcal{M}_{t+1}^{k}}}^{<0} < 0 \tag{2.6}$$

The above expression shows two channels through which a contractionary foreign monetary shock raises the rollover rate: (i) an increase in the borrowing cost in currency k (the direct effect); (ii) rising perceived default probability in the rollover period $t+2$ because of peso depreciation (the indirect effect). When $\sigma \to 0$, the Equation (2.6) simplifies to $\frac{\partial R_{t+1}^{ik}}{\partial \mathcal{M}_{t+1}^{k}} = \left(\frac{1}{1-\overline{\pi}_{t+1}^{iG}}\right)\overbrace{\frac{\partial R_{t+1}^{fk}}{\partial \mathcal{M}_{t+1}^{k}}}^{<0}$, where $\overline{\pi}_{t+1}^{i} \equiv \mathbb{E}_{t+1}\left[\pi_{t+2}^{iG}\right]$ is the rational expectation prior formed at $t+1$ for $t+2$. This underscores that expectations formation—anchored by policy credibility—determines whether the risk premium channel responds to the realized depreciation.

To derive the government default probability at period $t+2$, rewrite Equation (2.2) into $\pi_{t+2|t+1}^{iG} \equiv \mathbb{P}\left(\mathcal{T}_{t+2}^{i} \leq R_{t+1}^{ik}\mathcal{E}_{t+2}^{ik}\right)$. Assume a one-period foreign contractionary monetary shock such that $\mathcal{E}_{t+1}^{ik} > 0$ and $\mathcal{E}_{t+2}^{ik} = 1$ returns to 1, and that since $\mathcal{T}_{t}^{i}$ are independent and identically distributed, then we have:

$$\pi^{iG}_{t+2|t+1} \equiv F_T\left(R^{ik}_{t+1}\right) = F_T\left(\frac{\mathcal{R}\left(\mathcal{M}^k_{t+1}\right)}{1-\sigma F_T\left(R^{ik}_t \mathcal{E}\left(\mathcal{M}^k_{t+1}\right)\right) - (1-\sigma)\overline{\pi}^{iG}_{t+1}}\right)$$

Differentiating the above equation with respect to $\mathcal{M}^k_{t+1}$ gives:

$$\frac{\partial \pi^{iG}_{t+2|t+1}}{\partial \mathcal{M}^k_{t+1}} = F'_T(\cdot)\left(\frac{1}{1-\sigma F_T(\cdot) - (1-\sigma)\overline{\pi}^{iG}_{t+1}}\right)\overbrace{\frac{\partial R^{fk}_{t+1}}{\partial \mathcal{M}^k_{t+1}}}^{<0} + F'_T(\cdot)\left(\frac{\sigma R^{ik}_t R^{fk}_{t+1} F'_T(\cdot)}{\left(1-\sigma F_T(\cdot) - (1-\sigma)\overline{\pi}^{iG}_{t+1}\right)^2}\right)\overbrace{\frac{\partial \mathcal{E}^{ik}_{t+1}}{\partial \mathcal{M}^k_{t+1}}}^{<0} < 0$$

The above expression shows that a contractionary foreign monetary shock raises default probability through both higher borrowing costs and currency depreciation, the latter of which can be mitigated by anchoring expectations through strengthening credibility in developing countries.

This probabilistic framework thus reveals a self-fulfilling feedback loop absent from standard binary default models: a temporary monetary tightening abroad raises the peso-value of dollar-denominated debt; higher expected default risk pushes up rollover rates; the higher rates further worsen default odds. The 2011 euro-zone debt crisis provides a vivid illustration: as markets revised Greece's default probability upward, the bond yields surged, and the soaring risk premia made repayment increasingly untenable.

Two additional important implications follow from the above model: first, even under a pegged exchange rate regime, countries may still be pushed toward a default equilibrium, as an increase in the borrowing cost shifts the investor return curve rightward, pushing the economy to a default equilibrium (Panel (c) of Figure 1). Second, a natural question to ask is, if an interest rate shock always drives countries into default equilibrium, why haven't we seen more sovereign defaults in recent years. One explanation is that, after the 1980s Latin American Debt Crisis and 1997 Asian Financial Crisis, many emerging markets began to accumulate foreign exchange reserves, which have a macro-prudential effect by augmenting resources for servicing foreign currency denominated debt:

$$\pi_{t+1}^{iG} = \mathbb{P}(\mathcal{T}_{t+1}^{i} + FX^{ik} \cdot \mathcal{E}_{t+1}^{ik} \leq R_t^{ik}\mathcal{E}_{t+1}^{ik}) = \mathbb{P}(\mathcal{T}_{t+1}^{i} \leq (R_t^{ik} - FX^{ik})\mathcal{E}_{t+1}^{ik})$$

Therefore, the government default curve shifts downward, increasing the likelihood of a no-default equlibirum (Panel (d) of Figure 1).

2.2.6 From General Equilibrium to Local Average Treatment Effect (LATE)

The above analysis relies on the assumption that the initial steady-state equilibrium lies in the interior of the state space, which depends on the primitives, in particular, $\pi_0^i \in (0,1)$. If the initial intersection is a corner solution, then the steady-state π_0^i lies on one of the axes. First, in the "never default" region (Panel (a) of Figure 2), the curves meet only on the horizontal axis pre- and post-shock, so the default probability is (essentially) zero across the shocks we study. Second, in the "always default" region (Panel (b) of Figure 2), the curves intersect on the vertical axis pre- and post-shock, so the default probability always equals one.

For interior steady-state cases in which shocks to $\mathcal{M}_{t+1}^{k}$ can shift the initial equilibrium, there can be a single tangency or multiple crossings. Point A in Panel (c) of Figure 2 illustrates the "compliers" region central to the instrumental variable identification strategy developed in Section 4: these debtors' response path to a contractionary foreign monetary shock is the same as that illustrated in Section 2.2.5, delivering the local average treatment effect (LATE) for this group.

One caveat is that there is also a "non-instrumentable" region in which either the mapping is non-monotonic or foreign monetary shock is not the key driver for default. Point B in Panel (d) of Figure 2 illustrates a scenario in which after a one-period shock, default probability and rollover rate rise initially but subsequently fall back toward the initial equilibrium.[6] The instrumental variable design does not identify causal effects in this case, and recognizing this limitation clarifies scope and motivates complementary identification strategies (e.g., difference-in-difference) as robustness checks in Section 6.3.

As an illustrative stochastic baseline, assume the debtor *i's* fiscal capacity $\mathcal{T}_{t+1}^{i}$ and foreign money growth $\mathcal{M}_{t+1}^{k}$ are mutually *independent* and log-normally distributed. This statistical independent assumption implies exogeneity—country k's monetary decisions do not respond to any confounding factors (e.g., wars, financial crises) in debtor i.

Proposition (1): Assuming no domestic monetary response, let $\mathcal{T}_{t+1}^{i} \sim \text{Lognormal}(\mu_{t_i}, \sigma_{t_i}^2)$ and $\mathcal{M}_{t+1}^{k} \sim \text{Lognormal}(\mu_{m_k}, \sigma_{m_k}^2)$ be independent random variables, then we have

(a) $\ln \mathcal{T}_{t+1}^{i}$ and $\ln \mathcal{M}_{t+1}^{k}$ are independent, so $\ln \mathcal{T}_{t+1}^{i} + \ln \mathcal{M}_{t+1}^{k} \sim \text{Normal}(\mu_{t_i} + \mu_{m_k}, \sigma_{t_i}^2 + \sigma_{m_k}^2)$

(b) The default probability admits a closed-form expression

$$\pi_0^i \equiv \mathbb{P}\left(\mathcal{T}_{t+1}^{i} \leq R_t^{ik} \mathcal{E}_{t+1}^{ik}\right) = \Phi\left(\underbrace{\frac{\ln \sigma_{ik}}{\Xi}}_{\text{EXR responsiveness}} + \underbrace{\frac{-\mathbb{E}\left[\ln \frac{T_i'}{D_i}\right]}{\Xi}}_{\text{fiscal capacity}} + \underbrace{\frac{\overbrace{\mathbb{E}\left[\ln \mathcal{M}_{t+1}^{i}\right]}^{\text{assume }=0} - \mathbb{E}\left[\ln \mathcal{M}_{t+1}^{k}\right]}{\Xi}}_{\text{monetary prudence}} + \frac{\ln R_t^{ik}}{\Xi} \right) \tag{2.7}$$

where $\Xi \equiv \sqrt{\sigma_{t_i}^2 + \sigma_{m_k}^2} > 0$, and $\Phi(\cdot)$ denotes the standard-normal cumulative distribution function (CDF).

Proof: See Online Appendix A

Equation (2.7) delivers an interior steady state whenever the z-score is finite. Corner solutions arise only when the numerator is sufficiently negative ("never default" $\pi_0^i = 0$) or

[6] "Non-instrumentable" examples include the Greek debt crisis in 2010-2012, when foreign policy rates were falling rather than tightening; and Argentina's 2001 default, which was driven mainly by Brazil's devaluation and a rigid currency board system.

sufficiently positive ("always default" $\pi_0^i = 1$). Intuitively, the numerator can be decomposed into economically meaningful components: exchange rate responsiveness $\ln \sigma_{ik}$, fiscal capacity $-\mathbb{E}\left[\ln \frac{T_i'}{D_i}\right]$, monetary prudence $\mathbb{E}\left[\ln \mathcal{M}_{t+1}^i\right] - \mathbb{E}\left[\ln \mathcal{M}_{t+1}^k\right]$, interest burden $\ln R_{t+1}^{ik}$, and aggregate uncertainty $\Xi \equiv \sqrt{\sigma_{t_i}^2 + \sigma_{m_k}^2}$. This closed-form expression links the general equilibrium model to the empirical LATE interpretation.

2.3 Probabilistic Model that Micro-founds the Shift-share Instrument

The empirical design of using a shift-share instrumental variable to estimate causal output loss from sovereign default relies on differential exposure—encoded in endogenous currency denomination—to external monetary shocks. The probabilistic environment described in Section 0 naturally connects currency denomination to default risk, which in turn determines debtors' exposure to foreign monetary shocks. I build on the seminal Eaton and Kortum (2002) probabilistic trade framework to rationalize the relative stability of currency shares over time. The key prediction is that default probability responds to external monetary shocks through sticky currency shares, an empirical feature documented in Section 5.1.

Assume there is perfect competition in each currency market: a continuum of fund providers indexed by $j \in [0,1]$ lend out currency k at the same price—lending rate R_t^{ik} determined in t and due at $t+1$ (i.e., no market power or markups within a currency). Motivated by extreme depreciation episodes in developing countries, let the gross depreciation factors $\mathcal{E}_{t+1}^{i1}, \mathcal{E}_{t+1}^{i2},, \ldots, \mathcal{E}_{t+1}^{ik}, \ldots$ in an arbitrary period $t+1$ be random variables drawn independently (but not necessarily identically) across currencies from Fréchet distributions with a debtor-specific shape parameter θ^i and different currency-specific location parameters $\bar{e}^k$, that is, for all k, $\mathcal{E}_{t+1}^{ik} \sim \text{Frechet}(\bar{e}^k, \theta^i)$. Initially, assume debtor i

will borrow equal amounts from each currency. For model closure, assume government revenue $\mathcal{T}^i_{t+1}$ also follows a Fréchet distribution with the same shape parameter θ^i and a different debtor-specific location parameter $\overline{t}^i$.[7] The cumulative distribution function of these random variables are given as:

$$\mathbb{P}\left(\mathcal{E}^{ik}_{t+1} \leq \epsilon\right) = \exp\left(-\overline{e}^k \cdot \epsilon^{-\theta^i}\right) \qquad \text{and} \qquad \mathbb{P}\left(\mathcal{T}^i_{t+1} \leq \tau\right) = \exp(-\overline{t}^i \cdot \tau^{-\theta^i})$$

Using the max-stable property of the Fréchet distribution, the debtor i defaults when revenue cannot cover the largest currency-specific obligation, that is,

$$\pi^i_{t+1|t} = \mathbb{P}\left(\mathcal{T}^i_{t+1} \leq \max_k\{\mathcal{R}^{ik}_t \mathcal{E}^{ik}_{t+1}\}\right) \tag{2.8}$$

Appendix A2 derives a closed-form default probability as a function of the foreign-currency debt service:

$$\pi^i_{t+1|t} = \frac{\sum_k \overline{e}^k (R^{ik}_t)^{\theta^i}}{\overline{t}^i + \sum_k \overline{e}^k (R^{ik}_t)^{\theta^i}} \tag{2.9}$$

Intuitively, default risk increases with aggregate foreign-currency debt service $\sum_k \overline{e}^k (R^{ik}_t)^{\theta^i}$ and falls with fiscal capacity $\overline{t}^i$. It is also shaped by the tail thickness θ^i (extremeness of unexpected depreciations) and the currency-specific average depreciation level $\overline{e}^k$.

[7] Although these two assumptions may appear ad hoc, they align with real-world observations: emerging markets' currencies often experience extreme depreciation episodes (e.g., Thai baht, Indonesian rupiah, Malaysian ringgit during the 1997 Asian Financial Crisis; Mexico's 1994–1995 peso crisis). At the same time, government revenue in developing countries often exhibits "resource curse" heavily driven by external factors (e.g., oil-exporters such as Venezuela and Nigeria during the 1970s oil crisis). Consequently, windfall revenues can result in expenditure booms that undermine repayment capacity during subsequent currency crisis periods. A closer examination of the relationship between fiscal taxation, exchange rate and the implications for the stability of currency shares will be left to future work.

Assuming developing countries have strictly positive fiscal capacities to service debt (i.e., $\overline{t}^{\,i} > 0$) and a single foreign contractionary monetary shock (i.e., $\hat{R}_t^{i\$} > 0$ and for all $k \neq \$$, $\hat{R}_t^{ik} = 0$), linearizing Equation (2.9) around the steady state delivers a shift-share relationship that connects to the empirical design in Section 4 (see Appendix A3):

$$\hat{\pi}_{t+1|t}^{i} = \theta^{i}\left(1-\overline{\pi}^{i}\right)\left(\frac{\overline{e}^{\$}\left(\overline{R}^{i\$}\right)^{\theta^{i}}}{\sum_{k}\overline{e}^{k}\left(\overline{R}^{ik}\right)^{\theta^{i}}}\right)\hat{R}_t^{i\$} \tag{2.10}$$

where $\hat{\pi}_{t+1|t}^{i} \equiv \dfrac{d\pi_{t+1|t}^{i}}{\overline{\pi}^{i}}$ and $\hat{R}_t^{i\$} \equiv \dfrac{dR_t^{i\$}}{\overline{R}^{i\$}}$ are defined as the percentage deviations from steady state values.[8] Intuitively, the increase in default probability $\hat{\pi}_{t+1|t}^{i}$ is positively related to an exogenous shock $\hat{R}_t^{i\$}$ scaled by the endogenous exposure weight $\left(\dfrac{\overline{e}^{\$}\left(\overline{R}^{i\$}\right)^{\theta^{i}}}{\sum_{k}\overline{e}^{k}\left(\overline{R}^{ik}\right)^{\theta^{i}}}\right)$. For the instrument to be valid, the endogenous shares should be relatively stable over time so that variation in default probabilities is driven by the exogenous external monetary shock rather than contemporaneous reallocation across different currencies. Section 3.1 provides supporting evidence: currency denomination is sticky and does not change markedly with external monetary shocks.

For currency denomination shares, suppose the debtor i allocates borrowing across currencies by choosing the ex-post lowest debt service cost, i.e., $R_t^{i\$}\mathcal{E}_{t+1}^{i\$} \leq \min_{k\neq\$}\left\{R_t^{ik}\mathcal{E}_{t+1}^{ik}\right\}$.

[8] Since default probabilities and interest rates are easier to interpret in percentage points, Appendix A3 also derives an expression for that case. Define $\tilde{\pi}_{t+1|t}^{i} \equiv \pi_{t+1|t}^{i} - \overline{\pi}^{i}$ and $\tilde{R}_t^{i\$} \equiv R_t^{i\$} - \overline{R}^{i\$}$, both expressed as percentage point deviations from steady state values. Then we have $\tilde{\pi}_{t+1|t}^{i} = \left(\dfrac{\theta^{i}\overline{\pi}^{i}\left(1-\overline{\pi}^{i}\right)}{\overline{R}^{i\$}}\right)\left(\dfrac{\overline{e}^{\$}\left(\overline{R}^{i\$}\right)^{\theta^{i}}}{\sum_{k}\overline{e}^{k}\left(\overline{R}^{ik}\right)^{\theta^{i}}}\right)\tilde{R}_t^{i\$}$.

Under repeated draws, a continuum of price-taking lenders, and common debtor-specific parameters $\overline{e}^k = \overline{e}$ for all currency k, debtor i's borrowing share in currency \$ is:

$$s^i_{t+1|t}\left(\$\right) = \mathbb{P}\left(R^{i\$}_t \mathcal{E}^{i\$}_{t+1} \le \min_{k \neq \$}\left\{R^{i\$}_t \mathcal{E}^{i\$}_{t+1}\right\}\right) = \int_0^\infty \prod_{k \neq \$}\left(1 - \exp\left(-\overline{e}\left(\frac{R^{i\$}_t}{R^{ik}_t}\right)^{-\theta^i} \epsilon^{-\theta^i}\right)\right) dG_{\mathcal{E}^{i\$}}(\epsilon)$$

While the general form can only be solved with numerical methods, the above equation delivers a closed-form solution when there are only two currencies:

$$s^i_{t+1|t}(\$) = \frac{(R^{i£}_t)^\theta}{(R^{i£}_t)^\theta + (R^{i£}_t)^\theta} \tag{2.11}$$

with comparative statics $\frac{\partial s^i_{t+1|t}(\$)}{\partial R^{i\$}_t} < 0, \quad \frac{\partial s^i_{t+1|t}(\$)}{\partial R^{i£}_t} > 0$, and most interestingly:

$$\frac{\partial s^i_{t+1|t}(\$)}{\partial \theta^i} \propto -(r^{i\$}_t - r^{i£}_t)$$

where $r^{i\$}_t \equiv R^{i\$}_t - 1$ denotes the net interest rate. This expression implies that if the U.S. lending rate is lower than the British pound lending rate, the U.S. dollar share increases with θ^i, which captures a thinner-tailed Fréchet distribution. In other words, when extreme depreciations are more likely (i.e., smaller θ^i), currency denomination shares become less responsive to systematic relative cost differences, vice versa.

This expression parallels—though is not identical—to Eaton and Kortum (2002) and yields three testable implications: first, if one currency repeatedly delivers lowest service cost, borrowing tends to concentrate in that currency, and its dominance increases with increasing counterpart $\overline{e}^{£}$, which captures how extreme counterpart deprecation can happen. Second, currency share $s^i(\$)$ depends on both own and rival debt service cost, so a modest one-off shock may not necessarily change the composition sharply over short-horizons—currency denomination is relatively sticky, and the shape parameter θ^i governs the sensitivity to relative costs. Third, with repeated draws, exchange rate fluctuations average out over time and are absorbed into parameters that capture the friction in the currency market. Last but not least, the "all-currency denomination" form mirrors the structure of the default probability in Equation (2.9), motivating the shift-share design discussed in Section 4.

3. Data

3.1 Currency Denomination (Shares) and Monetary Shocks Measures (Shifts)

For the endogenous shares, I retrieved the data on the currency compositions of long-term public and publicly guaranteed debts for developing countries from the *International Debt Statistics* (IDS) published by the World Bank Group.[9] This country-year-currency specific variable covers external debts with maturities exceeding one year, held by either government or private debtors guaranteed for repayment by a public entity (e.g., a state-owned enterprise). The data are predominantly categorized into the currencies of six major advanced economies: U.S. dollars, Japanese yens, British pounds, Deutsche marks, French francs, and Swiss francs.[10]

[9] Due to debt relief policies, the graduation of some countries from debtor lists (e.g., Chile), and the loan-based nature of the data compilation, the historical data series from the *International Debt Statistics* (IDS) online portal have been revised. To ensure both accuracy and comprehensiveness, I also consulted data from the *Global Development Finance* (the predecessor of the IDS) using their 2006 & 2010 CD-ROMs, along with their corresponding publications from the 2010s. See Table A1 for more details.

[10] Data on currency shares of short-term debt (i.e., obligations with maturities under one year) are not available. Since 2000, the euro replaced the Deutsche Mark and the French Franc. In addition to the six major currencies listed above, the International Debt Statistics (IDS) also reports external debts denominated in 1. "Other Currencies" (e.g., Chinese Renminbi) 2. "Multiple Currencies" (e.g., currency-pool loans from the World Bank) and 3. Special Drawing Rights (SDRs). For categories 1 and 2, the IDS

Table 1 summarizes the currency compositions of external debts for all developing countries in the sample from 1970 to 2010. Despite the dominant role of U.S. dollars in foreign debt denomination, it highlights significant variations in the shares of external debts denominated in other advanced economies' currencies. Notably, at the 90th percentile, the share of debts denominated in Deutsche Marks accounted for 12.2% and in French Francs for 22.4% before 2000, with the Euro share increases to 47.3% after 2000.

Figure 3 visualizes the currency denomination trends of selected countries over time. Two patterns stand out: first, although the U.S. dollar dominates, the Japanese yen also plays an important role—especially for some Asian borrowers—suggesting differential exposure to external monetary shocks. Second, the grey dashed lines mark the annual level narrative contractionary U.S. monetary shocks classified by Romer and Romer (2023). Notice that the U.S. dollar shares do not change overnight. While the longer-run trends are evident, such effects rarely materialize within short horizons.

For the exogenous shifts, I retrieved long-term interest rate series (typically government bond rates) for the six advanced economies over the 1970-2010 period from the *Jordà-Schularick-Taylor Macrohistory Database* (Jordà et al. 2016). The U.S. and U.K. narrative monetary shocks series come from Cloyne and Hürtgen (2016).

3.2 Binary Default Indicators

For binary sovereign default classifications, I refer to the *Sovereign Default Database* compiled by Kuvshinov and Zimmermann (2019), who have integrated various default

data do not indicate whether the debt is in domestic currency or foreign currency. It also does not specify which specific foreign currencies are involved. While the IDS records (incomplete) debt stock for private non-guaranteed sector, it does not specify the currency denomination for this category. See Figure A1 for the overall structure of the dataset.

classifications—country-year varying binary indicators—drawn from the literature, including works by Beers and Chambers (2006), Beim and Calomiris (2000), Laeven and Valencia (2020), Reinhart and Rogoff (2011), and Reinhart and Trebesch (2016). This comprehensive database enables robustness checks using alternative default classifications from the literature. In the baseline LP-SSIV results, I use the first-year 0/1 default indicators classified by Standard and Poor's, as specified in Kuvshinov and Zimmermann (2019).[11] Since countries may have defaulted multiple times over the 1970-2010 period, in the LPDiD setup, I use in-default indicators along with a clean control setup to address negative weight effects (Borusyak and Hull 2024; Dube et al. 2025).

3.3 Continuous Default Measures

One important recent development in the sovereign default literature is the recognition of partial default: while the main discipline of debt repayments holds, countries may partially repay and renegotiate with individual creditors, blurring the boundary between default and non-default (Arellano et al. 2023). Therefore, a simple binary default indicator may not capture the rich variation in the extent of default.

To address this issue, I use the IDS arrears data—late or missed debt service payments—as a measure of partial default. Figure A2 plots the distribution of arrears, expressed in current U.S. dollars across all country-years in the sample, treating the world aggregate as the creditor. One salient feature is the mass at zero: many country-year

[11] Since Table 1 shows that sovereign debt can be denominated in different currencies, a natural concern is that countries may selectively default on one but not all currencies. In practice this is rare because within the same legal jurisdiction, most sovereign bonds are issued at the "pari passu" (Latin word for "equal footing") basis, implying no contractual seniority and equal treatment of creditors regardless of the currency denomination, which makes selective default uncommon (Aguiar and Amandor 2021; Wright 2014; Schumacher et al. 2012).

observations are recorded as exact zeros. This empirical pattern motivates separate analysis of extensive and intensive margins of default cost—a potentially informative perspective that the literature has yet to explore (Section 6.2).

