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Three Essays in Applied Macroeconomics and International Economics

By

JEFF KIN WAI CHEUNG

DISSERTATION

Submitted in partial satisfaction of the requirements for the degree of

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of the

UNIVERSITY OF CALIFORNIA

DAVIS

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2026

To Mr. Kwong Wai-Shing, Miss Jennifer Yu, the Wesleyan Freeman Asian Scholarship Program, and my family, for making this dissertation possible.

ABSTRACT

This dissertation contributes to the applied macroeconomics and international economics literatures. Using applied macro-econometric methods, I study three central questions in macroeconomics and international finance through three independent essays.¹

How costly is sovereign default? In [Chapter 1](#), 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 variation in developing countries' currency denomination of external debt and quasi-random interest rate movements in advanced economies to construct a novel shift-share instrumental variable for the default decisions. Using a local projection shift-share instrumental variable (LP-SSIV) framework, I causally estimate that sovereign defaults on average lead to an immediate 8% decline in real GDP per capita in the first year. The cumulative output losses peak at 18.5% in the second year, remain economically significant through the fourth year, and gradually dissipate by the sixth year. These temporary baseline LP-SSIV estimates, which I interpret as the policy-relevant *equilibrium* effects of sovereign default, are robust to narrative monetary shocks and alternative arrears-based continuous partial default measures. In contrast, specifications that account for the remaining unobserved mitigating channels—including spillover-corrected control function estimates, residualized LP-SSIV specifications, and difference-in-differences (DiD) estimators—produce similar short-run effects but substantially more long-lasting output losses. I interpret these persistent estimates as capturing the underlying *structural* cost of sovereign default. Lastly, I find that floating exchange rate regimes and lower external debt levels, especially short-term debt, effectively attenuate the output losses following default.

¹In accordance with the UC Responsible AI Principles, I acknowledge the use of ChatGPT for editing and revising this dissertation. I also gratefully acknowledge the excellent research assistance of Alejandro Munoz and Raphael Zhang for [Chapter 2](#) and [Chapter 3](#) respectively. All ideas, analysis, coding, writing, interpretations, and any remaining errors are entirely my own.

Are emerging markets’ central banks credibly committed to inflation targeting? In Chapter 2, I apply the “Optimal Policy Perturbation” (OPP) framework of [Barnichon and Mesters \(2023\)](#) to evaluate monetary policy credibility across emerging economies. Combining impulse responses from a Bayesian VAR with professional forecasts from the Economist Intelligence Unit, I exploit the dual interpretation of the OPP statistic. First, under the assumption that observed policy rates are locally optimal given policymakers’ information sets, I adopt a revealed-preference approach and recover time-varying implicit policy weights by minimizing deviations of the OPP statistic from zero. Second, under fixed benchmark weights, I use the OPP statistic to assess whether policy rates could have been adjusted to further reduce the policymaker’s loss, and to quantify the direction and magnitude of such adjustments. The evidence indicates that inflation-target regimes in emerging markets have become increasingly credible since the early 2000s, with stronger anchoring of inflation expectations and reduced inflation persistence. While central banks face sharper trade-offs during global crises, monetary policy behavior typically reverts toward inflation stabilization, highlighting the flexibility and resilience of modern inflation-target frameworks.

Are uncovered interest parity (UIP) coefficient estimates state-dependent? In Chapter 3, I construct a novel historical dataset to revisit the long-standing UIP puzzle. The UIP condition predicts that the higher-interest-rate currency should depreciate so as to equalize expected returns across currencies. Yet despite its central role in international macro models, this condition is often rejected in the empirical data. This paper argues that UIP coefficient estimates are inherently regime-dependent, a latent feature that becomes visible once one studies their term structure. Using historical Eurocurrency/LIBOR deposit rates at the 1-, 3-, 6-, and 12-month maturities, I document three empirical facts. First, rolling-window UIP coefficient estimates are negatively related to the volatility of interest rate differentials, and this monotonic relationship evolves gradually over time rather than shifting discretely around the 2008 Global Financial Crisis. Second, both the sign and the magnitude of UIP coefficients vary across maturities: under normal times, short-maturity currency trades earn