3.4 Other Macroeconomic Variables

The dependent variable—real GDP per capita in constant national currency—is retrieved from the *World Development Indicators* (WDI) published by the World Bank, as well as the GDP deflator. Nominal exchange rate data come from the *International Financial Statistics* published by the International Monetary Fund (IMF). The data on external debt stock, ratios of the short-term to total external debt, and the ratios of reserves-to-imports (in months) are retrieved from the *International Debt Statistics*. The GDP growth data are from the FRED website. In addition, I use the binary indicators for banking, currency, and political crises from Kuvshinov and Zimmermann (2019), supplemented with updated crisis information from Laeven and Valencia (2020). I also include the binary democracy indicators from Acemoglu et al. (2019), the Chinn and Ito (2006) capital openness index, and the IMF arrangement data from Vreeland (2007). Exchange rate regimes are classified as pegged if countries have a coarse code of 1 or 2, and as floating if they have a code of 3, 4, or 5, following the data compiled by Ilzetzki et al. (2019).

Table A2 reports summary statistics for all variables used in the empirical analysis, and Table A3 reports a balance test between "defaulters" and "never-defaulters—countries with and without default experience—and shows systematic differences between the two groups: defaulters exhibit slower output growth, higher inflation, higher currency depreciation, and higher debt-to-GDP ratios, more frequent banking and currency crises, and are less likely to maintain a fixed exchange rate regime, which is consistent with limited credibility under hard pegs. These imbalances imply that a plain OLS comparison may lead to omitted variable bias and is not an "apples-to-apples" comparison.

4. Empirical Strategy

The omitted variable bias is a central challenge when using OLS regression to establish causal statements. However, understanding the direction of the bias helps identify the mechanisms. Treat output loss magnitude as positive—larger values indicate more negative output growth—and consider two thought experiments: first, countries in financial crisis or at war are more likely to default on external debt and simultaneously experience larger output loss magnitude. These domestic factors are positively correlated with default probability and output loss magnitude, so the OLS estimate may be biased upward, overestimating the true causal effect. Second, default can relieve debt overhang, help debtors renegotiate better terms, and reduce the need for fiscal austerity. Such benefits are positively related to the default decision yet negatively related to output loss magnitude, so the OLS estimate may be biased downward, underestimating the causal effect. Since both channels plausibly apply, the net direction of bias is an empirical question that will be addressed in the analysis that follows.

I begin with the baseline local projection–ordinary least square (LP-OLS) regression. For country i in year t, the dependent variable is the long-difference of the log real GDP per capita $y_{it+h} - y_{it-1}$ for horizons $h = 0,1,\ldots,6$. The variable of interest is the first-year default indicator $D_{it} \in \{0,1\}$, which indicates whether year t is the first year of default for country i in a particular default episode. All regressions include country fixed effects and a comprehensive set of baseline controls $\mathbf{X}_{it}$, including up to two lags of (i) regression-specific variables such as the treatment variable (e.g., first-year default indicators), one lead and one lag of the shift-share instrument to address "lead-lag exogeneity" (Stock and Watson

2018), and the "incomplete shares" defined as one minus the sum of the major six named currency shares (Borusyak et al. 2022); (ii) continuous variables such as the first differences of log real GDP per capita, log GDP deflator, log nominal exchange rate, and log world GDP, as well as the levels of debt-to-GDP ratios and the Chinn and Ito (2006) capital openness index; and (iii) binary indicators for banking, currency, political crises (with 1 indicating war/coup/political transition), and democracy (Acemoglu et al. 2019). The same baseline control set is used across all OLS, IV (both first-stage and 2SLS), and all robustness specifications.

Henceforth, the LP-OLS specification is given as:

$$y_{it+h} - y_{it-1} = \alpha_{ih} + \beta_h D_{it} + \gamma_h \mathbf{X}_{it} + v_{it+h} \qquad h = 0,1,\ldots,6 \tag{4.1}$$

where the key coefficient estimate of interest is $\{\beta_h\}$, the impulse response of the cumulative output loss due to sovereign default in the horizon h.

To estimate the causal effect of sovereign default on output loss, I use the local projection–shift-share instrumental variable (LP-SSIV) approach. To be specific, I first define the "interest rate exposure", ΔIRE_{it}, for country i in year t as follows:[12]

$$\Delta IRE_{it} \equiv \sum\nolimits_k \text{currencyshare}_{it}^k \times \Delta\text{interestrate}_t^k \tag{4.2}$$

where $\text{currencyshare}_{it}^k$ represents the share of country i's external debts denominated in advanced economy *k's* currency in year *t*. $\Delta\text{interestrate}_t^k$ denotes the change in the long-

[12] Following Autor et al. (2013), I use "exposure" to denote share-weighted interest rate shocks constructed with pre-determined currency shares.

term interest rate in advanced economy k from year $t-1$ to year t, and it becomes a stationary after differencing (note that this variable is invariant to individual country i).[13]

Therefore, the first-stage regression is given as follows:

$$D_{it} = a_i + b\Delta IRE_{it-\ell} + g\mathbf{X}_{it} + \eta_{it} \tag{4.3}$$

Similarly, for the narrative monetary shocks, I define:

$$IRE_{it}^{USRR} \equiv \text{currencyshare}_{it}^{USD\$} \times \text{USRRshock}_t$$
$$IRE_{it}^{UKCH} \equiv \text{currencyshare}_{it}^{GBR£} \times \text{UKCHshock}_t$$

where a_i represents country fixed effects. $\Delta IRE_{it-\ell}$ denotes the ℓ lag of the "interest rate exposure" (i.e., the shift-share instrument). I use the monthly narrative monetary policy shocks series for the Federal Reserve and the Bank of England constructed by Cloyne and Hürtgen (2016). Figure A3 shows that the U.K. monetary shocks series is more volatile and tends to exhibit larger-amplitude shocks than its U.S. counterparts—especially during the 1975–1985 period—with more visible decline in volatility over time. Aggregating monthly shocks to annual measures by summing or averaging can therefore attenuate the monetary policy surprises by offsetting opposite-sign monthly values and make cross-country comparisons sensitive to within-year volatility. Accordingly, I define the annual monetary shock as the single largest-magnitude monthly shock in each year (sign retained) and then standardize the resulting annual series within the Federal Reserve and the Bank of England over the full samples respectively. I denote these annualized measures USRRshock_t and UKCHshock_t, and the first-stage regression is given as follows:

[13] Table A9–Table **A11** in the Online Appendix assess robustness to alternative shift-share instrument specifications, such as lagged currency shares weighting, short-term interest rates, and a leave-one-out variant that excludes U.S. dollar denominated debt. See Section 5.2 for detailed discussions.

$$D_{it} = a_i + b \cdot (IRE_{it}^{USRR}, IRE_{it}^{UKCH}) + g\mathbf{X}_{it} + \xi_{it}$$

Henceforth, the LP-SSIV specification is given as:

$$y_{it+h} - y_{it-1} = \alpha_{ih} + \beta_h \hat{D}_{it} + \gamma_h \mathbf{X}_{it} + v_{it+h} \qquad h = 0,1,\ldots,6 \qquad (4.4)$$

5. Baseline Results

5.1 First-stage Results

Table 2 presents significant positive first-stage relationships between the first-year default binary indicator D_{it} and the second lag of the "interest rate exposure" ΔIRE_{it-2}. Columns (1)–(3) show that first-stage coefficient estimates for the first, second, and third lags—ΔIRE_{it-1}, ΔIRE_{it-2}, and ΔIRE_{it-3}—are all positive, aligning with our economic intuition: when an advanced economy raises interest rates, the subsequent appreciation of its currency increases the likelihood of default for countries with a higher proportion of debt denominated in that currency. Nevertheless, only the second-lag coefficient estimate on ΔIRE_{it-2} is statistically significant, suggesting a delayed response of default risk to foreign interest rate hikes.

Table A7 and Table A8 examine mechanisms that can rationalize this delay. Table A7 shows that countries with larger shares of variable-rate debt (e.g., LIBOR- or U.S. prime-linked sovereign debt) are more susceptible to the interest rate shock. The coefficient estimates on the interaction terms $\Delta IRE_{it-\ell} \times \text{sharevarrate}_{it-1}$ are generally positive and statistically significant, implying that holding $\Delta IRE_{it-\ell}$ constant, a higher variable-rate share raises the default probability. Conversely, Table A8 shows that countries with longer average maturities are less vulnerable to $\Delta IRE_{it-\ell}$, as shown by negative coefficient estimates on the interaction term $\Delta IRE_{it-\ell} \times \text{Maturity}_{it-1}$, meaning that conditional on

$\Delta IRE_{it-\ell}$, longer maturities reduces the default probability. Together, these empirical patterns are consistent with a lagged transmission channel: floating rate liabilities are repriced more quickly when foreign interest rate rises, whereas longer maturities delay the pass-through.

In line with this timing, column (4) of Table 2 shows that the second lag remains the most robust instrument among the three lags considered. Historical experience echoes this delay: Mexico's 1982 default followed the onset of the Volcker era tightening by roughly two to three years, indicating prolonged financial strain before the default decision. To be specific, the first stage coefficient estimate of 0.027 implies that a one-unit increase in ΔIRE_{it-2} is associated with a 0.027 unit increase in the probability of sovereign defaults two years later.[14] Given the strength of this instrument and to satisfy the lead-lag exogeneity requirement, I use the second lag ΔIRE_{it-2} as the primary instrumental variable while including both the first ΔIRE_{it-1} and third lags ΔIRE_{it-3} as controls in all regressions.

[14] Since only linear OLS regressions yield residuals orthogonal to regressors and fitted values, two-stage least squares (2SLS) regression does not apply to nonlinear models—such as logistic regressions that often yield better interpretation for binary outcomes—for instrumental variable estimation (i.e., the "forbidden regressions"; see Chapter 4.61 of Angrist and Pischke (2009)). Nevertheless, column (2) of Table A6 reports logit regression results using the same baseline controls and country fixed effects as in Table 2 to complement OLS first-stage results (reproduced in column (1)). To interpret the logit results, the coefficient estimate on ΔIRE_{it-2} is 0.815, which means one unit increase in ΔIRE_{it-2} is associated with $\exp\left(0.815\right)-1=126\%$ higher odds-ratios of sovereign default. Alternatively, when ΔIRE_{it-2} increases from 0 to 1, the default probability increases by $\frac{\exp\left(-3.042+0.815\right)}{1+\exp\left(-3.042+0.815\right)}-\frac{\exp\left(-3.042\right)}{1+\exp\left(-3.042\right)}=5.18\%\text{pt}$ (unit in percentage point; %pt). Although this logit coefficient is slightly larger than its linear OLS counterpart (i.e., default probability increases by 2.6%pt), the results are consistent because of the nonlinear nature of logit function with slopes varying at different values of ΔIRE_{it-2}. Since the logit regression drops observations from groups with no within-group variation in the first-year default indicator (e.g., never-defaulters or always-defaulters), the number of observations in column (2) is smaller than that in the OLS sample. Therefore, column (3) reports OLS estimates on the same logit sample; interpretation follows as above.

Last but not least, column (5) shows that the shift-share instruments IRE_{it}^{USRR} and IRE_{it}^{UKCH} derived from the narrative monetary shocks also exhibit positive and statistically significant first-stage relationships with the first-year default indicator. This finding is important for two reasons: first, unlike the first difference in foreign interest rates, the narrative monetary shocks provide truly exogenous variations that are not correlated with any other predictable factors abroad. Second, both coefficient estimates of IRE_{it}^{USRR} and IRE_{it}^{UKCH} are positive and significant, suggesting that quasi-random monetary policies from outside the hegemon can also influence default decisions.[15]

5.2 Main Results: LP-OLS and LP-SSIV Regressions

Columns (1) and (2) in Table 3 show that the coefficient estimates for cumulative output loss from the LP-SSIV regressions are larger in magnitude but less persistent compared to those from the LP-OLS regressions. To be specific, the LP-OLS results in column (1) show that, on average, defaulting countries face a 2.66% output loss in the first year, with the cumulative output loss peaking at around 5.21% in the fifth year. The negative and statistically significant effects of defaults persist for up to sixth year, with the joint significance test yielding a p-value close to 0, confirming this persistence. However,

[15] Although Table 2 shows a smaller first-stage coefficient estimate on IRE_{it}^{USRR} than on IRE_{it}^{UKCH} (0.024 versus 0.053), Table A2 indicates that IRE_{it}^{USRR} has a much higher standard deviation than IRE_{it}^{UKCH} (0.437 versus 0.095). To make magnitudes comparable, I scale first-stage coefficient estimates by the regressors' dispersion: one-standard-deviation increase in IRE_{it}^{USRR} is associated with $0.024 \times 0.437 = 0.010$ increase in default probability, whereas one standard deviation increase in IRE_{it}^{UKCH} is associated with $0.053 \times 0.095 = 0.00535$ increase in default probability. After standardized, the implied first-stage effect is larger for the U.S. monetary shock, consistent with the hegemonic role of Federal Reserve's monetary policy.

the LP-OLS results alone cannot establish the direction of causality due to potential omitted variable bias.

Column (2) of Table 3 presents the baseline LP-SSIV results. The coefficient estimate for year 0 is -8%, though it is not statistically significant. This estimate becomes more negative to -18.5% in the second year and reaches statistical significance at the 5% level. It remains around -15% to -18% in the third and fourth year before reverting back to zero by the sixth year.

Panel (a) in Figure 4 visualizes the comparisons between LP-OLS and LP-SSIV estimates. While most coefficient estimates in the LP-SSIV regressions are larger in magnitude than those from the LP-OLS regressions, they are less persistent—indeed, the impulse response returns to zero after about six years, and the joint significance test yield a p-value that is not statistically significant, suggesting that defaulting on external debt may not have a long-lasting negative impact on countries.

The multiple weak instrument tests in Table 3 reinforce the validity of the shift-share instrument. The overall Kleibergen-Paap F-statistic is 21.14, well above the standard threshold of 10. At the same time, column (3) displays the p-values of the Anderson-Rubin F-statistics at each horizon, with statistical significance observed for second to the fourth year. Using the stacking method, column (4) shows the p-value in the second year barely rejects the null hypothesis on the equality between LP-OLS and LP-SSIV estimate at the 10% significance level, but other horizons lack insufficient statistical power to detect a significant difference.

Table 4 shows that while many coefficient estimates for "interest rate exposure" constructed with narrative monetary shocks are statistically insignificant, their magnitude and persistence are reassuringly close to the baseline results from Table 3. To be specific, using both IRE_{it}^{USRR} and IRE_{it}^{UKCH} as joint instruments yields a first-year coefficient estimate

(-8.15%) that closely resembles that (-8%) from differencing foreign interest rates as shown in Table 3. Furthermore, the negative impacts peak at -20.6% in the second year and then gradually dissipate, aligning closely with the peak value (-18.5%) and the overall shape of the impulse response as reported in Table 3. Regrettably, most coefficient estimates fail the weak instrument tests. However, it is important to note that these narrative monetary shocks series may contain measurement errors and noise, making it challenging to estimate the causal effect of monetary policy within a single economy, let alone aggregating them to assess causal effects in a completely different context.

In the Online Appendix, I examine the robustness of alternative shift-share instrument specifications. Table A9 follows Autor et al. (2013) and constructs the shift-share instrument using the lagged currency shares $\text{currencyshare}_{it-1}^{j}$, that is, $\Delta IREls_{it} \equiv \sum_{j} \text{currencyshare}_{it-1}^{j} \times \Delta\text{interestrate}_{t}^{j}$. Table A10 uses short-term interest rates in advanced economies (also from the JST Macrohistory Database) as exogenous shocks, that is, $\Delta IREstir_{it} \equiv \sum_{j} \text{currencyshare}_{it}^{j} \times \Delta\text{STinterestrate}_{t}^{j}$. Last but not least, Table A11 presents a leave-one-out specification that excludes U.S. dollar denominated debt, that is, $\Delta IREnoUSD_{it} \equiv \sum_{j \neq US\$} \text{currencyshare}_{it}^{j} \times \Delta\text{interestrate}_{t}^{j}$. Across these specifications, most results are highly consistent and quantitatively comparable. Although the excluding-U.S.-dollar-debt case exhibits weak instrument (Kleibergen–Paap F-statistics is 3.91), the point estimates remain stable for the first two horizons.

5.3 Addressing Exclusion Restriction: Control Function Approach

To address potential violations of the exclusion restriction assumption, I follow the method outlined by Jordà et al. (2020), who integrates the control function framework developed by Conley et al. (2012) and Wooldridge (2015) into local projection methods. To be specific, setting aside subscripts and control variables for clarity, let y be the dependent variable, D be the binary treatment variable, and z be an the instrumental variable, we can prove the following proposition:

Proposition (2): Suppose the true model is $y = \beta D + \phi z + v$ with $\phi \neq 0$—that is, the exclusion restriction fails. Let the first-stage regression be $D = bz + \eta$ with $b \neq 0$, then:

1) The 2SLS estimator based on z will be inconsistent:

$$\hat{\beta}^{IV} \xrightarrow{p} \beta + \frac{\phi}{b} \tag{5.1}$$

2) Consider the subsample of never-defaulting countries—i.e., $D = 0$ in all observed years. In this subsample, the structural equation simplifies to $y^{ND} = \phi^{ND} z^{ND} + v^{ND}$. Assume the instrument z remains exogenous for the never-defaulting countries, that is, $\mathbb{E}\left[z^{ND} v^{ND}\right] = 0$. Suppose further that the direct spillover of the instrument z in default episodes is proportional to that in never-default episodes, that is, $\phi = \lambda \phi^{ND}$ with the relative spillover parameter $\lambda > 0$. Define $\hat{\phi}^{ND}$ as the OLS estimator of ϕ^{ND} from the never-default subsample. Then the spillover-corrected 2SLS regression in the full sample,

$$y - \lambda \hat{\phi}^{ND} z = \beta D + (\phi - \lambda \hat{\phi}^{ND}) z + v \tag{5.2}$$

will yield a consistent estimator for the true treatment effect β

$$\hat{\beta}_{CF}^{IV}(\lambda) = \frac{\frac{1}{N}\sum_{i=1}^{N}\left(\left(y_i - \lambda \hat{\phi}^{ND} z_i\right) z_i\right)}{\frac{1}{N}\sum_{i=1}^{N} D_i z_i} \xrightarrow{p} \beta \tag{5.3}$$

as long as $\frac{1}{N}\sum_{i}^{N} z_i^2 \xrightarrow{p} \mathbb{E}\left[z^2\right] < \infty$ when $N \to \infty$.

Proof: The first part of the proof is detailed in Jordà et al. (2020). See Appendix A for the second part of the proof.

Panel (b) of Figure 4 shows the default cost estimates after controlling the spillover effects. To be specific, the case where $\lambda = 0$—i.e., no spillover—collapses back to our baseline LP-SSIV estimates. If we assume $\lambda = 1$, that is, the spillover experienced by defaulting countries is the same as that experienced by never-defaulting countries ($\phi = \phi^{ND}$), the green shaded area shows the bounds of the true causal effect on the cost of sovereign default. Although these IV spillover-corrected estimates closely align with the baseline LP-SSIV estimates, they exhibit a more persistent pattern, with output loss at around -20% in the sixth year, whereas the baseline estimates eventually return to zero by year 6. Alternative assumptions on the relative spillover parameter $\lambda > 0$ yield similar results: for example, if we assume $\lambda = 0.5$, that is, the defaulting countries experience less spillover than the never-defaulting countries, then the sixth-year estimate of output loss is around -10% (in turquoise). On the other hand, if we assume $\lambda = 1.5$ that defaulting countries experience larger spillover than never-defaulting countries, the sixth-year output loss can reach -35% (in orange). This persistence is not surprising, as an increase in foreign interest rates can generate positive spillovers through the trade channel: higher foreign interest rates depreciate local currency, boosting domestic exports and stimulating economic growth. Therefore, even modest positive spillovers imply that the true underlying cost of default is more severe and persistent than the baseline estimates suggest.

6. Robustness Checks

This section evaluates the robustness of the baseline LP-SSIV estimates along two dimensions. First, Sections 6.1 and 6.2 examine whether the main findings are sensitive to alternative default classifications, comparing results between using binary default indicators and continuous arrears-based partial default measures (Section 6.2) as the treatment variables. Second, Section 6.3 uses modern difference-in-differences (DiD) techniques, which rely on alternative identification assumptions (i.e., parallel trends) relative to the LP-SSIV design, to examine the stability of the causal interpretation. The DiD estimators can explicitly include year fixed effects, which helps address concerns that common time trends drive both foreign interest rate hikes and default decisions, but at the cost of less flexibility for exploring heterogeneous effects. The LP-SSIV approach, by contrast, can trace differential default costs across regimes (as shown in Section 8) but cannot absorb global time trends as directly as the DiD estimators. The consistency of the results across two empirical strategies with complementary strengths reinforces the credibility of the main findings.

6.1 Different Binary Default Classifications

Panel (c) in Figure 4 shows that the baseline results remain robust across alternative default classifications. Despite different methodologies being used for classification, the baseline coefficient estimates are consistently similar and quantitatively comparable across different classification schemes from previous studies by Reinhart and Rogoff (2011), Laeven and Valencia (2020), and Beim and Calomiris (2000). However, a notable exception is the default classification by Detragiache and Spilimbergo (2001), which shows significantly more negative coefficient estimates than others. This discrepancy is likely due to differences in sample sizes and time periods, as well as somewhat subjective judgments regarding the exact timing of default episodes.

6.2 Partial Default with Arrears Data

If we treat default intensity as a continuous measure rather than a binary indicator, the baseline estimation framework remains the same, requiring only minor renaming of the treatment variable. Let $A_{it} \in [0,\infty]$ denote arrears—late or missed debt service payments—of country i to the world aggregate measured in current U.S. dollars. As shown in Figure A2, there is a non-negligible mass at zeros in the arrears data. Regressing arrears levels can then be misleading, so I apply log-transformations and interpret coefficient estimates accordingly.

Define the log-transformation of A_{it} as follows: $\Lambda_1(A_{it}) = \ln(A_{it})$, $\Lambda_2(A_{it}) = \ln(1 + A_{it})$, and $\Lambda_3(A_{it}, D_{it}) = \ln\left(1 + 100 \times \frac{A_{it}}{D_{it}}\right)$.[16] The first two are common in the firm entry-exit literature to handle zeros and yield interpretable elasticities; the third follows Chen and Roth (2024)'s guidance for anchoring the interpretation at an economically meaningful variable. Using these variables Λ_{it} as proxies for partial default, the baseline LP-OLS and LP-SSIV specifications remain unchanged (as in Equations (4.3)–(4.4)):

$$\Lambda_{it} = a_i + b\Delta IRE_{it-\ell} + g\mathbf{X}_{it} + \eta_{it}$$

$$y_{it+h} - y_{it-1} = \alpha_{ih} + \beta_h \hat{\Lambda}_{it} + \gamma_h \mathbf{X}_{it} + v_{it+h} \qquad h = 0,1,\dots,6$$

Table 5 reports the main results using sovereign arrears as treatment intensity for partial default. Panel (a) shows the first-stage estimates. Among all three transformations, The first-stage coefficient estimates of $\Delta \ln(\text{arrears})$, $\Delta \ln(1 + \text{arrears})$, and $\Delta \ln\left(1 + 100 \times \frac{\text{arrear}}{\text{debt}}\right)$ are all statistically significant with the instrument ΔIRE_{it-3} instead of ΔIRE_{it-2} as in Table 2. Some signs differ because of transformation, and interpretation

[16] The analysis uses $\ln\left(1 + 100 \times \frac{\text{arrears}}{\text{debt}}\right)$ instead of $\ln\left(1 + \frac{\text{arrears}}{\text{debt}}\right)$ because for small arrears-to-debt ratios, $\ln\left(1 + \frac{\text{arrears}}{\text{debt}}\right) \approx \frac{\text{arrears}}{\text{debt}}$, so a one-unit increase represents a 100-percentage-point (100%pt) change in the arrears-to-debt ratio. However, Figure A2 indicates that the distribution of arrears-to-debt ratios is highly right-skewed, with most values clustered near zero and a long right tail. Rescaling the ratio by 100 before taking logs—$\ln\left(1 + 100 \times \frac{\text{arrears}}{\text{debt}}\right)$—makes a one-unit increase correspond to a 1-percentage-point (1%pt) increase in the arrears-to-debt ratio, providing a more intuitive interpretation of the coefficient estimates (Chen and Roth 2024).

differs because of intensive versus extensive margins. In column (1), the elasticity of arrears $\Delta \ln(\text{arrears})$ with respect to ΔIRE_{it-3} is 0.126, meaning one unit increase in ΔIRE_{it-3} is associated with 12.6% increase in arrears among strictly positive arrears observations (i.e., excluding zero). In column (2), I use $\Delta \ln(1+\text{arrears})$, which takes both the intensive and extensive margins into account. The first-stage coefficient estimate falls to 0.053 with the same positive sign, consistent with the $\Delta \ln(\text{arrears})$ (0.126) without the mass at zero.