larger premia in high-volatility states, whereas near the zero lower bound the premium shifts toward longer-maturity deposits. Third, these maturity patterns are difficult to reconcile with a pure “delayed-overshooting” explanation based solely on sluggish exchange rate responses to monetary shocks, but they are also hard to explain using a standard finance-based “risk premium” argument alone. Instead, the term-structure evidence from UIP regressions points to the coexistence of both mechanisms, with time-varying risk premia—shaped in part by the signs of yield curve slopes—playing a central role in the UIP puzzle.

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Wise men say that a PhD journey is never an easy path. Near the end of this long marathon, I finally came to realize that it is not merely an intellectual challenge, but also an emotional journey. Every moment of discovery often arrives alongside countless obstacles and moments of self-doubt. My opportunity to experience the joy of learning and discovery owes entirely to the immense generosity, kindness, and support I have received from the people who accompanied me throughout my life, both during and long before my PhD journey.

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Chapter 1

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

1.1 INTRODUCTION

How costly is sovereign default? The seminal [Eaton and Gersovitz \(1981\)](#) model posits that defaulting countries lose access to international capital markets, yet this mechanism alone struggles to explain the high levels of external debt typically observed prior to default episodes. To address this discrepancy, subsequent research has incorporated direct output losses following default to better reconcile theoretical models with empirical observations ([Arellano, 2008](#); [Yue, 2010](#); [Mendoza and Yue, 2012](#); [Uribe and Schmitt-Grohé, 2017](#)). In principle, however, a country that can default without facing enforceable penalties—since there are no international courts capable of enforcing sovereign debt contracts—may instead benefit from immediate debt relief. As a result, the direction and magnitude of the causal effect of sovereign default on output losses remain important empirical questions in the international macroeconomics literature.

Panel (a) of [Table 1.1](#) motivates the paper by presenting the observed output paths following four representative sovereign default or debt crisis episodes: Argentina (2001), Russia (1998), Ecuador (1999), and Greece (2010). Across these episodes, sovereign default

is typically followed by an immediate recession of roughly 6% decline in real GDP per capita, but substantial heterogeneity exists across episodes. Russia recovered quickly following its 1998 default, whereas Argentina and Ecuador experienced more gradual recoveries over the subsequent four years. By contrast, Greece’s debt crisis was associated with a considerably more prolonged recession. Overall, on average, sovereign defaults are associated with substantial output losses that are comparable in magnitude but considerably more persistent compared to those observed after major currency and financial crises reported in Panels (c) and (d). These empirical patterns also raise an important question: what determines the magnitude and persistence of default-induced recessions?

Since sovereign defaults typically occur during periods of severe economic distress, the descriptive evidence alone cannot establish a causal relationship between sovereign default and recessions. To estimate the causal effect of sovereign default on output losses, I introduce a novel empirical strategy—local projection with shift-share instrumental-variable (LP-SSIV) framework building on [Jordà, Schularick and Taylor \(2020\)](#) and [Jordà, Singh and Taylor \(2024\)](#). The shift-share instrument combines developing countries’ currency denomination of external debt as endogenous shares, as well as quasi-random interest rate movements in advanced economies as exogenous shifts. The underlying mechanism is intuitive: when an advanced economy raises its interest rates, its currency appreciates, increasing the local-currency debt burden denominated in that appreciated currency. Countries with greater exposure to that appreciating currency therefore face larger increases in debt-service costs and a higher default probability. Embedding this shift-share intuition within a local-projection (LP) framework allows me to trace the dynamic causal effects of sovereign default on output, thereby quantifying both the magnitude and persistence of default-induced recessions.