However, as Chen and Roth (2024) point out, the interpretation on coefficient estimates from either $\Delta \ln(\text{arrears})$ or $\Delta \ln(1+\text{arrears})$ should be treated with caution because they do not separate They recommend anchoring log changes at economically meaning variable. Accordingly, I use the arrear-to-debt ratio and anchor the change at the threshold when it increases from 0 to infinitesimal positive value, which I interpret as a threshold at which a country enters initial default. The first-stage coefficient estimate evaluated at this anchor $\Delta \ln\left(1+100\times\frac{\text{arrear}}{\text{debt}}\right)$ is 0.021, meaning 1 unit increase in ΔIRE_{it-3} is associated with a change of arrear-to-debt ratio from 0 to 0.00021 (i.e., 0.021 percentage point increase).

Panel (b) of Table 5 shows the LP-SSIV coefficient estimates using the three partial default measures. The output loss estimates from either $\Delta \ln(\text{arrears})$ or $\Delta \ln(1+\text{arrears})$ are relatively small in magnitude compared to those from $\Delta \ln\left(1+100\times\frac{\text{arrear}}{\text{debt}}\right)$ or baseline binary-default results as shown in Table 3. They peak by year 2 with cumulative output losses of -5.25% and -8.62% respectively. The Anderson-Rubin test statistics indicate that the instrument is only strong for the fourth year but have limited statistical power to reject the null hypotheses in other periods. Most impulse responses return to zero or slightly positive values within five years, resonating with the earlier results that default costs are not persistent.

Following Chen and Roth (2024), I interpret column (3) as the preferred specification that accounts for extensive and intensive margins. It implies an initial output loss of roughly -8% when the arrear-to-debt ratio increases from 0 to 0.00021—our proxy for the onset of default. The Anderson-Rubin test statistics suggest we cannot rule out the possibility of weak instrument except for year 4. Cumulative output loss peaks near -21.8% in the third year—comparable to the -18.5% peak in the second year as shown in Table 3—and then revert toward zero by the fifth year. Taken together, the arrears-based partial default measure—late or missed debt service repayment—reinforces the earlier finding that sovereign default is costly but temporary.

6.3 Difference-in-Difference Estimates on the Cost of Default

The credibility of using instrumental variables to establish causal statements relies on both the relevance and the exclusion restriction assumption. However, in macroeconomics, all variables are jointly determined, the validity of the exclusion restriction is often questionable. Therefore, an alternative design with different identification assumptions is useful for robustness checks. Difference-in-Differences (DiD) is another widely used empirical method to answer causal questions. It relies on the "parallel trends" assumption—absent treatment, treated and controlled units would have evolved similarly. Nevertheless, recent work (e.g., Roth et al. 2023; Callaway and Sant'Anna 2021; de Chaisemartin and D'Haultfœuille 2020; Borusyak et al. 2024; Sun and Abraham 2021) highlights pitfalls in multi-period settings where treatment status can switch on and off repeatedly (i.e., staggered treatment): the traditional two-way fixed effects estimators often use already-treated units from earlier periods in addition to never-treated units as controls, resulting in negative weights under heterogeneous treatment effects. This is especially problematic for estimating the cost of sovereign default using DiD, as repeated defaults occur over the sample period (e.g., Argentina defaulted in 1982, 1989, and 2001), making it challenging to ensure a fair comparison.

Building on the local projection framework introduced in Section 4, I follow the LP-DiD (local projection–difference-in-difference) design of Dube et al. (2025) and add a "clean control set", which offers a transparent solution to negative weight issues under staggered treatment. To be specific, since Chatterjee and Eyigungor (2012) estimate a quarterly "re-entry" probability of 0.0385 following default, which implies an average of 6.5 years of exclusion from international financial markets, I assume the effects of default gradually dissipate after six years, consistent with the gradual reversion toward zero in the empirical impulse response estimates shown in Panel (a) of Figure 4. Accordingly, I restrict the sample to countries that have not defaulted in the previous six years and apply the following LP-DiD specification:

$$y_{it+h} - y_{it-1} = \beta_h^{LP-DiD} \Delta D_{it} + \delta_t^h + \mathbf{X}_{it} + \epsilon_{it}^h \qquad h = -6, \ldots 0, \ldots 12 \tag{5.4}$$

restricting the estimation sample to a "clean control set":[17]

$$\begin{array}{ll} \text{default} & D_{it+j} = 1 \text{ for } 0 \leq j \leq h \quad \text{and} \quad D_{it-j} = 0 \text{ for } 1 \leq j \leq 6 \\ \text{clean control} & \Delta D_{it-j} = 0 \quad \text{for } -h \leq j \leq 6 \end{array}$$

In terms of specification, several changes are made from Equation (4.4): first, I include year fixed effects δ_t^h in line with the standard event-study setup to account for time-specific influences. Second, the independent variable used is the S&P in-default indicators, which equals to 1 for the entire period when a country remains in a default episode.

[17] Intuitively, the "clean control set" restrict samples to (i) treated in period t $D_{it} = 1, D_{it-1} = 0$ (ii) treated in both period t and $t-1$, $D_{it} = 1, D_{it-1} = 1$, (iii) untreated in both period t and $t-1$, $D_{it} = 0, D_{it-1} = 0$. In other words, it excludes the negative weights $D_{it} = 0, D_{it-1} = 1$ by design.

Panel (a) of Figure 5 shows LP-DiD estimates closely align with the spillover-corrected IV estimates in Panel (b) of Figure 4—the output loss remains persistent after controlling for spillovers, hovering around −30% even a decade later. This finding resonates with the more severe and enduring default cost once positive trade spillovers are accounted for. The stable parallel trends over the six-year pre-treatment window in Figure 5 further validate the robustness of the main results.

While the LP-DiD estimator is my preferred specification because of its transparency and theorical support for years of exclusion from the literature, one of the main concerns is the bias-variance tradeoff—the clean control set inevitably discards observations that may contain valuable information when fully utilized. Therefore, I also estimate the output loss using alternative DiD estimators from the literature.[18] Panel (b) of Figure 5 visualizes those coefficient estimates from imposing different assumptions (see Roth et al. (2023) for detailed comparisons). The central message—sovereign default entails significant output loss—remains.

7. Discussion Of Baseline Results

7.1 How Large is the Output Loss from a Sovereign Default?

Table 3 shows the causal estimates of the effect of a sovereign default on output loss: the initial drop is around −8%; it peaks at around −18.5%; persists through the fourth year; and gradually returns to zero by the sixth year. Are these causal estimates (reproduced in

[18] Since Callaway and Sant'Anna (2021) rely on never-treated units as controls, the effective sample size shrinks, resulting in larger point estimates and wider confidence intervals than those from other DiD estimators. For clarity, Figure A5 re-plots these DiD estimates from Figure 5 alongside the Callaway and Sant'Anna (2021) estimates for direct comparison.

Panel (a) of Table A5) consistent with observed default episodes, and how large are they relative to other crises? Panel (b) of Table A5 shows the cumulative output loss ($y_{t+h} - y_{t-1}$) with its counterparts from representative sovereign default episodes. Note that entries in Panels (b)–(d) are descriptive, not causal.

The initial −8% output loss is somewhat larger—but broadly consistent—with the average first-year output loss (around −5% to −7%) observed in the data, but the peak depth and persistence vary across episodes. For example, Argentina's 2001–2002 default and Greece's 2010 debt restructuring (not included in the baseline causal analysis due to missing currency denomination data, as Greece is classified as an advanced economy by the IDS) show peak losses around −16% to −18%, typically peaking by the second year, which is in line with the baseline causal estimates in Panel (a). However, the case of Greece in 2010 is worrisome, as it does not recover within six years, whereas Argentina returns to trend by the fifth year. In contrast, Russia's 1998 default and Ecuador's 1999 default show relatively smaller recessions: Russia rebounds within a year after an initial –5.3% loss, while Ecuador's output loss peaks at –7.2% in the second year but remains below trend until the fifth year. This heterogeneity in post-default outcomes is discussed further in Section 8.

Panel (c) of Table A5 shows that sovereign defaults appear as painful as typical currency crises, comparing output losses with Mexico's 1994–1995 Tequila Crisis and Thailand's 1997–1998 Asian Financial Crisis. The first-year output losses are comparable (−7.9% to −9.2%), but the impact of default is much more persistent: in currency crises, deviations from trend typically return to zero or positive within one to two years. One interpretation is that currency crises often involve depreciation that boosts exports, but they do not severely damage a country's reputation for honoring debt, allowing faster re-entry to international financial markets than after a sovereign default.

A useful benchmark is the 2008 Global Financial Crisis (Panel (d) of Table A5), which features around -5% decline in real GDP per capita in the United States, and it takes over five years for output deviations from trend to return to zero. At the same time, Sufi and Taylor (2021) show that over the 1870–2006 sample (excluding wars), a banking crisis typically lowers real GDP per capita by around -5% to -6% relative to trend over a six-year horizon, and the cumulative loss persists beyond the sixth year. The IV spillover-corrected estimates (Panel (b) of Figure 4)—consistent with the difference-in-difference estimates in Figure 5—suggest significantly larger and more persistent output losses (around -20% to -30%) well beyond ten years. In this regard, the sovereign default case is notable because, while its initial impact is larger than a typical banking crisis, output losses in our sample tend to revert toward zero. This pattern—default costs appearing more persistent once spillovers are accounted for—is consistent with the idea that, rather than sustained fiscal austerity, defaulting countries can obtain immediate debt relief or renegotiate terms, and that accompanying currency depreciation (the "Twin Ds") can help support exports and buffer output losses—the Mundell-Fleming mechanisms which are less available in typical banking crises.

7.2 Connections to Existing Literature

To replicate the high levels of external debt observed before default, theoretical models often include an ad-hoc direct output loss from sovereign default. Although my causal estimates—8% output loss in the first year, peaking at -18.5% by the year 2—significantly exceed -2% output loss parameter typically assumed in earlier default models (e.g., Aguiar and Gopinath 2006; Yue 2010), they align well with the approximately -10% output drops calibrated in more recent endogenous default models (e.g., Mendoza and Yue 2012; Arellano 2008).

In contrast, my causal estimates are notably higher than those reported in previous empirical studies. Early research using unbalanced panels with fixed effects or GMM regressions, such as Borensztein and Panizza (2008) and De Paoli et al. (2009), finds output loss ranging from −2.6% to −5%. However, using similar methodologies, Furceri and Zdzienicka (2012) report significantly higher and more persistent output loss at around −10% over eight years. This disparity highlights how default cost estimates can vary significantly depending on methodology used and sample selection. In addition to panel fixed effects models, Kuvshinov and Zimmermann (2019) apply local projection–inverse propensity score weighting (LP-IPSWRA) to estimate a first-year output loss of −2.7%, peaking at −3.7% within five years. Hébert and Schreger (2017) exploit exogenous variations around of legal decisions on 2001 Argentina's debt restructuring in a simultaneous equation model and find increase in default probability can significantly reduce the equity value of domestic firms, suggesting a potential channel through which sovereign default can lead to output loss.

It is particularly relevant to compare my findings with those of Farah-Yacoub et al. (2024), not only because we address similar questions, but also because we reach comparable conclusions despite using different methodologies and samples. Farah-Yacoub et al. (2024) apply local projection combined with synthetic control methods—analyzing counterfactual scenarios with varying control units—on historical data from 1815 to 2020. They show an −8.5% decline in GDP per capita in the first three years following a debt default, with negative effects persisting for up to 20 years. Despite differences in methods and sample period (1970-2010), my LP-SSIV estimates are strikingly similar to their results. To be specific, while the baseline LP-SSIV estimates initially show non-persistent output loss, accounting for spillover effects with either control function (Panel (b) in Figure 4) or difference-in-difference (Figure 5) aligns well with the conclusion in Farah-Yacoub et al.

(2024) that the cost of default is both significant and enduring. This consistency across various methods supports the unresolved yet widely held view that sovereign default is indeed costly.

8. State-dependent LP-SSIV Regressions

In this section, I estimate the heterogeneous costs of sovereign default across countries based on (1) exchange rate regimes (peg versus float), (2) external debt burden (debt-to-GDP ratio), (3) maturity mismatch (short-term versus long-term debt), and (4) reserve adequacy (whether reserves cover three months of imports). The key takeaway is that, while somewhat cliché, adopting a floating exchange rate regime, maintaining lower levels of external debt—especially short-term debt—and ensuring debt repayment when resources allow can substantially reduce the cost of default.

To be specific, for $h = 0,1,\ldots,6$, I run the following state-dependent LP-SSIV regression:

$$\begin{aligned} y_{it+h} - y_{it-1} = I_{it} \times [\alpha_{Aih} + \beta_{Ah}\hat{D}_{it} + \gamma_{Ah}\mathbf{X}_{it}] \\ +(1 - I_{it}) \times [\alpha_{Bih} + \beta_{Bh}\hat{D}_{it} + \gamma_{Bh}\mathbf{X}_{it}] + v_{it+h} \end{aligned} \quad (6.1)$$

where I_{it} takes value of 1 if country i is in economic state A with the coefficient estimates $\{\beta_{Ah}\}_h$ being the impulse response. Otherwise, $\{\beta_{Bh}\}_h$ will be the impulse response for countries *not* in economic state A (i.e., state B).

Starting with the role of exchange rate regimes, Panel (d) in Figure 4 shows that defaulting under a fixed exchange rate regime leads to a significantly larger output loss—averaging −35% in the first year—compared to defaulting under a floating exchange rate regime, though these negative impacts eventually dissipate within six years.[19] In contrast,

[19] Following the classification scheme developed by Ilzetzki et al. (2019), countries with a coarse code of 1 or 2 are categorized as having fixed exchange regimes, while those with codes 3, 4, or 5 are categorized as

defaults under a floating exchange rate regime generally yield insignificant coefficients. This provides additional empirical evidence for the notion that adopting a floating exchange rate regime can offer better protection against external shocks.

Second, Panel (e) in Figure 4 shows that countries with a higher debt burden—those with a debt-to-GDP ratio above the annual median—suffer a larger and more persistent output loss when they default. The coefficient estimates for these highly indebted countries are consistently negative across all time horizons, in contrast to those for countries with lower debt-to-GDP ratios. This result is not surprising, as heavily indebted countries are often required to implement more stringent and prolonged fiscal austerity measures when they negotiate their external debts with international creditors.

Third, Panel (f) of Figure 4 shows that countries defaulting with relatively more short-term debt—debt with maturities of less than one year—experience a notable increase in output loss. Surprisingly, countries defaulting with relatively less short-term debt, as compared to the annual median, also face substantial output loss, though with larger standard errors. It is reasonable to expect that owing a higher proportion of short-term debt worsens economic conditions, as countries may be forced to fire sale their domestic assets quickly to meet immediate obligations. However, it is somewhat counterintuitive that the output loss for countries with less short-term debt is even more severe, though not statistically significant. One possible explanation is that, in contrast to the short-term pain, countries with more long-term debt might need to endure prolonged fiscal austerity measures, which can further aggravate the cost of default.

having floating exchange rate regimes. Countries with code 6, which denotes dual exchange rate markets, are excluded from the sample.

Last but not least, Panel (g) in Figure 4 shows that countries with adequate reserves—that is, reserves sufficient to cover at least three months of imports—face disproportionately larger output losses when they default. While accumulating reserves is crucial for maintaining exchange rate stability, defaulting despite an apparent ability to repay may further damage the debtors' reputation with global creditors, leading to more severe consequences. Consistent with this pattern, Panel (h) in Figure 4 shows that countries defaulting under any IMF arrangement, as coded by Vreeland (2007), tend to experience more negative output losses, potentially because the IMF involvement signals to markets that the default episode is particularly severe. These two results highlight the need for the IMF to help bridge the information gaps between defaulting countries and global creditors, ensuring transparency and demonstrating that maintaining sufficient reserves, as well as engaging constructively with the IMF, is essential for domestic economic stability that will eventually benefit both parties.

9. Concluding Remarks

How much does it cost if a country defaults on its external debts? Creditors in the sovereign debt market face "limited commitment"—it is difficult to enforce contracts, so concerns over moral hazard naturally arise. In reality, sovereign defaults happen only occasionally, so why do countries repay their debt? One explanation—central in theoretical work seeking to match realistic debt level on the eve of default—is that defaulting countries suffer direct output loss, an ad hoc assumption for which a credible causal estimate has not yet been firmly established.

In this paper, I introduce a novel LP-SSIV approach to estimate the causal effect of sovereign default on output loss, leveraging aggregate variation in developing countries' endogenous currency denominations of external debt together with advanced economies' quasi-random monetary policies, proxied by both interest rates and narrative monetary shocks. I find that sovereign default reduces real GDP per capita by -8% in the first year. The output loss peaks at -18.5% in the second year; persists through the fourth year; and gradually fades by the sixth year. After accounting for the positive spillover effects using either the control function or difference-in-difference frameworks, the estimated output loss becomes larger and more persistent—approximately -30% even after a decade. This finding aligns well with existing literature and further confirms the substantial cost of sovereign default.

Although developing countries may not be able to fully shield themselves from the quasi-random monetary policies of advanced economies, adopting conventional strategies—transitioning from a pegged to a floating exchange rate regime and avoiding excessive accumulation of debt, especially short-term debt—can substantially reduce the negative impact of sovereign defaults.

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Table 1. Currency compositions of external debts and foreign base interest rates

	Currency shares			LT interest rates		
	(1)	(2)	(3)	(4)	(5)	(6)
	Mean	SD	P90	Mean	SD	P90
Panel (a): Major named currencies						
U.S. dollar	50.1	23.8	83.3	6.2	3.1	10.6
Japanese yen	5.1	8.9	15	3.8	3.1	8.1
British pound (sterling)	3.5	10.3	8.7	7.4	3.8	12
Swiss franc	0.9	4.1	2	3.4	2	5.6
Deutsche mark (pre-1999)	5.3	7.9	13.3	7.3	1.5	9
French franc (pre-1999)	6.9	13.1	21	9.2	2.8	13.1
Euro*	13.8	18.5	39.1	2.5	1.7	4.5
Panel (b): Miscellaneous categories						
Special Drawing Right	1.9	4.6	5.7			
Multiple currency	10.2	12.2	26.3			
Other currency	13.2	16.2	35.7			
Panel (c): Main vs. miscellaneous						
Named currencies (sum of panel (a))	72.4	19.6	96.6			
Miscellaneous (sum of panel (b))	25.2	19	51.3			

Notes: This table summarizes country-year-currency varying shares of long-term public and publicly guaranteed debt in developing countries and also interest rate movements among advanced economies during the 1970–2020 period. The "Euro*" category aggregates debt denominated in Deutsche mark, French franc, and other legacy ERM currencies (e.g., Italian lira, Spanish peseta) after 2000. The summary statistics for each currency are computed over the subsample in which that currency is present (e.g., U.S. dollar: 1970–2020; Deutsche mark: 1970–2000; Euro: 2001–2020). The specific currency composition within "other currency" or "multiple currency" is unknown; we therefore treat these categories, together with "special drawing rights", as "incomplete share"—defined as one minus the sum of all named currency shares—and use it as controls in the baseline regressions. Currency denomination data are retrieved from the *International Debt Statistics* provided by the World Bank Group, while interest rate data come from the long-term interest rate series (typically government bond rates) in the *JST Macrohistory database*.

Table 2. Strong first-stage between binary default indicators and $\Delta IRE_{it-\ell}$

Dependent variable: S&P first-year binary (0/1) default indicator					
	(1)	(2)	(3)	(4)	(5)
ΔIRE_{it-1}	0.007			0.009	
	(0.006)			(0.006)	
ΔIRE_{it-2}		0.027***		0.026***	
		(0.006)		(0.006)	
ΔIRE_{it-3}			0.010*	0.010*	
			(0.006)	(0.006)	
IRE_{it}^{USRR}					0.024**
					(0.010)
IRE_{it}^{UKCH}					0.053**
					(0.023)
Observations	3,898	3,898	3,898	3,898	3,235
R-squared	0.077	0.087	0.078	0.089	0.094
Country FE	Y	Y	Y	Y	Y
Baseline controls	Y	Y	Y	Y	Y

Notes: This table shows a strong first-stage relationship between the first-year S&P default dummy and the second lag of the shift-share instrument—the "interest rate exposure" ΔIRE_{it-2}. The first-year default dummy equals 1 if it is the first year a developing country defaults during a particular default episode, as reported by Standard and Poor's (S&P). All regressions include country fixed effects. The baseline control set includes up to two lags of: (i) regression-specific variables such as the treatment variable (e.g., first-year default indicator), one lead and one lag of the shift-share instrument (e.g., ΔIRE_{it-1} and ΔIRE_{it-3} are included as controls for ΔIRE_{it-2}) to address "lead-lag exogeneity" (Stock and Watson 2018), and the "incomplete shares"—defined as one minus the sum of major named currency shares (Borusyak et al. 2022); (ii) continuous variables such as the first differences of log real GDP per capita, log GDP deflator, log nominal exchange rate, and log world GDP; and levels of debt-to-GDP ratios and the Chinn–Ito (2006) capital openness index; and (iii) binary indicators for banking, currency, political crises (with 1 indicating war/coup/political transition), and democracy (Acemoglu et al. 2019). The same control set is used across OLS, IV (both first-stage and 2SLS), and all robustness specifications. Robust standard errors, clustered at the country level to address serial correlation, are reported in parentheses. Statistical significance is denoted by *** p<0.01, ** p<0.05, and * p<0.1.

Table 3. Baseline LP-OLS and LP-SSIV estimates on default cost with ΔIRE_{it-2}

	Output response ($y_{it+h} - y_{it-1}$)		AR test	OLS=IV
	LP-OLS	LP-SSIV	p-value	p-value
	(1)	(2)	(3)	(4)
$h=0$	-2.66***	-7.99	0.12	0.32
	(0.63)	(5.28)		
$h=1$	-3.81***	-18.48**	0.01**	0.07*
	(1.01)	(8.08)		
$h=2$	-4.76***	-15.78*	0.06*	0.20
	(1.14)	(8.60)		
$h=3$	-4.79***	-18.25*	0.05*	0.16
	(1.33)	(9.52)		
$h=4$	-5.28***	-9.47	0.38	0.70
	(1.51)	(10.72)		
$h=5$	-5.21***	-5.70	0.59	0.96
	(1.74)	(10.55)		
$h=6$	-5.17***	-0.52	0.96	0.68
	(1.81)	(11.17)		
Joint significance	0.00	0.10		
Instrument	ΔIRE_{it-2}			
KP weak IV	21.14			
Observations	3,898			
Country FE	Y	Y		
Baseline controls	Y	Y		

Notes: This table presents the baseline LP-OLS and LP-SSIV coefficient estimates of the cost of sovereign default. The sample period is 1970-2010. The dependent variable is the long-difference of the log of real GDP per capita $y_{t+h} - y_{t-1}$; the independent variable is a binary first-year default indicator; and the shift-share instrument is the second lag of "interest rate exposure" ΔIRE_{it-2}. This table also reports p-values from the joint significance test (i.e., whether coefficient estimates across all horizons $(\beta_0, \beta_1, ..., \beta_6)$ are simultaneously zero), Anderson-Rubin F-statistics (weak instrument test), the OLS=IV coefficient test at each horizon (stacking method), as well as the Kleibergen-Paap F-statistic (standard threshold of 10). All regressions include country fixed effects and the baseline controls specified in Table 2. In addition, it includes ΔIRE_{it-1} and ΔIRE_{it-3} as controls to address "lead-lag exogeneity", as well as "incomplete shares" controls. Robust standard errors, clustered at the country level to address serial correlations, are reported in parentheses. Statistical significance is indicated as *** p<0.01, ** p<0.05, * p<0.1.