The 1980s Latin American Debt Crisis illustrates this mechanism vividly. Following the recycling of petrodollars after the 1970s oil shocks, many Latin American countries accumulated substantial amounts of debt denominated in U.S. dollars. In 1979, U.S. interest rates rose sharply and unexpectedly under Paul Volcker’s monetary tightening, causing the

dollar to appreciate and substantially increasing their debt-service burdens. The result was a widespread wave of sovereign defaults beginning with Mexico in 1982, followed by a prolonged period of debt restructuring and sluggish economic growth commonly referred to as Latin America’s “lost decade.”

To preview the main results, the baseline LP-SSIV estimates show that on average, sovereign default causes an immediate 8% decline in real GDP per capita in the first-year. The cumulative output losses reach nearly 18% in the second year and remain stable through the fourth year before gradually fading to zero—that is, returning to trend—by the sixth year, suggesting that the effects are substantial but ultimately temporary. In addition to the typical binary default indicators, continuous arrears-based partial default measures—after accounting for the extensive margin—also deliver similar causal estimates. The baseline LP-SSIV estimates are robust to a broad set of macroeconomic controls, narrative monetary policy shocks, and recent methodological guidance from the shift-share designs, panel instrumental variables, and time-series regressions (Christian and Barrett, 2024; Borusyak, Hull and Jaravel, 2022, 2025; Borusyak and Hull, 2024; Stock and Watson, 2018). Taken together, the baseline LP-SSIV results provide strong empirical evidence that sovereign default imposes sizeable output costs.

A central challenge for the baseline LP-SSIV interpretation is the exclusion restriction: it is possible that foreign monetary tightening affects economic activity through channels other than sovereign default, in particular through the trade channel—sovereign default is often accompanied by large currency depreciations—an empirical pattern known as the “Twin Ds”—which may stimulate exports and partially offset the adverse consequences of default. To address this concern, I augment the baseline LP-SSIV framework with a control-function approach (Jordà, Singh and Taylor, 2024; Jordà, Schularick and Taylor, 2020; Wooldridge, 2015; Conley, Hansen and Rossi, 2012). To be specific, I use never-defaulting countries to estimate the direct spillover effects of the shift-share instrument on output growth and thereby construct spillover-corrected informative bounds on the causal cost of sovereign default. While

the short-run effects remain broadly similar to the baseline LP-SSIV estimates, the long-run output losses become substantially more persistent, with output remaining between -10% to -30% below trend even six years after default. These findings remain robust when the comparison group is restricted to observationally similar countries using propensity-score matching based on either long-run averages or pre-default characteristics.

I further investigate whether observed interest rate changes adequately isolate exogenous monetary policy variation. A potential concern is that realized interest rate movements may contain predictable components related to domestic macroeconomic conditions or broader global forces that simultaneously affect multiple advanced economies. Following [Jordà, Schularick and Taylor \(2020\)](#) and [Jordà, Singh and Taylor \(2024\)](#), I residualize interest rate changes abroad with respect to a broad set of macroeconomic controls in advanced economies, as well as the time-varying commodity price movements. The resulting residualized LP-SSIV estimates produce output paths that closely resemble the baseline LP-SSIV estimates in the short run but again yield substantially more persistent long-run output losses, ranging from -10% to -30% below trend after six years, which closely align with the spillover-corrected LP-SSIV estimates. Similar patterns emerge from alternative shift-share specifications proposed in the recent literature, including controls for interactions between year fixed effects and incomplete shares as well as year-level re-centered instruments ([Borusyak and Hull, 2023](#); [Borusyak, Hull and Jaravel, 2025, 2022](#)).