Table 4. Default cost estimates with narrative shocks IRE_{it}^{USRR} and IRE_{it}^{UKCH}

	Output response $(y_{it+h} - y_{t-1})$		AR test	OLS=IV
	LP-OLS	LP-SSIV	p-value	p-value
	(1)	(2)	(3)	(4)
$h=0$	-2.54***	-8.15	0.68	0.56
	(0.66)	(9.64)		
$h=1$	-3.42***	-20.55	0.34	0.35
	(1.04)	(17.72)		
$h=2$	-4.07***	-8.16	0.63	0.87
	(1.17)	(20.00)		
$h=3$	-3.85***	-2.96	0.76	0.94
	(1.34)	(21.55)		
Joint significance	0.00	0.58		
Instrument	IRE_{it}^{USRR} and IRE_{it}^{UKCH}			
KP weak IV	4.29			
Observations	3,235			
Country FE	Y	Y		
Baseline controls	Y	Y		

Notes: This table presents the LP-SSIV coefficient estimates for the cost of sovereign default using shift-share instruments based on narrative monetary shocks. To be specific, IRE_{it}^{USRR} represents the "interest rate exposure" using endogenous currency denomination shares weighted by the Romer-Romer narrative U.S. monetary shocks, while IRE_{it}^{UKCH} is based on the Cloyne-Hürtgen U.K. monetary shocks (Romer and Romer 2004; Cloyne and Hürtgen 2016). Figure A3 shows that U.K. monetary shocks series is markedly more volatile and exhibits larger shocks in magnitude than its U.S. counterpart. Also, the within-year positives and negatives monthly monetary shocks often offset each other. To mitigate this attenuation, Table 4 defines the annual USRR and/or UKCH monetary shock as the single largest-magnitude monthly shock in each year, standardizes the annual series within each country, and then constructs the narrative shift-share instrument by interacting these standardized shocks with currency shares denominated in U.S. dollars and British pounds respectively. The sample period (1975-2007) in this table differs from that (1970-2010) in Table 3 due to data constraints, resulting in different LP-OLS estimates and numbers of observations as compared to Table 3. The dependent and independent variables, as well as all statistical tests (joint significance, KP weak IV, AR test, OLS=IV test), are consistent with those in Table 3. The control variables include baseline controls from Table 2, plus one lead and one lag of both IRE_{it}^{USRR} and IRE_{it}^{UKCH}, as well as adjustments for incomplete currency shares excluding U.S. dollars or British pounds. Robust standard errors, clustered at the country level to account for serial correlations, are reported in parentheses. Statistical significance is indicated as *** p<0.01, ** p<0.05, * p<0.1.

Table 5. Default cost estimates with sovereign arrears and $\Delta IRE_{it-\ell}$

Panel (a): Strong first-stage relationship between sovereign arrears and $\Delta IRE_{it-\ell}$

Dependent variable	$\Delta \ln\left(\text{arrears}\right)$	$\Delta \ln\left(1+\text{arrears}\right)$	$\Delta \ln\left(1+100\times\frac{\text{arrear}}{\text{debt}}\right)$
	(1)	(2)	(3)
ΔIRE_{it-1}	0.046	-0.004	-0.006
	(0.040)	(0.024)	(0.011)
ΔIRE_{it-2}	0.070*	0.002	-0.003
	(0.041)	(0.030)	(0.015)
ΔIRE_{it-3}	0.126***	0.053**	0.021**
	(0.044)	(0.023)	(0.010)
R-squared	0.092	0.055	0.059

Panel (b): Default cost estimates—extensive and intensive margins

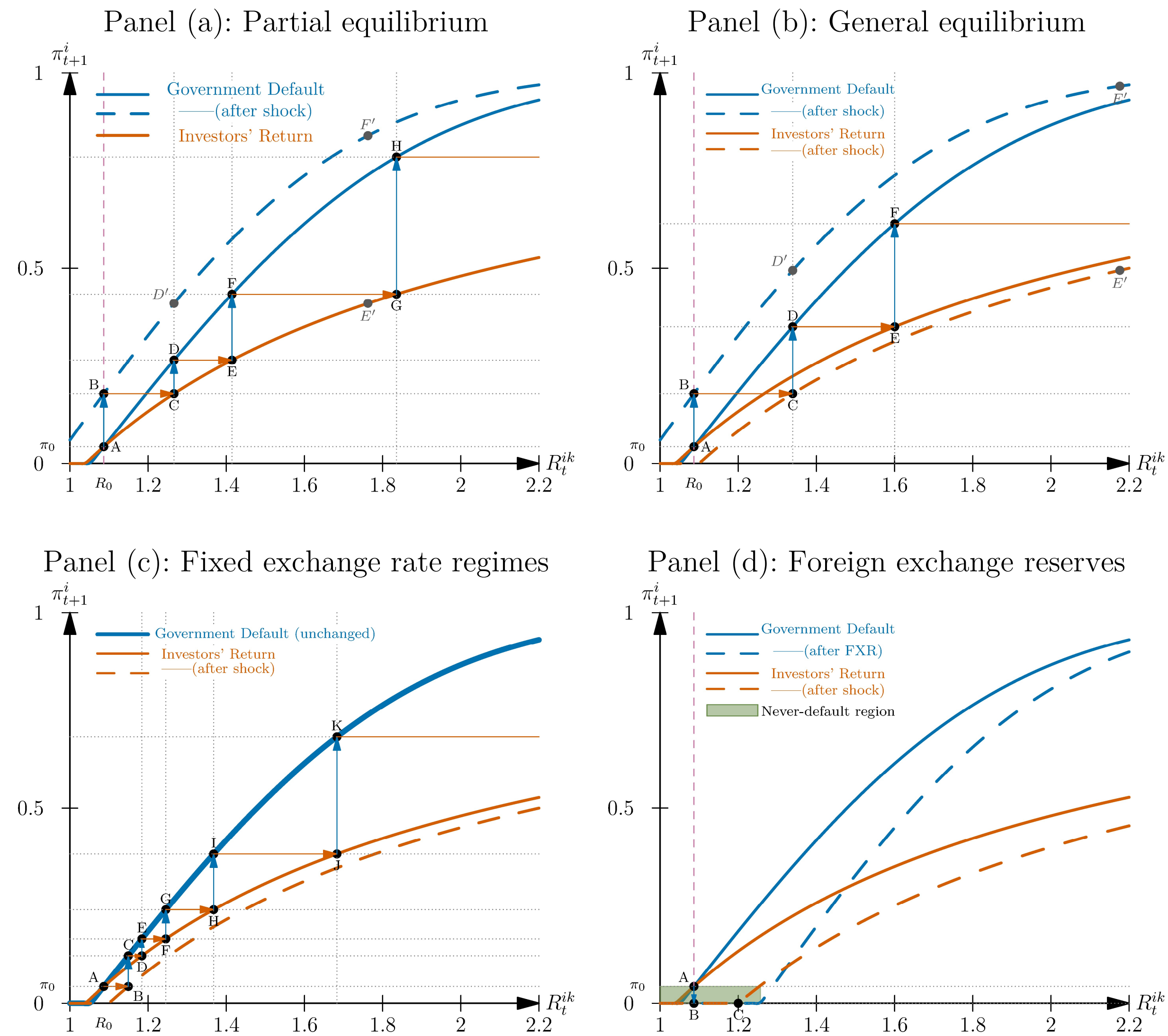
Dependent variable: cumulative loss in real GDP per capita (%)

Independent variable	$\Delta \ln\left(\text{arrears}\right)$		$\Delta \ln\left(1+\text{arrears}\right)$		$\Delta \ln\left(1+100\times\frac{\text{arrear}}{\text{debt}}\right)$	
	LP-SSIV	AR	LP-SSIV	AR	LP-SSIV	AR
	(1)	(2)	(3)	(4)	(5)	(6)
$h=0$	-1.77	0.18	-3.16	0.18	-8.03	0.17
	(1.41)		(2.74)		(7.17)	
$h=1$	-3.02	0.17	-4.04	0.28	-10.42	0.26
	(2.34)		(3.98)		(10.19)	
$h=2$	-5.25*	0.04**	-8.62	0.05**	-21.84	0.04**
	(2.96)		(5.28)		(13.71)	
$h=3$	-4.58	0.14	-7.36	0.14	-18.67	0.13
	(3.40)		(5.57)		(14.05)	
$h=4$	-4.02	0.25	-6.15	0.28	-15.66	0.27
	(3.77)		(5.98)		(14.92)	
$h=5$	-1.25	0.75	1.98	0.76	4.51	0.78
	(3.88)		(6.61)		(16.40)	
Instrument			ΔIRE_{it-3}			
Joint significance	0.1		0.28		0.3	
KP weak IV	8.17		5.45		4.7	
Observations	2,946		4,690		4,690	

Country FE	Y	Y	Y
Baseline controls	Y	Y	Y

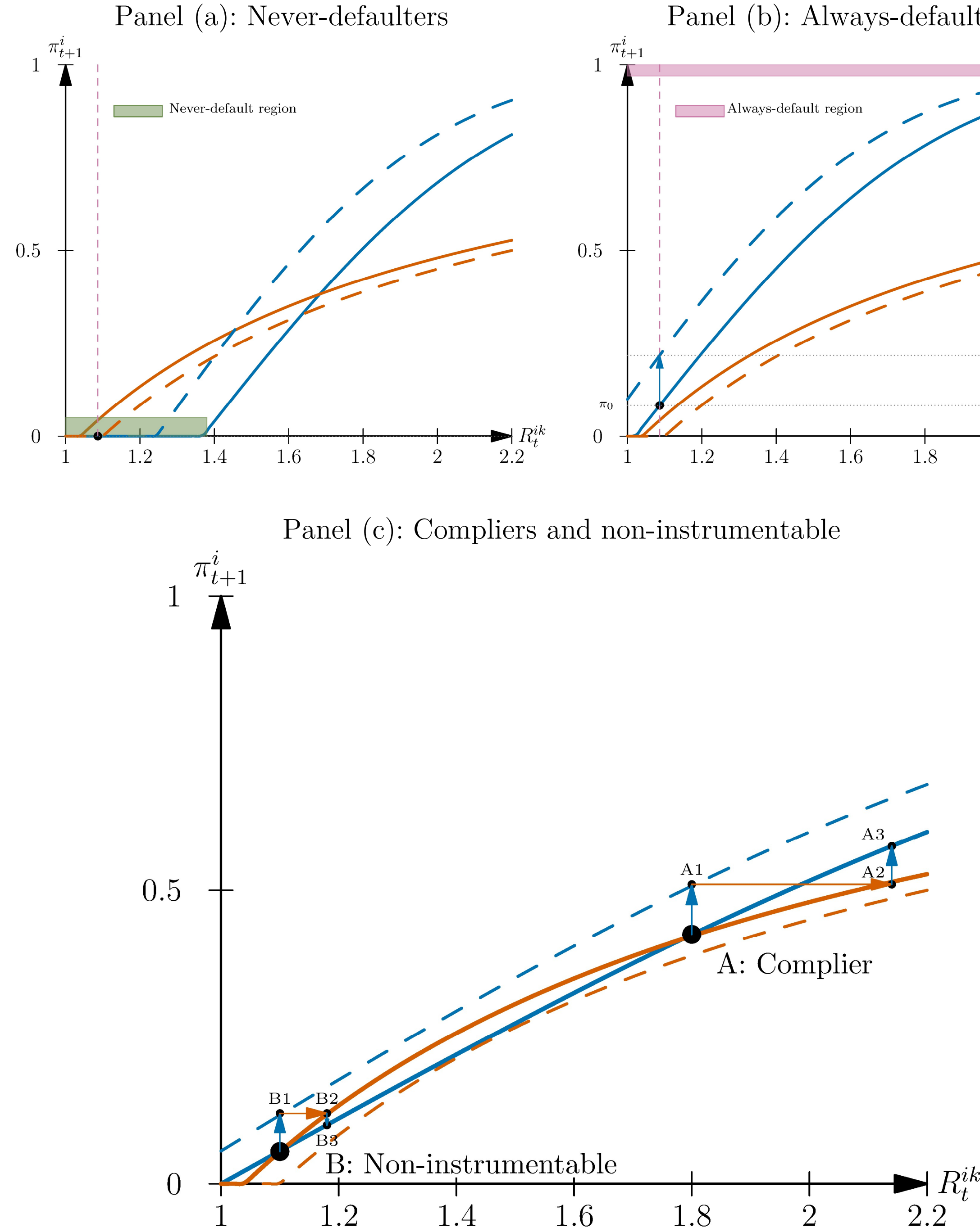
Notes: This table presents the first-stage and LP-SSIV estimates of the output loss from sovereign default using continuous default intensity measures that account for both extensive and intensive margins. The IDS arrears data refer to late repayments in the public and publicly guaranteed (PPG) sector of debtor countries, which sum both principal and interest arrears, use the world aggregate as the counterpart, and admit zeros as well as strictly positive numbers. To distinguish missing entries from real zeros, I compare values across different versions of the IDS database (see Table A1). Following Chen and Roth (2024), column (3) in Panel (a) and columns (5)–(6) are the preferred specifications, as these regressions anchor the interpretation of the extensive and intensive margins via the variable $\frac{\text{arrear}}{\text{debt}}$. In Panel (a), the dependent variable set includes continuous default measures such as $\Delta \ln(\text{arrears})$, $\Delta \ln(1+\text{arrears})$, and $\Delta \ln(1+\frac{\text{arrear}}{\text{debt}}\times 100)$. In particular, column (3) shows that a one-unit increase in ΔIRE_{it-3} is associated with a 0.021 increase in $\frac{\text{arrear}}{\text{debt}}$, which is close to zero and provides a natural default threshold. In Panel (b), the dependent variable is the long-difference of the log of real GDP per capita $y_{t+h}-y_{t-1}$ and is interpreted as a deviation from trend. The shift-share instrument is the third lag ΔIRE_{it-3} for all columns (1)–(6). The sample period is 1970–2020. The joint significance test evaluates whether coefficient estimates $(\beta_0,\beta_1,\ldots,\beta_5)$ across all horizons are jointly zero. The Kleibergen-Paap F-statistic for weak instruments (KP weak IV) is commonly used with a standard threshold of 10. The p-value for the Anderson-Rubin test is calculated from the Anderson-Rubin F-statistic. All regressions include country fixed effects and baseline controls specified in Table 2, as well as two leads ΔIRE_{it-1} and ΔIRE_{it-2} to account for lead-lag exogeneity and incomplete shares for currency compositions that do not sum to one (Stock and Watson 2018; Borusyak et al. 2022). Robust standard errors, clustered at the country level to address serial correlation, are reported in parentheses. Statistical significance is indicated as *** p<0.01, ** p<0.05, * p<0.1.

Figure 1. Multiple equilibria in the probabilistic sovereign default model



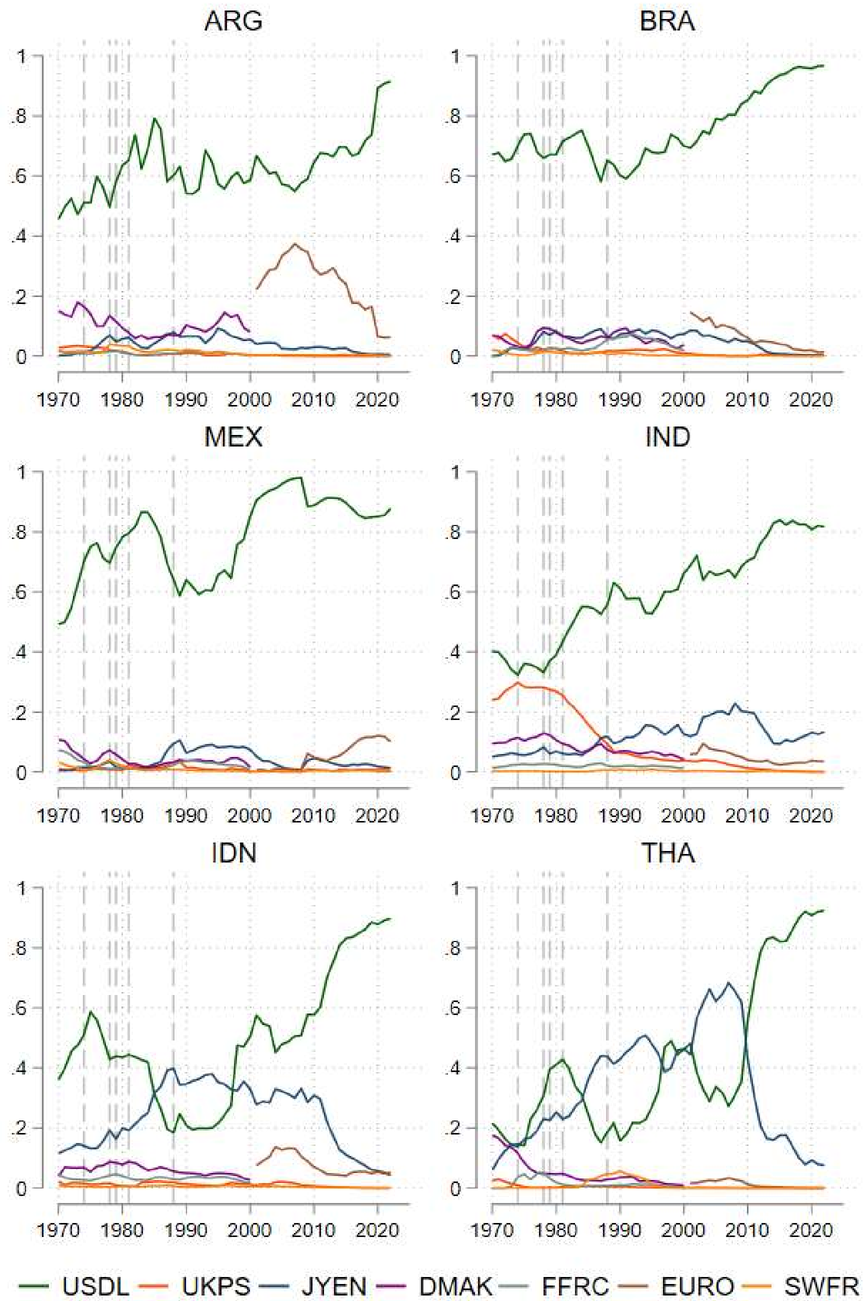
Notes: This figure illustrates multiple equilibria in the probabilistic sovereign default from Section 2. The vertical axis is the default probability; the horizontal axis is the gross rollover interest factor. "Shock" denotes a foreign monetary tightening; solid curves are pre-shock and dashed curves are post-shock. Panel (a) shows that holding investors' return schedule fixed, the shock raises peso-value of dollar-debt obligation, shifting the default curve up. Panel (b) shows that an increase in default risk premium and dollar borrowing costs can generate a self-fulfilling move to the default equilibrium. Panel (c) shows that while fixed exchange rate regimes keep default curve unchanged, higher dollar rates shift investors' return schedule up, raising default risk and hence the risk premium. Panel (d) shows that reserves accumulation shifts default curve down by providing additional dollar resources unaffected by peso depreciation, thereby enlarging the never-default region (shaded in green).

Figure 2. Multiple equilibria and local average treatment effect



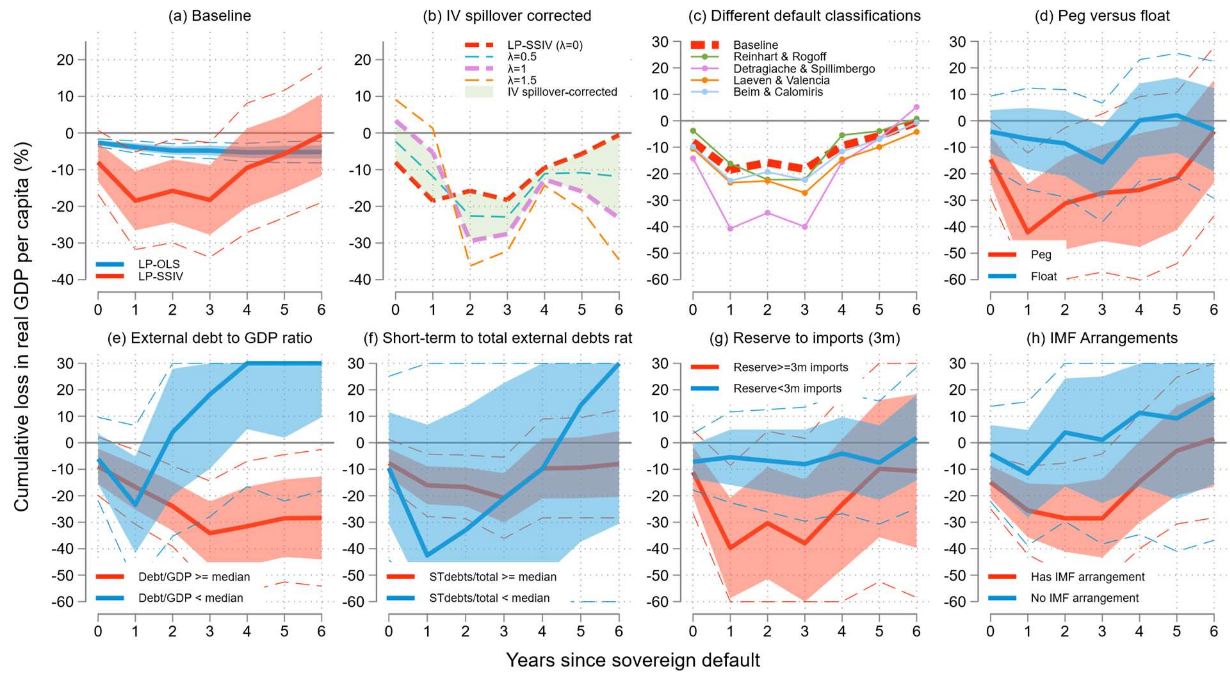
Notes: This figure classifies initial steady states in the probabilistic sovereign default model in Section 2 into four scenarios: (a) never-defaulters: intersections lie on the horizontal axis pre- and post-shock; (b) always-defaulters: intersections lie on the vertical axis pre- and post-shock; (c) compliers: interior intersections move monotonically after shock, identifying the local average treatment effect. (d) non-instrumentable: either non-monotonic response to shock or defaults are driven by factors other than the foreign monetary shock.

Figure 3. Overall trends in currency denomination for selected countries



Notes: This figure visualizes the currency denomination trends of selected countries over time. The data are based on the public and publicly guaranteed debt from the World Bank's International Debt Statistics (IDS). The currencies shown are USDL (U.S. dollar), UKPS (pound sterling), JYEN (Japanese yen), DMAK (Deutsche mark), FFRC (French franc), EURO (replacement for Deutsche mark and French franc post-2000), and SWFR (Swiss franc). Gray vertical dashed lines indicate years with narrative contractionary U.S. monetary policy shocks classified by Romer and Romer (2023). Two takeaways: (1) differential exposure (e.g., greater yen use in some Asian borrowers) and (2) sticky currency shares.

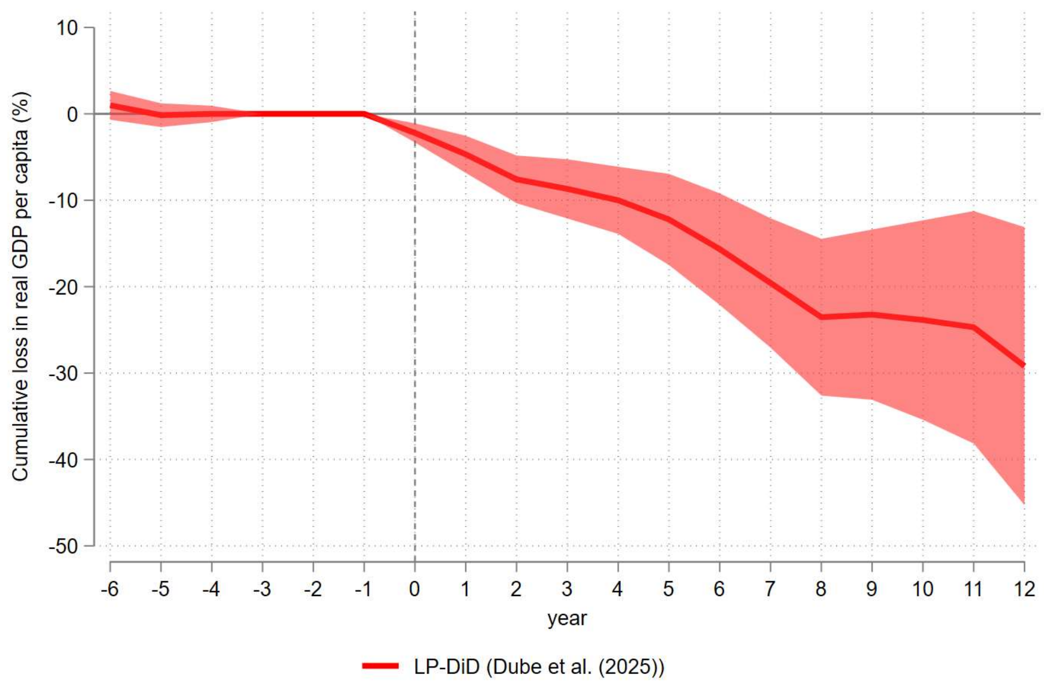
Figure 4. Baseline results, different default classifications, and state-dependent cost



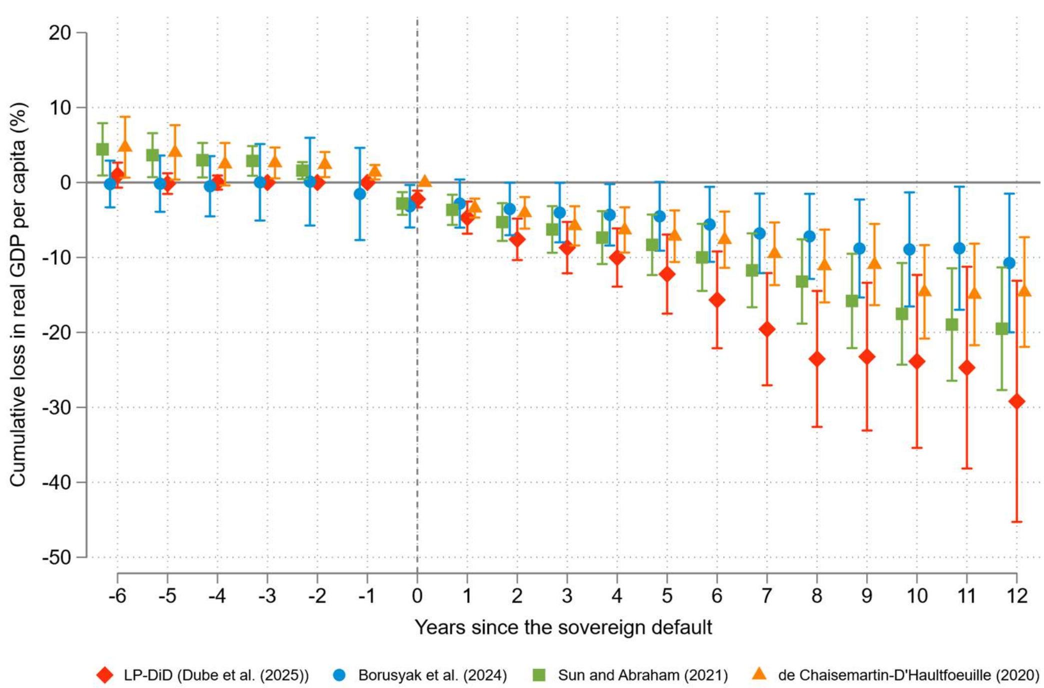
Notes: This figure presents the baseline and state-dependent LP-SSIV estimates of the cost of sovereign default. The solid lines report the point estimates; the shaded areas are one-standard-deviation (68%) confidence intervals; and dashed lines are the 90% confidence intervals. Panels (a) and (b) use an output loss scale from −40% to 20%, while Panels (c)-(d) use a scale from −60% to 30%. Panel (a) shows the baseline LP-OLS and LP-SSIV coefficient estimates reported in Table 3. Panel (b) shows the causal IV spillover-corrected estimates using the control function approach, with the green shaded area (i.e., $\lambda = 1$) presenting the case in which defaulting countries experience the same magnitude of spillover as the never-defaulting countries. Panel (c) illustrates robustness checks based on different default classifications from the Sovereign Default Dataset by Kuvshinov and Zimmermann (2019). Panel (d) shows that countries with pegged exchange rate regimes (coarse codes 1 or 2 in Ilzetzki et al. (2019)) face a more severe cost of default. Panel (e) indicates that countries with higher external debt-to-GDP ratios (above the annual median) experience a more pronounced cost of default. Panel (f) shows that countries with higher short-term debt levels (above the annual median) likewise incur a greater cost of default. Panel (g) shows that defaulting countries that have sufficient reserves (greater than 3 months of imports) face a disproportionately larger negative impact. Panel (h) shows that countries defaulting under any IMF arrangement tend to experience greater output loss (Vreeland 2007). All regressions include country fixed effects and control variables as specified in Table 3. Robust standard errors are clustered at the country level to address serial correlation.

Figure 5: Estimating default cost with difference-in-difference estimators

Panel (a): Local projection—difference-in-difference



Panel (b): Alternative difference-in-difference estimators



Notes: This figure plots the estimated cost of sovereign default using various difference-in-difference (DiD) estimators that address staggered treatment in the literature. All DiD specifications include country and year fixed effects, as well as the baseline controls specified in Table 3. The LPDiD specification uses the long-difference in log real GDP per capita as the dependent variable and first-differences of controls variables (up to two lags). Assuming the effects of negative weights from repeated defaults (i.e., contamination effects) dissipate after 6 years, the clean control set is defined as a 6-year window, restricting the sample to countries that have not defaulted in the past 6 years. Following the standard practice in the literature, the other DiD estimators use the level of log real GDP per capita as the dependent variable, include only lagged (not contemporaneous) controls in levels, and exclude lagged outcome. Robust standard errors are clustered at the country level, and 90% confidence intervals are shown in the figure. See Figure A5 for additional results with Callaway and Sant'Anna (2021) and the main text for reference on these DiD estimators.

Monetary Shocks, Currency Exposures, and the Cost of Sovereign Default

Online Appendix

Jeff Kin Wai Cheung (UC Davis)

Appendix

Additional Tables

Additional Figures

A1. Closed-form Steady State Default Probability Under Log-Normal Distributions

Proposition (1) follows almost immediately from the fact that $\ln(\bullet)$ is an injective function on the strictly positive domain and the properties of normal distribution. Fix an arbitrary debtor *I*, currency *k*, and period $t+1$, and suppress time subscript $t+1$ when clear for simplification. Assume $\left(\mathcal{T}^{i}, \mathcal{M}^{k}\right)$ are both strictly positive and mutually independent random variables, then.

Lemma (A1): If $\mathcal{T}^{i} > 0, \mathcal{M}^{k} > 0$ and $\mathcal{T}^{i} \perp \mathcal{M}^{k}$, then $\ln \mathcal{T}^{i} \perp \ln \mathcal{M}^{k}$

Proof: This is a direct application of integration by substitution and using that the fact that $\ln(\bullet)$ is an *injective* function at the *strictly positive* domain. It is sufficient to show that for any mutually *independent* random variables $X, Y > 0; X \perp Y$, then we have $\ln X > \ln Y$. First, let $U \equiv \ln X, V \equiv \ln Y$, so we have $X = \exp(U), Y = \exp(V)$. Therefore, the Jacobian matrix is given as:

$$J \equiv \begin{pmatrix} {}^{\partial X}\!/_{\partial U} & {}^{\partial X}\!/_{\partial V} \\ {}^{\partial Y}\!/_{\partial U} & {}^{\partial Y}\!/_{\partial V} \end{pmatrix} = \begin{pmatrix} \exp(U) & 0 \\ 0 & \exp(V) \end{pmatrix} \Rightarrow \det(J) = \exp(U) \cdot \exp(V)$$

The above result shows that the Jacobian determinant is well-defined. Since the $X \perp Y$ implies $f_{X \times Y}(x, y) = f_X(x) f_Y(y)$ for all x, y in the domain, we have

$$\begin{aligned} g_{U \times V}(u, v) &= f_{X \times Y}(\exp(u), \exp(v)) \det(J) \\ &= f_X(\exp(u)) \cdot f_Y(\exp(v)) \cdot \exp(u) \cdot \exp(v) \\ &= \underbrace{f_X(\exp(u)) \exp(u)}_{g_U(u)} \cdot \underbrace{f_Y(\exp(v)) \exp(v)}_{g_V(v)} \\ &= g_U(u) \cdot g_V(v) \qquad \blacksquare \end{aligned}$$

To derive a closed-form solution for the steady state π_0^i, assume $\left(\mathcal{T}^i, \mathcal{M}^k\right)$ are mutually independent random variables drawn from log-normal distributions with different parameters:

Lemma (A2): If $\mathcal{T}^i \sim \text{Lognormal}\left(\tau_{ti}, \sigma_{ti}^2\right), \mathcal{M}^k \sim \text{Lognormal}\left(\mu_{mk}, \sigma_{mk}^2\right)$, and $\mathcal{T}^i \perp \mathcal{M}^k$, then we have $\mathcal{Z}^{ik} \equiv \ln \mathcal{T}^i \mathcal{M}^k \sim \text{Normal}\left(\mu_{ti} + \mu_{mk}, \sigma_{ti}^2 + \sigma_{mk}^2\right)$

Proof: By definition of the log-normal distribution, we have $\ln \mathcal{T}^i \sim \text{Normal}\left(\mu_{ti}, \sigma_{mk}^2\right)$ and likewise $\ln \mathcal{M}^k \sim \text{Normal}\left(\mu_{mk}, \sigma_{mk}^2\right)$. By **Lemma (A1)**, since $\mathcal{T}^i \perp \mathcal{M}^k$, we have $\ln \mathcal{T}^i \perp \ln \mathcal{M}^k$, which means they are mutually independent of each other. Therefore, by the property of normal distribution that the sum of two independent normally distributed random variables still follow normal distribution, we have

$$\mathcal{Z}^{ik} \equiv \ln \mathcal{T}^i \mathcal{M}^k = \ln \mathcal{T}^i + \ln \mathcal{M}^k \sim \text{Normal}\left(\mu_{ti} + \mu_{mk}, \sigma_{ti}^2 + \sigma_{mk}^2\right) \qquad \blacksquare$$

The mutually independent assumption is a strong assumption because it places restriction on the entire joint probability density function (PDF) of two random variables (e.g., $X \perp Y \Leftrightarrow f_{X \times Y}(x, y) = f_X(x) f_Y(y)$). Therefore, observing $\text{cov}(\mathcal{T}_i', \mathcal{M}_j') = 0$ in the data is a necessary, but not a sufficient condition, for them to be mutually independent.

Lemma (3): If $\mathcal{Z}^{ik} \equiv \ln \mathcal{T}^i \mathcal{M}^k \sim \text{Normal}\left(\mu_{ti} + \mu_{mk}, \sigma_{ti}^2 + \sigma_{mk}^2\right)$, then

$\dfrac{\mathcal{Z}^{ik} - (\mu_{ti} + \mu_{mk})}{\sqrt{\sigma_{ti}^2 + \sigma_{mk}^2}} \sim \text{Normal}\left(0,1\right)$. Hence,

$$\mathbb{P}\left(\mathcal{Z}^{ik} \leq z\right) = \mathbb{P}\left(\frac{\mathcal{Z}^{ik} - (\mu_{ti} + \mu_{mk})}{\sqrt{\sigma_{ti}^2 + \sigma_{mk}^2}} \leq \frac{z - (\mu_{ti} + \mu_{mk})}{\sqrt{\sigma_{ti}^2 + \sigma_{mk}^2}}\right) = \Phi\left(\frac{z - (\mu_{ti} + \mu_{mk})}{\sqrt{\sigma_{ti}^2 + \sigma_{mk}^2}}\right)$$

where $\Phi\left(\cdot\right)$ is the standard-normal cumulative density function (CDF).

Last but not least, assume domestic money supply growth at a constant rate and an incomplete passthrough of foreign monetary policy on exchange rate, that is, $\mathcal{E}^{ik} = \overline{\sigma}^{ik} \frac{\bar{\mathcal{M}}^{i}}{\mathcal{M}^{k}}$ where $\overline{\sigma}^{ik}$ measures how responsive peso depreciation to foreign monetary policy. Since there are only two countries in the world, domestic debtor i and foreign country k, we have:

$$
\begin{aligned}
\pi_0^i &\equiv \mathbb{P}\left(\mathcal{T}^i \le \overline{R}^{ik}\mathcal{E}^{ik}\right) \\
&= \mathbb{P}\left(\mathcal{T}^i \le \overline{R}^{ik} \cdot \overline{\sigma}^{ik} \frac{\bar{\mathcal{M}}^i}{\mathcal{M}^k}\right) \\
&\overset{(*)}{=} \mathbb{P}\left(\ln \mathcal{T}^i \mathcal{M}^k \le \ln \overline{R}^{ik}\overline{\sigma}^{ik}\bar{\mathcal{M}}^i\right) \\
&= \Phi\left(\frac{\ln \overline{R}^{ik}\overline{\sigma}_i^{ik}\bar{\mathcal{M}}^i - (\mu_{ti} + \mu_{mk})}{\sqrt{\sigma_{ti}^2 + \sigma_{mk}^2}}\right) \\
&= \Phi\left(\frac{\ln \overline{\sigma}_i^{ik} + \left(\ln \overline{R}^{ik} - \mu_{ti}\right) + (\ln \bar{\mathcal{M}}^i - \mu_{mk})}{\Xi}\right) \qquad \text{where} \quad \Xi \equiv \sqrt{\sigma_{ti}^2 + \sigma_{mk}^2} > 0 \\
&= \Phi\left(\frac{\ln \overline{\sigma}_i^{ik}}{\Xi} + \frac{-\mathbb{E}\left[\ln \frac{\mathcal{T}^i}{D^i}\right]}{\Xi} + \frac{\mathbb{E}\left[\ln \bar{\mathcal{M}}^i\right] - \mathbb{E}\left[\ln \mathcal{M}^k\right]}{\Xi} + \frac{\ln \overline{R}^{ik}}{\Xi}\right)
\end{aligned}
$$

Notice that since $\ln \mathcal{M}_{t+1}^k = \ln M_{t+1}^k - \ln M_t^k \approx \%\Delta M_{t+1}^k$, and if $\frac{T^i}{D^i} \approx 1$, we have $\ln \frac{\mathcal{T}^i}{D^i} \approx \frac{\mathcal{T}^i}{D^i} - 1 \propto \left(\mathcal{T}^i - D^i\right)$. Therefore, $\mathcal{T}_i'$ serves as a natural measure of *fiscal capacity* on whether on average debtor I's revenue exceeds its debt obligation. Similarly, if $R^{ik} \approx 1$, then $\ln R^{ik} \approx R^{ik} - 1 = r^{ik}$, where r^{ik} is the net interest rate.

The probability expression in Equation (2.7) can be decomposed into four economically meaningful components. The first term, $\ln \overline{\sigma}^{ik}$, captures the sovereign i's elasticity of exchange rate to foreign monetary shocks—a stronger exchange rate response to foreign monetary shocks ($\overline{\sigma}^{ik}$) magnifies the peso value of debt obligation and thus pushes default probability upwards.

The second term, $-\mathbb{E}\left[\ln \dfrac{\mathcal{T}^i}{D^i}\right]$, reflects the sovereign *i's fiscal capacity*. When the ex-post revenue $\mathcal{T}^i_{t+1}$ in period $t+1$ *systematically* exceeds ex-ante required debt D^i_t, this term becomes negative and default probability falls, vice versa. Since it depends on the entire *joint* distribution of the revenue-to-debt ratio rather than on occasionally windfalls (e.g., commodity booms), only fundamental reforms towards more prudent fiscal policy, such as broadening the tax base and reducing government expenditure, can reduce the likelihood of default.

The third term, $\mathbb{E}\left[\ln \mathcal{M}^i - \ln \mathcal{M}^k\right]$, measures sovereign *i*'s *monetary prudence*, that is, the gap between the expected domestic and/or foreign money growth rates. Rapid domestic monetary expansion depreciates the peso, making foreign-currency debt harder to service and increasing default probability. This mechanism is consistent with the sharp decline in default episodes after many emerging markets adopted the inflation-target framework, which effectively aligns domestic money growth with the 2 percent inflation rate in many advanced economies.

Last but not least, the fourth term, $\ln R^{ik}$, represents the *interest burden* on external debt. While its direct effect is intuitive (i.e., higher interest leads to higher default probabilities), The next session shows that the default probabilities themselves feed back into rollover rates of debt into the next period. Under adaptive expectations, even a small upward revision in perceived default risk can sharply raise short-term borrowing cost and drive the economy towards a self-fulfilling default equlibirum.

To plot the government default function with default probability on the vertical axis and real interest rate charged on the horizontal axis, we can rewrite the above equation:

$$\pi^i = \Phi\left(\frac{\ln \bar{\sigma}^{ik}}{\Xi} + \frac{-\mathbb{E}\left[\ln \frac{T_i'}{D_i}\right]}{\Xi} + \frac{-\mathbb{E}\left[\ln \mathcal{M}^k\right]}{\Xi} + \frac{1}{\Xi}\left(R^{ik} - 1\right)\right) \quad \text{if} \quad R^{ik} \approx 1$$

A2. An Eaton and Kortum (2002) Probabilistic Approach to Currency Shares

Assume a continuum of loan suppliers $j \in [0,1]$ and a countable set of currencies $k \in \mathbb{N}$. Within each currency k, perfect competition implies that all lenders charge the same gross interest rate $R_t^{ik} \geq 1$—determined at t and due at $t+1$—to debtor i. Fix an arbitrary period $t+1$ and a debtor i, the gross depreciation factor against currency k, denoted as $\mathcal{E}_{t+1}^{ik}$, is drawn independently across currency from Fréchet distributions with the same debtor-specific shape parameter θ^i but different currency-specific location parameter $\overline{e}^k$. For model closure, I assume debtor i's tax revenue also follows a Fréchet distribution with the same shape parameter θ^i but debtor-specific location parameter $\overline{t}^i$. For simplicity, suppress subscripts when unambiguous and relabel $\left(\pi_{t+1}^i, \overline{e}^k, \theta^i,, \overline{t}^i, R_t^{ik}, \mathcal{E}_{t+1}^{ik}, \mathcal{T}_{t+1}^i\right) \to \left(\pi, \overline{e}_k, \theta, \overline{t}, R_k, \mathcal{E}_k, \mathcal{T}\right)$, so $\forall k \in \mathbb{N}$, $\mathcal{E}_k \sim \text{Frechet}(\overline{e}_k, \theta)$ and also $\mathcal{T} \sim \text{Frechet}(\overline{t}, \theta)$. From Equation (2.8), default occurs when

$$\begin{aligned}
\pi &\equiv \mathbb{P}\left(\mathcal{T} \leq \max_k \{R_k \mathcal{E}_k\}\right) \\
&= \int_0^\infty \mathbb{P}\left(\exists k : \tau \leq R_k \mathcal{E}_k\right) dG_{\mathcal{T}}(\tau) \\
&= \int_0^\infty \left(1 - \mathbb{P}\left(\forall k : \tau > R_k \mathcal{E}_k\right)\right) dG_{\mathcal{T}}(\tau) \\
&= 1 - \int_0^\infty \left(\prod_k \mathbb{P}\left(\mathcal{E}_k < \frac{\tau}{R_k}\right)\right) dG_{\mathcal{T}}(\tau) \\
&= 1 - \int_0^\infty \exp\left(-\tau^{-\theta} \sum_k \overline{e}_k R_k^\theta\right) \cdot \left(-\overline{t} \exp\left(-\overline{t} \cdot \tau^{-\theta}\right)\right) d(\tau^{-\theta})
\end{aligned}$$

Let $u \equiv \tau^{-\theta}$, $du = d\left(\tau^{-\theta}\right) = -\theta \tau^{-\theta-1}$. When $\tau \to 0, u \to \infty$; when $\tau \to \infty, u \to 0$. Also, define $\Phi \equiv \overline{t} + \sum_k \overline{e}_k R_k^\theta$. The above equation can be further simplified to:

$$
\begin{aligned}
\pi &= 1-(-\overline{t})\int_{\infty}^{0}\exp\left(-\Phi u\right)du \\
&= 1-\overline{t}\int_{0}^{\infty}\exp\left(-\Phi u\right)du \\
&= 1-\overline{t}\left[-\frac{1}{\Phi}\exp\left(-\Phi u\right)\right]_{0}^{\infty} \\
&= 1-\frac{\overline{t}}{\overline{t}+\sum_k \overline{e}_k R_k^{\theta}} \\
&= \frac{\sum_k \overline{e}_k R_k^{\theta}}{\overline{t}+\sum_k \overline{e}_k R_k^{\theta}}
\end{aligned}
$$

Therefore, relabel the variable names gives:

$$
\pi_{t+1}^{i} = \frac{\sum_k \overline{e}^{k}\left(R_t^{ik}\right)^{\theta^i}}{\overline{t}^{\,i}+\sum_k \overline{e}^{k}\left(R_t^{ik}\right)^{\theta^i}}
$$

For currency denomination choice, fix an arbitrary currency \$. A debtor will borrow in \$ if its implied debt service cost is less than or equal to the minimum across all other currencies, that is, $R^{i\$}\mathcal{E}^{i\$} \le \min_{k\neq\$}\left\{R^{k}\mathcal{E}^{ik}\right\}$. Since there is a continuum of creditors $j\in[0,1]$ supplying currency \$ in perfect competition, whether or not the debtor i borrows in currency \$ from creditor j follows a Bernoulli distribution with the probability density function:

$$
1^{ij}(\$) = \begin{cases} 1 & \text{with } \mathbb{P}\left(R^{i\$} \le \min_{k\neq\$}\{R^{ik}\mathcal{E}^{ik}\}\right) \\ 0 & \text{with } 1-\mathbb{P}\left(R^{i\$} \le \min_{k\neq\$}\{R^{ik}\mathcal{E}^{ik}\}\right) \end{cases}
$$

To invoke the law of large numbers, I further assume nominal depreciation factor across different currencies follows identical and independent Fréchet distribution with the same debtor-specific locational parameter $\overline{e}^{i}$ and same shape parameter θ^{i} as before. Henceforth, define the share that debtor i borrows in currency \$ from all debtors as $s^{i}(\$)$, we have

$$s^i(\$) \equiv \mathbb{P}\Big(R^{i\$}\mathcal{E}^{i\$} \le \min_{k\neq\$}\Big\{R^{i\$}\mathcal{E}^{i\$}\Big\}\Big) = \int_0^\infty \prod_{k\neq\$}\left(1-\exp\left(-\overline{e}^i\left(\frac{R^{i\$}}{R^{ik}}\right)^{-\theta}\epsilon^{-\theta}\right)\right)dG_{\mathcal{E}^{i\$}}(\epsilon)$$

While a closed-form solution does not exist in general and must be computed numerically, it does exist when there are only two currencies:

$$\begin{aligned}
s^i(\$) &\equiv \int_0^1 1^{ij}(\$)dj = \mathbb{E}[1(\$)] \\
&= \mathbb{P}\Big(R^{i\$}\mathcal{E}^{i\$} \le R^{i£}\mathcal{E}^{i£}\Big) \\
&= \int_0^\infty \mathbb{P}\Big(R^{i\$}\epsilon \le R^{i£}\mathcal{E}^{i£}\Big)dG_{\mathcal{E}^{i\$}}(\epsilon) \\
&= \int_0^\infty \left(1-\mathbb{P}\left(\mathcal{E}^{i£} \le \frac{R^{i\$}}{R^{i£}}\epsilon\right)\right)dG_{\mathcal{E}^{i\$}}(\epsilon) \\
&= \int_0^\infty \left(1-\exp\left(-\overline{e}^i\left(\frac{R^{i\$}}{R^{i£}}\epsilon\right)^{-\theta}\right)\right)(-\overline{e}^i)(-\theta\epsilon^{-\theta-1})\Big(\exp(-\overline{e}^i\cdot\epsilon^{-\theta})\Big)dG_{\mathcal{E}^{i\$}}(\epsilon) \\
&= 1-(-\overline{e}^i)\int_0^\infty \exp\left(-\underbrace{\left(\overline{e}^i\left(\frac{R^{i\$}}{R^{i£}}\right)^{-\theta}+\overline{e}^i\right)}_{\Phi}\epsilon^{-\theta}\right)d(\epsilon^{-\theta}) \\
&[u=\epsilon^{-\theta}, du=-\theta\epsilon^{-\theta-1}; \epsilon=0, u=\infty; \epsilon=\infty, u=0] \\
&= 1-(-\overline{e}^i)(-1)\frac{1}{-\Phi}\underbrace{\left(\frac{1}{\exp(-\Phi\cdot\infty)}-\frac{1}{\exp(-\Phi\cdot 0)}\right)}_{-1} \\
&= 1-\frac{\overline{e}^i}{\underbrace{\left(\overline{e}^i\left(\frac{R^{i\$}}{R^{i£}}\right)^{-\theta}+\overline{e}^i\right)}_{\Phi}} = 1-\left(\frac{\overline{e}^i(R^{i£})^\theta}{\overline{e}^i(R^{i\$})^\theta+\overline{e}^i(R^{i£})^\theta}\right) \\
&= \frac{(R^{i£})^\theta}{(R^{i\$})^\theta+(R^{i£})^\theta}
\end{aligned}$$

Comparative static analysis indicates that $\dfrac{\partial s^i(\$)}{\partial R^{i\$}_{t+1}} < 0, \dfrac{\partial s^i(\$)}{\partial R^{i£}_{t+1}} > 0$, and in particular,

$$
\begin{aligned}
\frac{\partial s^i(\$)}{\partial \theta^i} &= \frac{\partial}{\partial \theta^i}\left(\frac{\exp(\theta \ln R^{i\pounds})}{\exp(\theta \ln R^{i\$}) + \exp(\theta \ln R^{i\pounds})}\right) \\
&= \frac{\Big((R^{i\$})^{\theta} + (R^{i\pounds})^{\theta}\Big) \cdot \ln R^{i\pounds} \cdot (R^{i\pounds})^{\theta} - (R^{i\pounds})^{\theta} \cdot \Big(\ln R^{i\$} \cdot (R^{i\$})^{\theta} + \ln R^{i\pounds} \cdot (R^{i\pounds})^{\theta}\Big)}{+} \\
&= -\frac{\Big(R^{i\$} R^{i\pounds}\Big)\Big(\ln R^{i\$} - \ln R^{i\pounds}\Big)}{+} \\
&\propto -(r^{i\$} - r^{i\pounds})
\end{aligned}
$$

In other words, the effect of tail thickness on the currency share depends on the relative borrowing costs: when the U.S. lending rate is cheaper than British pounds, the U.S. dollar share increases with θ^i (i.e., thinner tail). Therefore, less extreme depreciation makes currency shares more responsive to systematic cost differences.

A3. Relationship Between Default Probability and Shift-share Instrument

To linearize Equation (2.9),

$$\pi_{t+1|t}^{i} = \frac{\sum_k \overline{e}^k \left(R_t^{ik}\right)^{\theta^i}}{\overline{t}^{\,i} + \sum_k \overline{e}^k \left(R_t^{ik}\right)^{\theta^i}}$$

we first define $A_t^i \equiv \sum_k \overline{e}^k \left(R_t^{ik}\right)^{\theta^i}$, as well as steady state values $\overline{R}^{ik}$, $\overline{A}^i \equiv \sum_k \overline{e}^k \left(\overline{R}^{ik}\right)^{\theta^i}$, $\overline{\pi}^i = \dfrac{\overline{A}^i}{\overline{t}^{\,i} + \overline{A}^i}$. Define $\hat{\pi}_{t+1|t}^i \equiv \dfrac{d\pi_{t+1|t}^i}{\overline{\pi}^i}$ and $\hat{R}_t^{ik} \equiv \dfrac{dR_t^{ik}}{\overline{R}^{ik}}$ as percentage deviations from steady state values. Taking total differential over the rewritten equation $\pi_{t+1|t}^i = \dfrac{A_t^i}{\overline{t}^{\,i} + A_t^i}$ yields:

$$d\pi_{t+1|t}^{i} = \frac{\left(\overline{t}^{\,i} + \overline{A}^i\right)(1) - \overline{A}^i(1)}{\left(\overline{t}^{\,i} + \overline{A}^i\right)^2} dA_t^i = \frac{\overline{t}^{\,i}}{\left(\overline{t}^{\,i} + \overline{A}^i\right)^2} dA_t^i$$

$$dA_t^i = \theta^i \sum_k \overline{e}^k \left(\overline{R}^{ik}\right)^{\theta^i - 1} dR_t^{ik} \frac{\overline{R}^{ik}}{\overline{R}^{ik}} = \theta^i \sum_k \overline{e}^k \left(\overline{R}^{ik}\right)^{\theta^i} \frac{dR_t^{ik}}{\overline{R}^{ik}} = \theta^i \sum_k \overline{e}^k \left(\overline{R}^{ik}\right)^{\theta^i} \hat{R}_t^{ik}$$

Combining all equations above, we have:

$$\begin{aligned}
\hat{\pi}_{t+1|t}^{i} &= \frac{1}{\overline{\pi}^i} d\pi_{t+1|t}^i \\
&= \left(\frac{\overline{t}^{\,i} + \overline{A}^i}{\overline{A}^i}\right)\left(\frac{\overline{t}^{\,i}}{\left(\overline{t}^{\,i} + \overline{A}^i\right)^2}\right)\left(\theta^i \sum_k \overline{e}^k \left(\overline{R}^{ik}\right)^{\theta^i} \hat{R}_t^{ik}\right) \\
&= \left(\frac{\overline{t}^{\,i}}{\overline{t}^{\,i} + \overline{A}^i}\right)\left(\frac{\theta^i \sum_k \overline{e}^k \left(\overline{R}^{ik}\right)^{\theta^i} \hat{R}_t^{ik}}{\overline{A}^i}\right) \\
&= \theta^i \left(1 - \overline{\pi}^i\right) \frac{\sum_k \overline{e}^k \left(\overline{R}^{ik}\right)^{\theta^i} \hat{R}_t^{ik}}{\sum_k \overline{e}^k \left(\overline{R}^{ik}\right)^{\theta^i}}
\end{aligned}$$

where the last step follows from $\overline{\pi}^i = \dfrac{\overline{A}^i}{\overline{t}^i + \overline{A}^i} \Rightarrow 1 - \overline{\pi}^i = \dfrac{\overline{t}^i}{\overline{t}^i + \overline{A}^i}$. Assume a single contractionary monetary shock occurs in currency \$, that is, $\hat{R}_t^{i\$} > 0$ and $\forall k \neq \$, \hat{R}_t^{ik} = 0$, then we have:

$$\hat{\pi}_{t+1|t}^i = \theta^i \left(1 - \overline{\pi}^i\right)\left(\frac{\overline{e}^{\$}\left(\overline{R}^{i\$}\right)^{\theta^i}}{\sum_k \overline{e}^k \left(\overline{R}^{ik}\right)^{\theta^i}}\right)\hat{R}_t^{i\$} \qquad \blacksquare$$

To interpret the results as percentage points changes from steady state values, define $\tilde{\pi}_{t+1|t}^i \equiv \pi_{t+1|t}^i - \overline{\pi}^i = d\pi_{t+1|t}^i$ and $\tilde{R}_t^{ik} \equiv R_t^{ik} - \overline{R}^{ik} = dR_t^{ik}$ (both in percentage points deviations), then we have:

$$\begin{aligned}
\tilde{\pi}_{t+1|t}^i &= d\pi_{t+1|t}^i \\
&= \frac{\left(\overline{t}^i + \overline{A}^i\right)(1) - \overline{A}^i(1)}{\left(\overline{t}^i + \overline{A}^i\right)^2} dA_t^i \\
&= \frac{\overline{t}^i}{\left(\overline{t}^i + \overline{A}^i\right)^2}\left(\theta^i \sum_k \overline{e}^k \left(\overline{R}^{ik}\right)^{\theta^i - 1} dR_t^{ik}\left(\frac{\overline{R}^{ik}}{\overline{R}^{ik}}\right)\right)\left(\frac{\overline{A}^i}{\overline{A}^i}\right) \\
&= \theta^i \left(\frac{\overline{t}^i}{\overline{t}^i + \overline{A}^i}\right)\left(\frac{\overline{A}^i}{\overline{t}^i + \overline{A}^i}\right)\left(\frac{\sum_k \overline{e}^k \left(\overline{R}^{ik}\right)^{\theta^i}}{\overline{A}^i}\right) \\
&= \theta^i \overline{\pi}^i \left(1 - \overline{\pi}^i\right)\left(\frac{1}{\sum_k \overline{e}^k \left(\overline{R}^{ik}\right)^{\theta^i}}\right)\left(\sum_k \overline{e}^k \left(\overline{R}^{ik}\right)^{\theta^i} \frac{\tilde{R}_t^{ik}}{\overline{R}^{ik}}\right)
\end{aligned}$$

Again, a single contractionary monetary shock occurs in currency \$, that is, $\tilde{R}_t^{i\$} \equiv R_t^{i\$} - \overline{R}^{i\$} > 0$ and $\forall k \neq \$, \tilde{R}_t^{ik} = 0$, then we have:

$$\tilde{\pi}_{t+1|t}^i = \frac{\theta^i \overline{\pi}^i \left(1 - \overline{\pi}^i\right)}{\overline{R}^{i\$}}\left(\frac{\overline{e}^{\$}\left(\overline{R}^{i\$}\right)^{\theta^i}}{\sum_k \overline{e}^k \left(\overline{R}^{ik}\right)^{\theta^i}}\right)\tilde{R}_t^{i\$} \qquad \blacksquare$$

A4. Control Function Approach to Address Exclusion Restriction

Suppose the true model is given as

$$y = \beta D + \phi z + v \tag{A4.1}$$

where y is the dependent variable, D is a binary treatment indicator, and z is the instrument for D. Assume $\left(y, D, z\right)$ are covariance-stationary so that standard OLS asymptotic inference applies.[20] If the treatment variable D is not truly exogenous, then $\mathbb{E}\left[Dv\right] \neq 0$. Assume z is an exogenous instrument for D but it violates the exclusion restriction—$\mathbb{E}\left[zv\right] = 0$ but $\phi \neq 0$. Estimating a two-stage least squares (2SLS) regression:

$$D = bz + \eta; \qquad y = \beta D + \phi z + v$$

will yield an inconsistent estimator because

$$\begin{aligned}
\hat{\beta}^{IV} &= \frac{\operatorname{cov}(y, \hat{D})}{V(\hat{D})} \\
&= \frac{\operatorname{cov}(\beta D + \phi z + v, bz)}{V(bz)} \\
&= \frac{\beta b \operatorname{cov}(D, z)}{b^2 V(z)} + \frac{\phi b\, V(z)}{b^2 V(z)} + \frac{b \operatorname{cov}(v, z)}{b^2 V(z)} \\
&\xrightarrow{p} \beta + \frac{\phi}{b}
\end{aligned}$$

The key idea of the control function approach is to, first, estimate the instrument's direct (spillover) effect on the dependent variable in a subsample in which treatment never occurs. Then, I subtract that spillover component from the dependent variable in the full sample and estimate the 2SLS regression again. This allows us to assess how the IV estimate

[20] The long-difference $y_{it+h} - y_{it-1}$—the dependent variable—is stationary. The first-year default indicator is stationary in the panel, and the shift-share instrument ΔIRE is stationary because the augmented Dickey-Fuller test—run for each country over all years—shown in Table A12 strongly rejects a unit root.

of β changes when we vary the assumed strength of the spillover—that is, to compute upper and lower bounds for true treatment estimator β. To be specific, consider the subsample of countries that have never defaulted (i.e., $D=0$ in all observed years). In this subsample, the structural equation simplifies to $y^{ND}=\phi^{ND}z^{ND}+v^{ND}$. Assume the instrument remains exogenous in this subsample, that is, $\mathbb{E}\left[z^{ND}v^{ND}\right]=0$. Then the OLS estimator in the never-default subsample satisfies:

$$\hat{\phi}^{ND}=\frac{\mathrm{cov}(\phi^{ND}z^{ND},z^{ND})}{V(z^{ND})}\xrightarrow{p}\phi^{ND}$$

Next, assume that the spillover effect of the instrument z in default episodes is proportional to the one identified in never-default episodes, that is, $\phi=\lambda\phi^{ND}$ with $\lambda>0$. Since $\hat{\phi}^{ND}\xrightarrow{p}\phi^{ND}$, we have $\lambda\hat{\phi}^{ND}\xrightarrow{p}\lambda\phi^{ND}=\phi$. Subtracting $\lambda\hat{\phi}^{ND}z$ from the dependent variable y in the full sample, we have:

$$y-\lambda\hat{\phi}^{ND}z=\beta D+(\phi-\lambda\hat{\phi}^{ND})z+v$$

Multiply both sides by z and taking expectations yields the orthogonality condition for spillover-corrected IV estimator $\hat{\beta}_{CF}^{IV}(\lambda)$:

$$\mathbb{E}\left[\left(y-\lambda\hat{\phi}^{ND}z\right)z\right]=\beta\mathbb{E}\left[Dz\right]+\left(\phi-\lambda\hat{\phi}^{ND}\right)\mathbb{E}\left[z^2\right]+\mathbb{E}\left[zv\right]$$

since the term $\left(\phi-\lambda\hat{\phi}^{ND}\right)\rightarrow 0$ as sample size grows. Therefore, a 2SLS regression that uses the spillover-corrected dependent variable together with the original instrument yields a consistent estimator of the true treatment effect β:

$$\hat{\beta}_{CF}^{IV}(\lambda)=\frac{\mathrm{cov}(y-\lambda\hat{\phi}^{ND}z,z)}{\mathrm{cov}(D,z)}=\frac{\frac{1}{N}\sum_{i=1}^{N}\left(\left(y_i-\lambda\hat{\phi}^{ND}z_i\right)z_i\right)}{\frac{1}{N}\sum_{i=1}^{N}D_iz_i}\xrightarrow{p}\frac{\mathbb{E}\left[\left(y-\lambda\hat{\phi}^{ND}z\right)z\right]}{\mathbb{E}\left[Dz\right]}=\beta \quad \text{(A4.3)}$$

as long as $\frac{1}{N}\sum_{i=1}^{N} z_i^2 \xrightarrow{p} \mathbb{E}\left[z^2\right] < \infty$ when $N \to \infty$ ■

Table A1: Data sample and availability

Country	Period	IDS data availability	Country	Period	IDS data availability
AGO	1989-2010	online, CD2010, CD2006	LAO	1984-2010	online, CD2010, CD2006
ALB	1991-2010	online, CD2010, CD2006	LBN	1988-2010	online, CD2010, CD2006
ARG	1970-2010	online, CD2010, CD2006	LBR	2000-2010	online, CD2010, CD2006
ARM	1993-2010	online, CD2010, CD2006	LKA	1970-2010	online, CD2010, CD2006
AZE	1993-2010	online, CD2010, CD2006	LSO	1970-2010	online, CD2010, CD2006
BDI	1970-2010	online, CD2010, CD2006	LTU	1995-2010	CD2010, CD2006
BEN	1970-2010	online, CD2010, CD2006	LVA	1995-2010	CD2010, CD2006
BFA	1970-2010	online, CD2010, CD2006	MAR	1970-2010	online, CD2010, CD2006
BGD	1972-2010	online, CD2010, CD2006	MDA	1995-2010	online, CD2010, CD2006
BGR	1981-2010	online, CD2010, CD2006	MDG	1970-2010	online, CD2010, CD2006
BIH	1999-2010	online, CD2010, CD2006	MDV	1995-2010	online, CD2010, CD2006
BLR	1993-2010	online, CD2010, CD2006	MEX	1970-2010	online, CD2010, CD2006
BLZ	1980-2010	online, CD2010, CD2006	MKD	1993-2010	online, CD2010, CD2006
BOL	1970-2010	online, CD2010, CD2006	MLI	1970-2010	online, CD2010, CD2006
BRA	1970-2010	online, CD2010, CD2006	MMR	1970-2010	online, CD2010, CD2006
BRB	1974-2005	CD2006	MNG	1992-2010	online, CD2010, CD2006
BTN	1982-2010	online, CD2010, CD2006	MOZ	1991-2010	online, CD2010, CD2006
BWA	1970-2010	online, CD2010, CD2006	MRT	1970-2010	online, CD2010, CD2006
CAF	1970-2010	online, CD2010, CD2006	MUS	1976-2010	online, CD2010, CD2006
CHL	1970-2010	CD2010, CD2006	MWI	1980-2010	online, CD2010, CD2006
CHN	1981-2010	online, CD2010, CD2006	MYS	1970-2010	CD2010, CD2006
CIV	1970-2010	online, CD2010, CD2006	NER	1970-2010	online, CD2010, CD2006
CMR	1970-2010	online, CD2010, CD2006	NGA	1970-2010	online, CD2010, CD2006
COD	1970-2010	online, CD2010, CD2006	NIC	1970-2010	online, CD2010, CD2006
COG	1970-2010	online, CD2010, CD2006	NPL	1971-2010	online, CD2010, CD2006
COL	1970-2010	online, CD2010, CD2006	PAK	1970-2010	online, CD2010, CD2006
COM	1980-2010	online, CD2010, CD2006	PAN	1970-2010	online, CD2010, CD2006
CPV	1981-2010	online, CD2010, CD2006	PER	1970-2010	online, CD2010, CD2006
CRI	1970-2010	online, CD2010, CD2006	PHL	1970-2010	online, CD2010, CD2006
CZE	1993-2004	CD2006	PNG	1970-2010	online, CD2010, CD2006
DMA	1981-2010	online, CD2010, CD2006	POL	1990-2008	CD2010, CD2006
DOM	1970-2010	online, CD2010, CD2006	PRY	1970-2010	online, CD2010, CD2006
DZA	1970-2010	online, CD2010, CD2006	ROU	1990-2010	online, CD2010, CD2006
ECU	1970-2010	online, CD2010, CD2006	RUS	1992-2010	online, CD2010, CD2006
EGY	1970-2010	online, CD2010, CD2006	RWA	1971-2010	online, CD2010, CD2006
ERI	1994-2005	online, CD2010, CD2006	SDN	1970-2010	online, CD2010, CD2006
EST	1995-2005	CD2006	SEN	1970-2010	online, CD2010, CD2006
ETH	1981-2010	online, CD2010, CD2006	SLE	1970-2010	online, CD2010, CD2006
FJI	1970-2010	online, CD2010, CD2006	SLV	1970-2010	online, CD2010, CD2006
GAB	1970-2010	online, CD2010, CD2006	STP	2001-2010	online, CD2010, CD2006
GEO	1992-2010	online, CD2010, CD2006	SVK	1993-2006	CD2006

GHA	1970-2010	online, CD2010, CD2006	SWZ	1970-2010	online, CD2010, CD2006
GIN	1986-2010	online, CD2010, CD2006	SYC	1980-2010	CD2010, CD2006
GMB	1970-2010	online, CD2010, CD2006	SYR	1970-2004	online, CD2006
GNB	1974-2010	online, CD2010, CD2006	TCD	1970-2010	online, CD2010, CD2006
GNQ	1984-2006	CD2006	TGO	1970-2010	online, CD2010, CD2006
GRD	1977-2010	online, CD2010, CD2006	THA	1970-2010	online, CD2010, CD2006
GTM	1970-2010	online, CD2010, CD2006	TJK	1992-2010	online, CD2010, CD2006
GUY	1970-2010	online, CD2010, CD2006	TKM	1993-2010	online, CD2010
HND	1970-2010	online, CD2010, CD2006	TTO	1970-2005	CD2006
HRV	1995-2006	CD2006	TUN	1970-2010	online, CD2010, CD2006
HTI	1970-2010	online, CD2010, CD2006	TUR	1970-2010	online, CD2010, CD2006
HUN	1991-2006	CD2006	TZA	1988-2010	online, CD2010, CD2006
IDN	1970-2010	online, CD2010, CD2006	UGA	1982-2010	online, CD2010, CD2006
IND	1970-2010	online, CD2010, CD2006	UKR	1992-2010	online, CD2010, CD2006
IRN	1979-2010	online, CD2010, CD2006	URY	1970-2010	CD2010, CD2006
JAM	1970-2010	online, CD2010, CD2006	UZB	1992-2010	online, CD2010, CD2006
JOR	1976-2010	online, CD2010, CD2006	VEN	1970-2008	online, CD2010, CD2006
KAZ	1992-2010	online, CD2010, CD2006	VNM	1988-2010	online, CD2010, CD2006
KEN	1970-2010	online, CD2010, CD2006	YEM	1990-2010	online, CD2010, CD2006
KGZ	1992-2010	online, CD2010, CD2006	ZAF	1994-2010	online, CD2010, CD2006
KHM	1993-2010	online, CD2010, CD2006	ZMB	1970-2010	online, CD2010, CD2006
KNA	1984-2010	CD2010, CD2006	ZWE	1970-2008	online, CD2010, CD2006

Notes: This table provides details on the country sample and the availability of *International Debt Statistics* (IDS) used in the analysis. The term "Online" indicates that IDS data were obtained from the World Bank online data portal. The term "CD2010" indicates that the data were retrieved from the 2010 *Global Development Finance* CD-ROM (the predecessor to IDS), likewise for "CD2006". These CD-ROMs can be accessed through interlibrary loan service.

Table A2: Descriptive statistics for variables used in the regression analysis

	(1) Mean	(2) SD	(3) Min	(4) Median	(5) Max	(6) #obs
Panel (a) Treatment-related variables						
Binary default	0.026	0.159	0	0	1	3898
$\Delta\ln(\text{arrear})$	0.050	1.464	-18.604	0.079	11.649	2718
$\Delta\ln(1+\text{arrear})$	0.023	0.937	-9.063	0.000	7.921	3898
$\Delta\ln(1+\frac{\text{arrears}}{\text{debt}})$	-0.000	0.432	-4.078	0.000	3.796	3898
ΔIRE	-0.105	0.637	-2.668	-0.128	4.162	3898
IRE^{USRR}	0.005	0.437	-3.523	0.075	1.763	3402
IRE^{UKCH}	0.004	0.095	-1.502	0.000	1.195	3402
Panel (b) Continuous variables						
$\Delta\ln(\text{rgdppc})$	0.019	0.057	-0.538	0.023	0.877	3898
$\Delta\ln(\text{GDPdefl})$	0.157	0.344	-0.379	0.082	5.593	3898
$\Delta\ln(\text{xr})$	0.122	0.485	-9.375	0.028	13.450	3898
Debt-to-GDP ratio	0.551	0.571	0.018	0.411	10.874	3898
Chinn-Ito Index	0.346	0.304	0	0.163	1	3898
Average maturities	22.658	10.285	1.333	20.765	50.111	3712
Floating rate share	0.208	0.207	0.000	0.150	0.940	3848
Panel (c) Binary variables						
Banking crisis	0.025	0.157	0	0	1	3898
Currency crisis	0.041	0.198	0	0	1	3898
Democracy	0.472	0.499	0	0	1	3898
Peg	0.698	0.459	0	1	1	3834

Notes: This table reports descriptive statistics for all variables used in the empirical analysis.

Table A3: Balance test between defaulters and never-defaulters

	(1) Defaulters	(2) Never-defaulters	(3) Difference	(4) p-value
Panel (a) Treatment-related variables				
Binary default	0.047	0.000	0.047***	0.000
$\Delta\ln(\text{arrear})$	0.124	-0.071	0.195***	0.001
$\Delta\ln(1+\text{arrear})$	0.059	-0.020	0.080***	0.007
$\Delta\ln(1+\frac{\text{arrears}}{\text{debt}})$	0.018	-0.022	0.041***	0.003
ΔIRE	-0.091	-0.123	0.032	0.111
IRE^{USRR}	0.002	0.009	-0.007	0.633
IRE^{UKCH}	0.004	0.004	0.000	0.942
Panel (b) Continuous variables				
$\Delta\ln(\text{rgdppc})$	0.010	0.031	-0.021***	0.000
$\Delta\ln(\text{GDPdefl})$	0.193	0.114	0.079***	0.000
$\Delta\ln(\text{xr})$	0.161	0.075	0.086***	0.000
Debt-to-GDP ratio	0.627	0.460	0.166***	0.000
Chinn-Ito Index	0.347	0.345	0.002	0.841
Panel (c) Binary variables				
Banking crisis	0.031	0.019	0.012**	0.013
Currency crisis	0.056	0.023	0.034***	0.000
Democracy	0.468	0.476	-0.008	0.613
Peg	0.629	0.782	-0.154***	0.000

Notes: This table reports a balance test for macroeconomic variables and shows that there are systematic differences between countries with default experience and those without: defaulters have, on average, lower output growth, higher inflation, higher currency depreciation, and higher debt-to-GDP ratios; they are also more likely to experience banking and currency crises and less likely to main a fixed exchange rate regime, which is consistent with limited credibility under hard pegs. A second observation is that "interest rate exposure" measures do not differ significantly between defaulters and never-defaulters; nor do the Chinn-Ito (2006) index (capital account openness) or the Acemoglu et al. (2019) democracy indicator. "Defaulters" are defined as countries that have defaulted on their external debt at least once over the 1970–2010 period as specified by Kuvshinov and Zimmermann (2019), while "Never-defaulters" are countries that have no defaults in the same period.

Table A4: Argentina's bilateral "other-currency" share (selected years and creditors)

(1) Counterpart	(2) Indicator	(3) 1970	(4) 1980	(5) 1990	(6) 2000	(7) 2010	(8) 2020
Panel (a) "Other-currency" to all creditors (world aggregate)							
WLD	d_{it}	5,893.2	27,322.5	62,477.6	150,062.9	126,642.4	275,474.9
	$s_{it}^{\text{OTHC,all}}$	4.7%	3.5%	1.0%	6.3%	1.6%	2.8%
Panel (b) "Other-currency" to selected advanced economies (bilateral)							
Bondholders	d_{ijt}	386.1	832.1	11,543.0	80,158.0	56,025.1	103,598.7
	s_{ijt}^{OTHC}	—§	—§	—§	7.8%	0.5%	2.9%
	$s_{ijt}^{\text{OTHC,all}}$				4.1%	0.2%	1.1%
AUS	d_{ijt}	0	—†	0	—†	0.6	0.1
	s_{ijt}^{OTHC}	—¶	—¶	—¶	—¶	100%	—§
	$s_{ijt}^{\text{OTHC,all}}$					0.0%	0.0%
CAN	d_{ijt}	10.8	304.9	50.9	120.3	147.5	45.1
	s_{ijt}^{OTHC}	51.7%	59.6%	49.0%	38.1%	73.0%	41.1%
	$s_{ijt}^{\text{OTHC,all}}$	0.1%	0.7%	0.0%	0.0%	0.1%	0.0%
DNK	d_{ijt}	—†	—†	20.6	9.5	10.7	3.3
	s_{ijt}^{OTHC}	—¶	—¶	59.3%	24.2%	0.0%	27.6%
	$s_{ijt}^{\text{OTHC,all}}$			0.1%	0.0%	0.0%	0.0%
SWE	d_{ijt}	3.3	20.9	38.3	169.0	314.4	4.0
	s_{ijt}^{OTHC}	92.0%	68.1%	39.5%	2.1%	4.5%	32.8%
	$s_{ijt}^{\text{OTHC,all}}$	0.1%	0.1%	0.0%	0.0%	0.0%	0.0%
Panel (c) "Other-currency" to selected emerging markets (bilateral)							
BRA	d_{ijt}	0.4	17.4	151.6	46.2	457.4	239.3
	s_{ijt}^{OTHC}	0	0	—§	0	0	13.1%
	$s_{ijt}^{\text{OTHC,all}}$						0.0%
CHN	d_{ijt}	—†	—†	—†	—†	25	22,928
	s_{ijt}^{OTHC}	—¶	—¶	—¶	—¶	0	5.6%
	$s_{ijt}^{\text{OTHC,all}}$					0.0%	0.5%
IADB	d_{ijt}	183.0	848.8	2,623.3	7,709.9	10,311.6	13,368.2

s_{ijt}^{OTHC}	45.6%	19.5%	10.6%	3.4%	0.8%	0.0%
$s_{ijt}^{\text{OTHC,all}}$	1.4%	0.6%	0.4%	0.2%	0.1%	0.0%

Notes: This table reports the *Argentina*'s "other-currency" denomination vis-à-vis its creditors for selected years (the dataset is available at annual level). The empirical analysis in this paper uses the *world aggregate* and therefore does *not* take a stance on whether "other-currency" should be classified as local- or foreign-currency borrowing, although the patterns are consistent with the "original sin" literature—most developing countries' debt are denominated in foreign currency. d_{it} presents the total external debt stock of country i in year t (in million current US dollars). $s_{it}^{\text{OTHC,all}}$ is the "other-currency" share of country i in year t vis-a-via all creditors in the world. s_{ijt}^{OTHC} is the bilateral "other-currency" share of country i with respect to its creditors j in year t, and $s_{ijt}^{\text{OTHC,all}}$ is the "bilateral" OTHC share between i and j against the total external debt of all creditors, that is $s_{ijt}^{\text{OTHC,all}} = \frac{s_{ijt}^{\text{OTHC}} \times d_{ijt}}{d_{it}}$. Inferred IDS's coding convention: —†: data entries are missing (either blank or "."); —¶ IDS codes OTHC ("other-currency share") as 100, but often then the bilateral external debts are either 0 or missing; —§: IDS codes OTHC as 0.001, 0.002, 0.003, or 0.004, which appears to reflect encoding practice and carry special meanings rather than measurement. **0** on the above table are actual recorded values. Since this paper relies on world-aggregate shares and levels, a detailed treatment of bilateral debt denomination, in particular local- and/or foreign-currency denomination, is left to a separate analysis.

Table A5: Comparison of cumulative output loss across financial crises

Variable of interest: cumulative output loss $y_{t+h} - y_{t-1}$ (h in years)	$h=0$	$h=1$	$h=2$	$h=3$	$h=4$	$h=5$	$h=6$
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Panel (a): Baseline LP-SSIV causal effect of sovereign default on output loss							
Baseline LP-SSIV	-8.0%	-18.5%	-15.8%	-18.3%	-9.5%	-5.7%	-0.5%
Panel (b): Representative sovereign default episodes							
ARG2001	-5.6%	-18.2%	-10.8%	-3.2%	4.3%	11.0%	18.6%
RUS1998	-5.3%	1.2%	11.2%	16.6%	21.6%	29.1%	36.5%
ECU1999	-6.6%	-7.2%	-5.0%	-2.7%	-1.8%	4.4%	7.8%
GRC2010	-5.8%	-16.3%	-23.1%	-24.9%	-23.8%	-23.3%	-23.4%
Panel (c): Currency/financial crises							
MEX1995	-7.9%	-3.7%	1.5%	5.8%	6.9%	10.2%	8.2%
THA1998	-9.2%	-5.8%	-2.5%	0.0%	5.1%	11.2%	16.4%
Panel (d): Benchmark—2008 Global Financial Crisis							
USA2008	-0.8%	-4.3%	-2.5%	-1.7%	-0.2%	1.0%	2.5%

Notes: This table compares the magnitude and the persistence of cumulative output loss from sovereign default with other financial crises. The values are cumulative output loss measured by the log-difference of $y_{t+h} - y_{t-1}$, where y_{t+h} is the log value of the h-years ahead real GDP per capita in constant price and local currency series retrieved from the World Development Indicator. Panel (a) reports baseline LP-SSIV estimates from Table 3. ARG2001 refers to Argentina 2001–2002 default; RUS1998 refers to Russia's 1998 default; ECU1999 refers to Ecuador 1999 default; and GRC2010 refers to Greece's 2010 debt restructuring (not a legal default). For output loss during financial crises (Panels (b)–(d)), I use the first contraction year, where MEX1995 refers to the Mexico 1994–1995 Tequila Crisis; THA1998 refers to the 1997–1998 Asian Financial Crisis; and last but not least, USA2008 refers to the 2008 Global Financial Crisis, which serves as a benchmark for comparison.

Table A6: Comparison between linear and logistic regression on the default probability

Dependent variable: S&P first-year binary (0/1) default indicator			
	Linear	Logistic	Linear (same sample as (2))
	(1)	(2)	(3)
ΔIRE_{it-1}	0.009	0.308	0.014
	(0.006)	(0.191)	(0.010)
ΔIRE_{it-2}	0.026***	0.815***	0.040***
	(0.006)	(0.213)	(0.009)
ΔIRE_{it-3}	0.010*	0.166	0.014
	(0.006)	(0.218)	(0.010)
Observations	3,898	2,030	2,030
R-squared	0.089		0.093
Country FE	Y	Y	Y
Baseline controls	Y	Y	Y

Notes: This table reports the first-stage regressions estimated by both a linear OLS regression and a non-linear logit regression as a robustness check. Since only linear OLS regressions yield first-stage residuals orthogonal to fitted values and regressors, two-stage least squares (2SLS) does not apply to a logit second stage (the "forbidden regression"). Column (1) reproduces the OLS first-stage results from Table 2, and column (2) reports logit estimates with the same set of baseline controls and country fixed effects as in Table 2. Since logit regression drops observations from groups within no within-group variation in the first-year default indicator (e.g., always-defaulters or never-defaulters), column (3) reports the OLS results using the same sample as in column (2). Robust standard errors, clustered at the country level to account for serial correlations, are reported in parentheses. Statistical significance is indicated as *** p<0.01, ** p<0.05, * p<0.1. See Footnote 14 in the main text for interpretation.

Table A7: Amplification of $\Delta IRE_{it-\ell}$ through floating rate liabilities

Dependent variable: S&P first-year binary default indicator				
	(1)	(2)	(3)	(4)
$\text{sharevarrate}_{it-1}$	0.0341	0.0524**	0.0440*	0.0587**
	(0.0223)	(0.0245)	(0.0244)	(0.0259)
ΔIRE_{it-1}	0.0143*			0.0143
	(0.0084)			(0.0090)
$\Delta IRE_{it-1} \times \text{sharevarrate}_{it-1}$	-0.0290			-0.0202
	(0.0278)			(0.0277)
ΔIRE_{it-2}		0.0163**		0.0155**
		(0.0068)		(0.0066)
$\Delta IRE_{it-2} \times \text{sharevarrate}_{it-1}$		0.0470**		0.0471**
		(0.0193)		(0.0183)
ΔIRE_{it-3}			-0.0002	0.0003
			(0.0061)	(0.0067)
$\Delta IRE_{it-3} \times \text{sharevarrate}_{it-1}$			0.0489**	0.0453*
			(0.0241)	(0.0232)
Observations	3,872	3,872	3,872	3,872
R-squared	0.0788	0.0891	0.0810	0.0936
Country FE	Y	Y	Y	Y
Baseline controls	Y	Y	Y	Y

Notes: This table reports coefficient estimates on interaction terms between $\Delta IRE_{it-\ell}$ and $\text{sharevarrate}_{it-1}$ to investigate the mechanisms behind the delayed response of default probability to foreign interest rate hikes. The variable $\text{sharevarrate}_{it-1}$ is defined as the share of external debt stock denominated at variable rate—LIBOR or the U.S. prime linked interest rates— out of total external debt stock. The main finding is that, conditional on $\Delta IRE_{it-\ell}$, the default probability increases with the share of external debt stock under variable rate contracts. The data come from the World Bank's International Debt statistics. All columns include country fixed effects and the set of baseline controls specified in Table 2. Robust standard errors, clustered at the country level to account for serial correlations, are reported in parentheses. Statistical significance is indicated as *** p<0.01, ** p<0.05, * p<0.1.

Table A8: Longer debt maturities mitigate first-stage effect on $\Delta IRE_{it-\ell}$

Dependent variable: S&P first-year binary default indicator				
	(1)	(2)	(3)	(4)
Maturity_{it-1}	-0.0010**	-0.0011***	-0.0011***	-0.0011***
	(0.0004)	(0.0004)	(0.0004)	(0.0004)
ΔIRE_{it-1}	-0.0050			-0.0055
	(0.0147)			(0.0150)
$\Delta IRE_{it-1} \times \text{Maturity}_{it-1}$	0.0005			0.0005
	(0.0005)			(0.0006)
ΔIRE_{it-2}		0.0551***		0.0537***
		(0.0133)		(0.0132)
$\Delta IRE_{it-2} \times \text{Maturity}_{it-1}$		-0.0013***		-0.0013**
		(0.0005)		(0.0005)
ΔIRE_{it-3}			0.0308**	0.0239
			(0.0152)	(0.0148)
$\Delta IRE_{it-3} \times \text{Maturity}_{it-1}$			-0.0010*	-0.0007
			(0.0006)	(0.0006)
Observations	3,571	3,571	3,571	3,571
R-squared	0.0852	0.0969	0.0875	0.0997
Country FE	Y	Y	Y	Y
Baseline controls	Y	Y	Y	Y

Notes: This table reports coefficient estimates on interaction terms between $\Delta IRE_{it-\ell}$ and Maturity_{it-1} to investigate the mechanisms behind the delayed response of default probability to foreign interest rate hikes. The variable Maturity_{it-1} is defined as the average maturity of new commitments (in years). The main finding is that, conditional on $\Delta IRE_{it-\ell}$, longer average maturities lower the default probability. The data come from the World Bank's International Debt statistics. All columns include country fixed effects and the set of baseline controls specified in Table 2. Robust standard errors, clustered at the country level to account for serial correlations, are reported in parentheses. Statistical significance is indicated as *** p<0.01, ** p<0.05, * p<0.1.

Table A9: Robustness check using lagged shares in the shift-share instrument

Panel (a): First-stage results with lagged currency shares ($\Delta IREls_{it}$)

Dependent variable: S&P first-year binary default indicator

	(1)	(2)	(3)	(4)	(5)
$\Delta IREls_{it}$	0.008				0.009
	(0.006)				(0.006)
$\Delta IREls_{it-1}$		0.027***			0.027***
		(0.006)			(0.006)
$\Delta IREls_{it-2}$			0.011*		0.010*
			(0.006)		(0.006)
$\Delta IREls_{it-3}$				0.000	0.003
				(0.004)	(0.004)
R-squared	0.078	0.087	0.079	0.077	0.089
Observations	3,898	3,898	3,898	3,898	3,898
Country FE	Y	Y	Y	Y	Y
Baseline controls	Y	Y	Y	Y	Y

Panel (b): Two-stage least square (2SLS) results with lagged currency shares ($\Delta IREls_{it-1}$)

	Output response		AR test	OLS=IV
VARIABLES	LP-SSIV	LP-SSIV (lagged shares)	p-value	p-value
	(1)	(2)	(3)	(4)
$h=0$	-7.99	-8.35	0.109	0.295
	(5.28)	(5.32)		
$h=1$	-18.48**	-19.44**	0.008***	0.052*
	(8.08)	(7.95)		
$h=2$	-15.78*	-17.27**	0.028**	0.125
	(8.60)	(8.18)		
$h=3$	-18.25*	-20.21**	0.025**	0.095*
	(9.52)	(9.19)		
$h=4$	-9.47	-11.55	0.271	0.553
	(10.72)	(10.50)		
$h=5$	-5.70	-7.48	0.468	0.829

	(10.55)	(10.31)		
$h=6$	-0.52	-1.63	0.881	0.751
	(11.17)	(10.87)		
Joint significance	0.104	0.053*		
Instrument	ΔIRE_{it-2}	$\Delta IREls_{it-1}$		
KP weak IV	21.14	22.05		
Observations	3,898			
Country FE	Y	Y		
Baseline controls	Y	Y		

Notes: This table presents the robustness check of constructing the shift-share instrument with lagged currency shares. To be specific, I define $\Delta IREls_{it} \equiv \sum_j \text{currencyshare}_{it-1}^{j} \times \Delta\text{interestrate}_t^j$, which differs from Equation (4.2) in the lagged exposure term $\text{currencyshare}_{it-1}^{j}$ (instead of contemporaneous $\text{currencyshare}_{it}^{j}$ in the original ΔIRE_{it} specification). Panel (a) reports strong first-stage results between the binary default indicators and the new $\Delta IREls_{it-1}$. Panel (b) shows the two-stage least square (2SLS) coefficient estimates using $\Delta IREls_{it-1}$ as the instrumental variable, with column (1) reproducing the LP-SSIV baseline results from Table 3. The dependent variable is the long-difference of the log of real GDP per capita $y_{t+h} - y_{t-1}$. The independent variable is a binary indicator for sovereign default in the first year, while the shift-share instrument is the first lag of "interest rate exposure (lagged shares)" $\Delta IREls_{it-1}$. The sample period is 1970-2010. The joint significance test evaluates whether coefficient estimates in all periods $(\beta_0, \beta_1, \ldots, \beta_6)$ are simultaneously zero. The Kleibergen-Paap F-statistic for weak instruments (KP weak IV) is commonly used with a standard threshold of 10. The p-value for the Anderson-Rubin test is calculated from the Anderson-Rubin F-statistic. The p-value for the equality of OLS and IV coefficients (i.e., OLS = IV) in each period is derived using the stacking method. All regressions include country fixed effects and baseline controls specified in Table 2, as well as one lead and one lag of the instrument ($\Delta IREls_{it}$ and $\Delta IREls_{it-2}$) to account for lead-lag exogeneity and incomplete shares for currency compositions that do not sum to one (Stock and Watson 2018; Borusyak et al. 2022). Robust standard errors, clustered at the country level to address serial correlations, are reported in parentheses. Statistical significance is indicated as *** p<0.01, ** p<0.05, * p<0.1.

Table A10: Robustness check using short-term interest rates in shift-share instrument

Panel (a): First-stage results with short-term interest rates ($\Delta IREstir_{it}$)

Dependent variable: S&P first-year binary default indicator

	(1)	(2)	(3)	(4)	(5)
$\Delta IREstir_{it}$	-0.005**				0.000
	(0.003)				(0.003)
$\Delta IREstir_{it-1}$		0.000			-0.000
		(0.003)			(0.004)
$\Delta IREstir_{it-2}$			0.013***		0.013***
			(0.002)		(0.003)
$\Delta IREstir_{it-3}$				0.008***	0.006**
				(0.002)	(0.003)
R-squared	0.078	0.087	0.079	0.077	0.089
Observations	3,898	3,898	3,898	3,898	3,898
Country FE	Y	Y	Y	Y	Y
Controls	Y	Y	Y	Y	Y

Panel (b): 2SLS results with short-term interest rates ($\Delta IREstir_{it-2}$)

	Output response		AR test	OLS=IV
VARIABLES	LP-SSIV	LP-SSIV (ST interest rate)	p-value	p-value
	(1)	(2)	(3)	(4)
$h=0$	-7.99	-12.42*	0.054*	0.152
	(5.28)	(6.75)		
$h=1$	-18.48**	-23.08**	0.017**	0.064*
	(8.08)	(10.20)		
$h=2$	-15.78*	-31.09***	0.001***	0.012**
	(8.60)	(10.23)		
$h=3$	-18.25*	-34.74***	0.000***	0.006***
	(9.52)	(10.64)		
$h=4$	-9.47	-33.36***	0.000***	0.008***
	(10.72)	(10.64)		
$h=5$	-5.70	-5.67	0.53	0.953

	(10.55)	(9.00)		
$h = 6$	-0.52	12.82	0.165	0.065*
	(11.17)	(9.88)		
Joint significance	0.104	0.006***		
Instrument	ΔIRE_{it-2}	$\Delta IREstir_{it-2}$		
KP weak IV	21.14	22.24		
Observations	3,898			
Country FE	Y	Y		
Baseline controls	Y	Y		

Notes: This table presents the robustness check that constructs the shift-share instrument with short-term interest rate series (instead of long-term interest rate series) retrieved from the JST Macrohistory Database. To be specific, I define $\Delta IREstir_{it} \equiv \sum_j \text{currencyshare}_{it}^j \times \Delta\text{STinterestrate}_t^j$, which differs from Equation (4.2) in that it uses the short-term interest rates $\Delta\text{STinterestrate}_t^j$ (instead of long-term $\Delta\text{interestrate}_t^j$ in the original ΔIRE_{it} specification). Panel (a) reports strong first-stage results between the binary default indicators and the new $\Delta IREstir_{it}$. Panel (b) shows the two-stage least square (2SLS) coefficient estimates using $\Delta IREstir_{it-2}$ as the instrumental variable, with column (1) reproducing the LP-SSIV baseline results from Table 3. The dependent variable is the long-difference of the log of real GDP per capita $y_{t+h} - y_{t-1}$. The independent variable is a binary indicator for sovereign default in the first year, while the shift-share instrument is the second lag of "interest rate exposure (short-term interest rates)" $\Delta IREstir_{it-2}$. The sample period is 1970-2010. The joint significance test evaluates whether coefficient estimates in all periods $(\beta_0, \beta_1, \ldots, \beta_6)$ are simultaneously zero. The Kleibergen-Paap F-statistic for weak instruments (KP weak IV) is commonly used with a standard threshold of 10. The p-value for the Anderson-Rubin test is calculated from the Anderson-Rubin F-statistic. The p-value for the equality of OLS and IV coefficients (i.e., OLS = IV) in each period is derived using the stacking method. All regressions include country fixed effects and baseline controls specified in Table 2, as well as one lead and one lag of the instrument ($\Delta IREstir_{it-1}$ and $\Delta IREstir_{it-3}$) to account for lead-lag exogeneity and incomplete shares for currency compositions that do not sum to one (Stock and Watson 2018; Borusyak et al. 2022). Robust standard errors, clustered at the country level to address serial correlations, are reported in parentheses. Statistical significance is indicated as *** p<0.01, ** p<0.05, * p<0.1.

Table A11: Robustness check of excluding U.S. dollar debt in shift-share instrument (leave-one-out)

Panel (a): First-stage results with excluding U.S. dollar debt ($\Delta IREnoUSD_{it}$)

Dependent variable: S&P first-year binary default indicator

	(1)	(2)	(3)	(4)	(5)
$\Delta IREnoUSD_{it}$	-0.018*				-0.017*
	(0.009)				(0.010)
$\Delta IREnoUSD_{it-1}$		0.022			0.023*
		(0.014)			(0.014)
$\Delta IREnoUSD_{i-2}$			0.025**		0.016*
			(0.011)		(0.009)
$\Delta IREnoUSD_{it-3}$				0.017*	0.013
				(0.010)	(0.010)
R-squared	0.078	0.078	0.078	0.078	0.081
Observations	3,898	3,898	3,898	3,898	3,898
Country FE	Y	Y	Y	Y	Y
Controls	Y	Y	Y	Y	Y

Panel (b): 2SLS results with excluding U.S. dollar debt ($\Delta IREnoUSD_{it-2}$)

	Output response		AR test	OLS=IV
VARIABLES	LP-SSIV	LP-SSIV (exclude USD)	p-value	p-value
	(1)	(2)	(3)	(4)
$h=0$	-7.99	-1.55	0.920	0.585
	(5.28)	(15.38)		
$h=1$	-18.48**	-9.22	0.730	0.549
	(8.08)	(26.05)		
$h=2$	-15.78*	22.36	0.492	0.291
	(8.60)	(36.85)		
Instrument	ΔIRE_{it-2}	$\Delta IREnoUSD_{it-2}$		
KP weak IV	21.14	3.91		
Observations	3,898			

Country FE	Y	Y
Baseline controls	Y	Y

Notes: This table presents the robustness check that constructs the shift-share instrument while excluding the U.S. dollar denomination debt. I define $\Delta IREnoUSD_{it} \equiv \sum_{j \neq US\$} \text{currencyshare}_{it}^{j} \times \Delta \text{interestrate}_{t}^{j}$, which differs from Equation (4.2) by dropping the USD-weighted term $\text{currencyshare}_{it}^{US\$} \times \Delta \text{interestrate}_{t}^{US\$}$. Panel (a) reports first-stage results between the binary default indicators and the new $\Delta IREnoUSD_{it}$. Panel (b) shows the two-stage least square (2SLS) coefficient estimates using $\Delta IREnoUSD_{it-2}$ as the instrumental variable, with column (1) reproducing the LP-SSIV baseline results from Table 3. Since this $\Delta IREnoUSD_{it}$ variable is a weak instrument in this specification (KP F-statistic is 3.91), this table reports results only through horizon $h = 2$. The dependent variable is the long-difference of the log of real GDP per capita $y_{t+h} - y_{t-1}$. The independent variable is a binary indicator for sovereign default in the first year, while the shift-share instrument is the second lag of "interest rate exposure (excluding U.S. dollar debt)" $\Delta IREnoUSD_{it-2}$. The sample period is 1970-2010. The joint significance test evaluates whether coefficient estimates in all periods $(\beta_0, \beta_1, \ldots, \beta_6)$ are simultaneously zero. The Kleibergen-Paap F-statistic for weak instruments (KP weak IV) is commonly used with a standard threshold of 10. The p-value for the Anderson-Rubin test is calculated from the Anderson-Rubin F-statistic. The p-value for the equality of OLS and IV coefficients (i.e., OLS = IV) in each period is derived using the stacking method. All regressions include country fixed effects and baseline controls specified in Table 2, as well as one lead and one lag of the instrument ($\Delta IREnoUSD_{it-1}$ and $\Delta IREnoUSD_{it-3}$) to account for lead-lag exogeneity and incomplete shares for currency compositions that do not sum to one (Stock and Watson 2018; Borusyak et al. 2022). Robust standard errors, clustered at the country level to address serial correlations, are reported in parentheses. Statistical significance is indicated as *** p<0.01, ** p<0.05, * p<0.1.

Table A12: Persistence and unit-root tests for U.S. dollar share and the shift-share IV

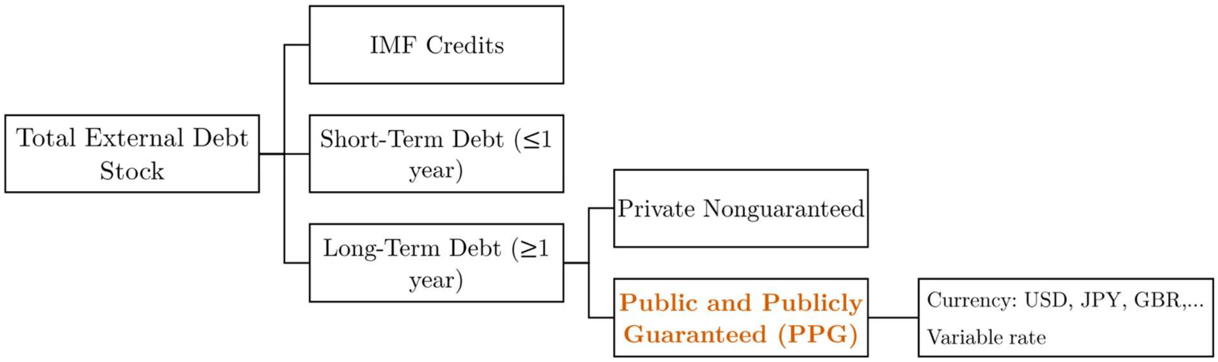
Country	U.S. dollar share		SSIV		Country	U.S. dollar share		SSIV	
	ρ	DF p-value	ρ	DF p-value		ρ	DF p-value	ρ	DF p-value
	(1)	(2)	(3)	(4)		(5)	(6)	(7)	(8)
AGO	0.88	0.53	-0.2	0.00***	LKA	1	0.96	0.2	0.00***
ALB	0.81	0.25	-0.12	0.00***	LSO	0.88	0.38	0.23	0.00***
ARG	0.78	0.28	0.17	0.00***	LTU	0.83	0.68	-0.09	0.00***
ARM	0.76	0.01***	-0.13	0.00***	LVA	0.9	0.67	-0.07	0.01***
AZE	0.73	0.33	-0.17	0.00***	MAR	0.93	0.63	0.24	0.00***
BDI	0.86	0.08*	0.12	0.00***	MDA	0.85	0.52	-0.06	0.00***
BEN	0.92	0.4	0.16	0.00***	MDG	0.81	0.08*	0.21	0.00***
BFA	0.93	0.17	0.23	0.00***	MDV	0.94	0.77	-0.05	0.00***
BGD	0.91	0.01**	0.15	0.00***	MEX	0.91	0.22	0.16	0.00***
BGR	0.97	0.89	-0.08	0.00***	MKD	0.97	0.89	-0.11	0.00***
BIH	0.79	0.41	0.06	0.18	MLI	0.98	0.87	0.2	0.00***
BLR	0.91	0.75	-0.05	0.01***	MMR	0.9	0.41	0.28	0.00***
BLZ	0.97	0.45	-0.17	0.00***	MNG	0.99	0.95	-0.07	0.00***
BOL	0.99	0.93	0.19	0.00***	MOZ	0.8	0.27	-0.09	0.00***
BRA	1	0.95	0.18	0.00***	MRT	0.74	0.04**	0.28	0.00***
BRB	0.68	0.05**	0.08	0.00***	MUS	0.87	0.24	0.16	0.00***
BTN	0.85	0.26	-0.11	0.00***	MWI	0.89	0.05*	-0.17	0.00***
BWA	0.8	0.07*	0.2	0.00***	MYS	0.94	0.71	0.23	0.00***
CAF	0.9	0.1	0.21	0.00***	NER	0.95	0.55	0.29	0.00***
CHL	0.97	0.88	0.2	0.00***	NGA	0.82	0.19	0.2	0.00***
CHN	0.95	0.25	0.02	0.00***	NIC	0.9	0.3	0.12	0.00***
CIV	0.88	0.39	0.28	0.00***	NPL	0.93	0.2	0.19	0.00***
CMR	1	0.95	0.29	0.00***	OMN	0.9	0.48	-0.04	0.00***
COD	0.92	0.76	0.18	0.00***	PAK	1.03	0.99	0.25	0.00***
COG	0.98	0.87	0.2	0.00***	PAN	0.96	0.79	0.2	0.00***
COL	1	0.96	0.17	0.00***	PER	0.97	0.87	0.17	0.00***
COM	0.78	0.06*	-0.16	0.00***	PHL	0.99	0.92	0.28	0.00***
CPV	0.98	0.92	-0.09	0.00***	PNG	0.97	0.86	0.18	0.00***
CRI	1	0.94	0.14	0.00***	POL	1.04	0.98	-0.33	0.00***
CZE	0.78	0.82	-0.36	0.01**	PRY	1.02	0.99	0.21	0.00***
DJI	1.01	0.98	.	.	ROU	0.83	0.33	-0.2	0.00***
DMA	0.82	0.05*	-0.09	0.00***	RUS	0.85	0.19	-0.09	0.00***
DOM	0.97	0.86	0.18	0.00***	RWA	0.8	0.03**	0.2	0.00***
DZA	1.02	0.98	0.26	0.00***	SDN	0.82	0.01***	0.11	0.00***
ECU	0.97	0.83	0.15	0.00***	SEN	0.84	0.06*	0.3	0.00***
EGY	0.91	0.42	0.19	0.00***	SLB	0.82	0.00***	-0.16	0.00***
ERI	0.86	0.65	-0.21	0.03**	SLE	0.93	0.45	0.12	0.00***
EST	0.45	0.00***	-0.37	0.00***	SLV	1	0.95	0.09	0.00***
ETH	0.95	0.76	-0.03	0.00***	SOM	0.83	0.00***	0.13	0.1

FJI	0.99	0.91	0.19	0.00***	SRB	0.9	0.47	-0.34	0.12
GAB	0.94	0.7	0.3	0.00***	STP	0.77	0.11	0.35	0.53
GEO	0.16	0.00***	-0.14	0.00***	SVK	0.57	0.46	-0.68	0.00***
GHA	0.92	0.47	0.15	0.00***	SWZ	1.04	0.99	0.25	0.00***
GIN	0.97	0.65	-0.17	0.00***	SYC	0.96	0.88	-0.17	0.00***
GMB	0.94	0.21	0.21	0.00***	SYR	0.94	0.46	0.09	0.00***
GNB	0.78	0.06*	0.12	0.00***	TCD	0.96	0.36	0.15	0.00***
GNQ	0.88	0.02**	-0.27	0.00***	TGO	0.93	0.48	0.17	0.00***
GRD	0.95	0.23	0	0.00***	THA	1.01	0.98	0.28	0.00***
GTM	1	0.97	-0.02	0.00***	TJK	0.98	0.9	-0.13	0.00***
GUY	0.81	0.00***	0.19	0.00***	TKM	0.89	0.63	-0.26	0.00***
HND	1.01	0.97	0.16	0.00***	TON	0.93	0.68	0.07	0.00***
HRV	0.84	0.8	-0.47	0.00***	TTO	0.59	0.04**	0.24	0.00***
HTI	1.02	0.99	0.18	0.00***	TUN	0.88	0.36	0.28	0.00***
HUN	0.8	0.6	-0.37	0.00***	TUR	0.99	0.94	0.22	0.00***
IDN	1.01	0.97	0.3	0.00***	TZA	1	0.96	-0.11	0.00***
IND	0.99	0.9	0.16	0.00***	UGA	0.9	0.07*	-0.05	0.00***
IRN	0.65	0.06*	-0.18	0.00***	UKR	0.62	0.07*	-0.12	0.00***
JAM	0.96	0.67	0.19	0.00***	URY	0.88	0.43	0.16	0.00***
JOR	0.99	0.93	0.16	0.00***	UZB	0.54	0.06*	-0.16	0.00***
KAZ	0.91	0.31	-0.08	0.00***	VCT	0.96	0.18	-0.06	0.00***
KEN	0.99	0.9	0.23	0.00***	VEN	0.83	0.27	0.15	0.00***
KGZ	0.72	0.00***	-0.11	0.00***	VNM	0.89	0.36	-0.03	0.00***
KHM	0.88	0.44	0.18	0.05**	VUT	0.91	0.28	-0.04	0.00***
KNA	0.92	0.11	-0.1	0.00***	WSM	0.77	0.00***	0.12	0.00***
LAO	1.01	0.98	0.09	0.00***	YEM	0.95	0.73	-0.14	0.00***
LBN	0.91	0.58	-0.21	0.00***	ZAF	0.64	0.19	-0.11	0.00***
LBR	0.93	0.56	0.09	0.23	ZMB	0.96	0.83	0.12	0.00***
LCA	1	0.96	-0.07	0.00***	ZWE	0.93	0.52	0.14	0.00***

Notes: This table reports persistence and unit-root test results for both the U.S. dollar currency share and the shift-share instrument ΔIRE_{it} estimated separately for each country. The AR(1) coefficient ρ is derived from the OLS regression of each series on its first lag. The "DF p-value" is the MacKinnon approximate p-value from the Augmented Dickey-Fuller unit-root test. *** p<0.01, ** p<0.05, * p<0.1

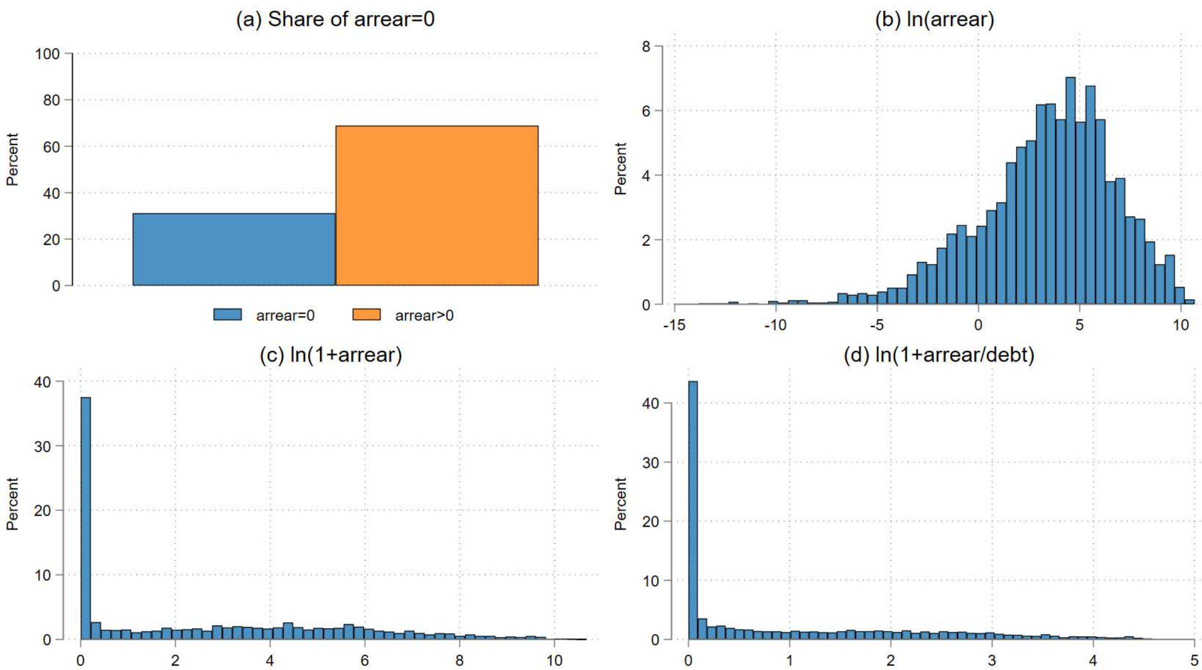
Figure A1: Data Structure of the World Banks' International Debt Statistics

Panel (a): Sector-specific



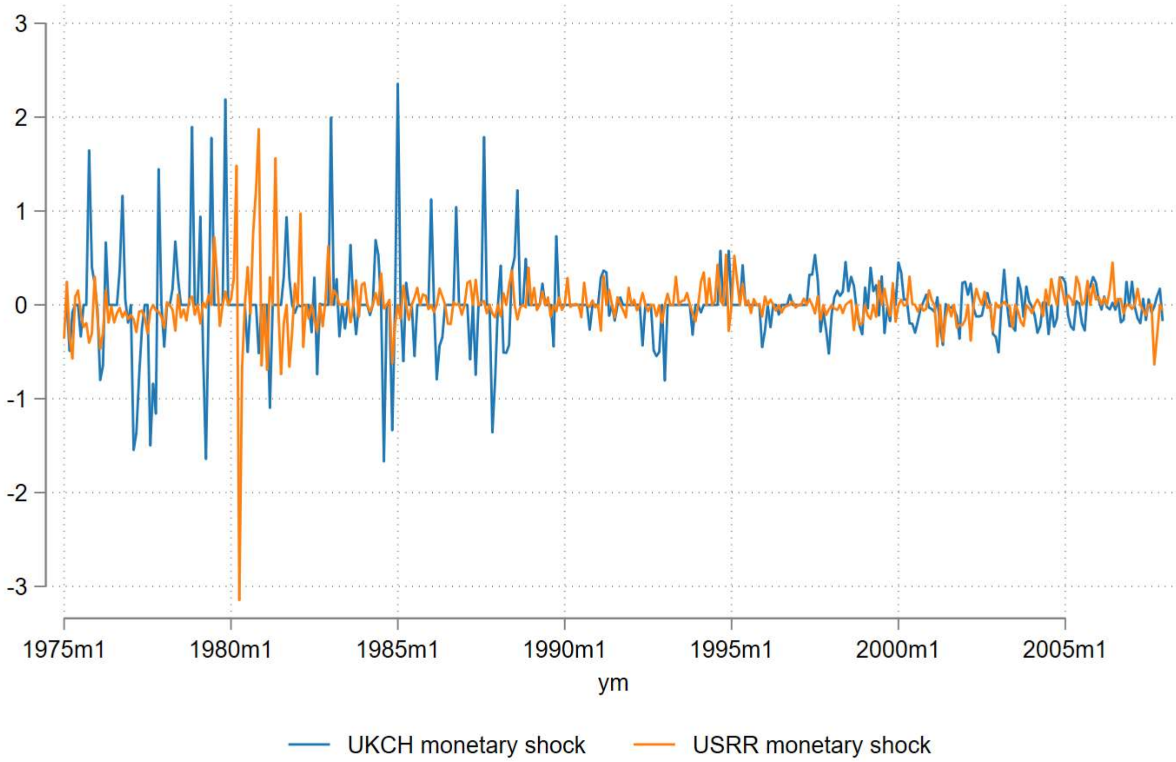
Notes: This figure summarizes the overall data structure of the World Bank's International Debt Statistics relevant to this paper. The currency denomination data are only available for long-term public and publicly guaranteed (PPG) debt. The currency denomination data are not available for short-term or private-sector debt.

Figure A2: Distribution of sovereign arrears (late repayments) for 1970–2020



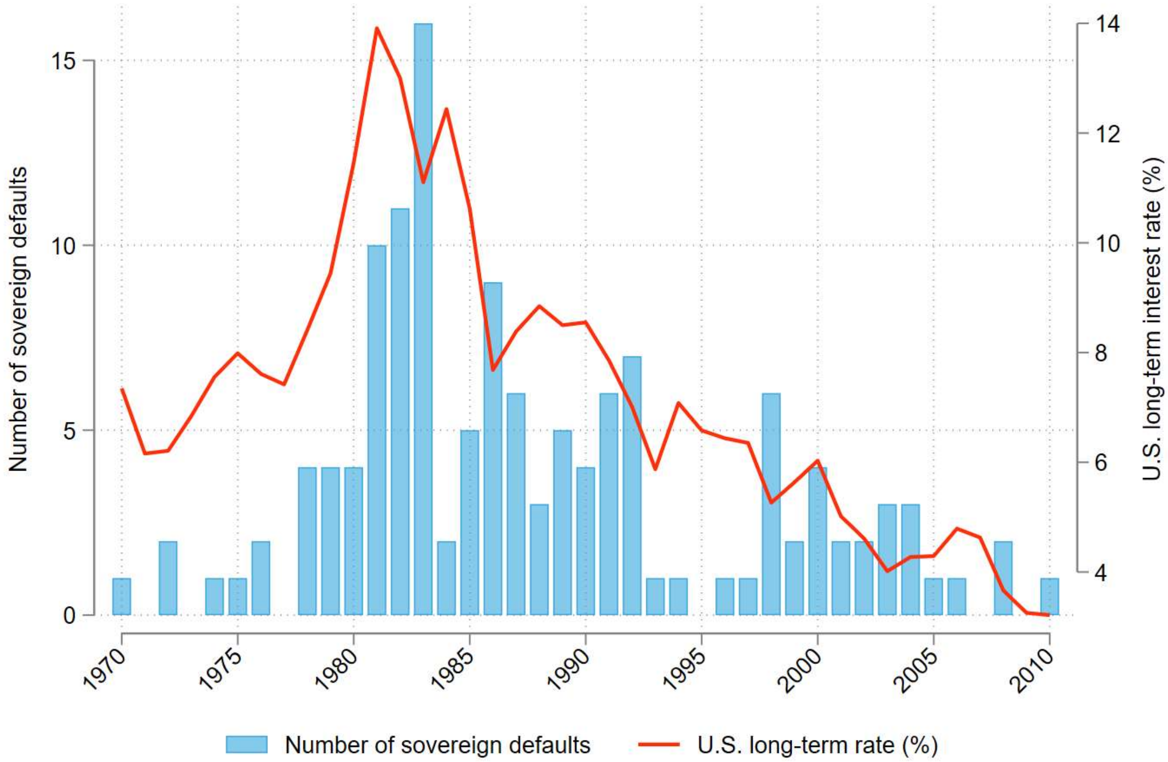
Notes: This figure shows the distribution of sovereign arrears from the World Bank's International Debt Statistics (IDS) over the 1970–2020 period. The data include the sum of principal and interest arrears for each country vis-à-vis all creditors (i.e., the world). All monetary values are expressed in current U.S. dollars, and the arrear-to-debt ratio is defined as the total arrears divided by the external debt stock in the same year. Only Panel (b) excludes 0 values, while (a), (c), and (d) use the whole sample.

Figure A3: Narrative monetary shocks from Cloyne and Hürtgen (2016)



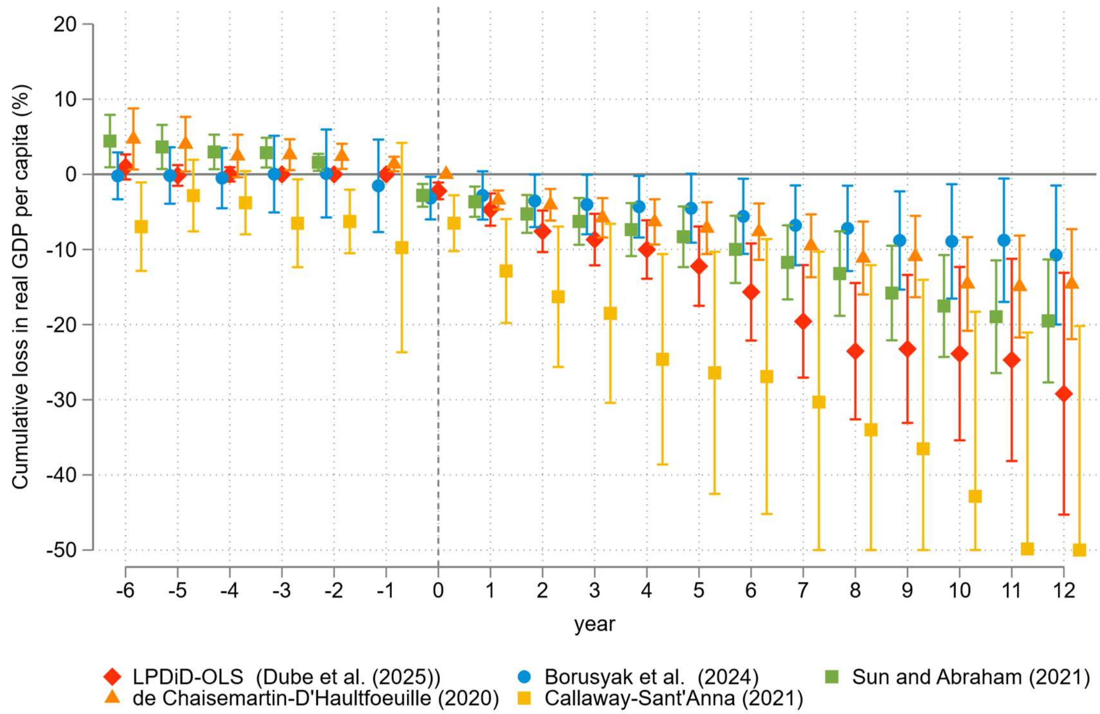
Notes: This figure plots monthly narrative monetary shocks series from Cloyne and Hürtgen (2016). Notice that the U.K. monetary shocks series is markedly more volatile and exhibits larger shocks in magnitude than its U.S. counterpart. Also, the within-year positives and negatives monthly monetary shocks often offset each other. To mitigate this attenuation, Table 4 defines the annual USRR and/or UKCH monetary shock as the single largest-magnitude monthly shock in each year, standardizes the annual series within each country, and then constructs the narrative shift-share instrument by interacting these standardized shocks with currency shares denominated in U.S. dollars and British pounds respectively.

Figure A4: Co-movement between sovereign defaults and U.S. long-term interest rates



Notes: This figure plots the annual number of sovereign defaults (left axis) and the U.S. long-term interest rate (right axis). The close co-movement highlights an identification challenge with year fixed effects: global interest rate cycles are strongly correlated with waves of sovereign default, so year fixed effects will absorb most of the variations in both the regressors (binary default indicators) and the instrument. A difference-in-differences (DiD) design, by contrast, can help isolate the variation in default decisions while explicitly controlling for time (year) fixed effects. See Section 6.3 for the DiD application, which delivers similar results. The sovereign default is the binary indicator from the S&P classification retrieved from Kuvshinov and Zimmermann (2019), and the U.S. long-term interest rates are retrieved from the JST Macrohistory database.

Figure A5: All difference-in-difference estimators on the cost of sovereign default



Notes: This figure plots the estimated cost of sovereign default using various difference-in-difference (DiD) estimators that address staggered treatment in the literature. Relative to Figure 5, it additionally reports results from Callaway-Sant'Anna (2021). All DiD specifications include country and year fixed effects, as well as the baseline controls specified in Table 3. The LPDiD specification uses the long-difference in log real GDP per capita as the dependent variable and first-differences of controls variables (up to two lags). Assuming the effects of negative weights from repeated defaults (i.e., contamination effects) dissipate after 6 years, the clean control set is defined as a 6-year window, restricting the sample to countries that have not defaulted in the past 6 years. Following the standard practice in the literature, the other DiD estimators use the level of log real GDP per capita as the dependent variable, include only lagged (not contemporaneous) controls in level, and exclude lagged outcome. Robust standard errors are clustered at the country level, and 90% confidence intervals are shown in the figure. See the main text for reference on these DiD estimators.