Taken together, these findings suggest that global time-varying factors and mitigating trade cushions through currency depreciation tend to attenuate the baseline LP-SSIV estimates. However, sovereign defaults often cluster within major historical episodes, implying that aggressively absorbing global time variation may also remove an important source of identification. To address this concern, I complement the LP-SSIV analysis with a difference-in-differences (DiD) framework. The DiD estimators directly control for global time trends through year fixed effects, are robust to spillovers, and rely on an alternative identification assumption based on parallel trends. Most importantly, in the sovereign default settings,

countries may repeatedly default over the sample period (e.g., Argentina), so recent advances in the DiD literature provide a timely and useful framework for addressing the negative weight concerns that arise under staggered treatment.

Motivated by the finding that the baseline LP-SSIV estimates largely dissipate within six years, I first implement a local projection difference-in-differences (LP-DiD) estimator following Dube et al. (2025). The resulting LP-DiD estimates closely match the baseline LP-SSIV results in the short run while producing considerably more persistent long-run output losses, comparable to those obtained from the spillover-corrected and residualized LP-SSIV specifications. Alternative DiD estimators proposed in the literature yield similar conclusions.

How should we interpret these results? I interpret the spillover-corrected LP-SSIV estimates, the residualized LP-SSIV estimates, and the LP-DiD estimates as capturing the underlying *structural* cost of sovereign default after accounting for all mechanisms that partially offset its adverse economic consequences, such as the export booms associated with post-default currency depreciation. By contrast, I interpret the baseline LP-SSIV estimates as capturing the policy-relevant *equilibrium* effects of sovereign default, since these offsetting channels are present in actual default episodes and therefore shape the realized economic costs observed by policymakers.

The LP-SSIV framework also allows me to explore heterogeneity in default costs. Within a state-dependent local-projection framework, I find that countries under pegged exchange rate regimes experience substantially larger output losses than those with flexible exchange rates, consistent with their limited ability to benefit from post-default depreciation. Heavily indebted countries also tend to experience larger short-run output losses, especially when a substantial share of their external liabilities consists of short-term debt with maturities of less than one year, although these estimates are less precisely measured. Interestingly, defaults occurring despite adequate foreign exchange reserve holdings—sufficient to cover at least three months of imports—are associated with larger output losses, possibly because

such episodes signal unwillingness rather than inability to repay.

Last but not least, to rigorously justify the LP-SSIV framework, I develop a simple probabilistic sovereign default model based on [Cole and Kehoe \(2000\)](#). The corner solutions in this multiple-equilibria model naturally yield a local average treatment effect (LATE) interpretation: the estimated causal effects apply only to countries induced to default by foreign monetary tightening. I further draw on the seminal [Eaton and Kortum \(2002\)](#) probabilistic Fréchet model to characterize frictions that limit rapid rebalancing of sovereign debt portfolios. Together, these assumptions generate a shift-share representation of sovereign default probabilities and provide theoretical foundations for the empirical LP-SSIV design.

1.1.1 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, 2013](#)), 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, Hoggarth and Saporta, 2009](#); [Borensztein and Panizza, 2009](#); [Levy Yeyati and Panizza, 2011](#)). 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](#); [Furceri and Zdzienicka, 2012](#); [Kuvshinov and Zimmermann, 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, Graf von Luckner and Reinhart \(2024\)](#) 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 LP-SSIV estimates 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-differences approaches to estimate causal effects in fields such as international trade and immigration (Autor, Dorn and Hanson, 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, Kolesár and Morales, 2019; Borusyak, Hull and Jaravel, 2022; Goldsmith-Pinkham, Sorkin and Swift, 2020; Borusyak and Hull, 2024; de Chaisemartin and DHaultfoeuille, 2020; Roth et al., 2023; Sun and Abraham, 2021; Borusyak, Hull and Jaravel, 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—by Eichengreen and Hausmann (1999). On the theoretical side, Coppola, Krishnamurthy and Xu (2024) 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 1.2](#) develops a novel probabilistic sovereign default model that motivates the key mechanism underpinning the empirical strategy. [Section 1.3](#) describes the data. [Section 1.4](#) outlines the empirical strategy. [Section 1.5](#) presents the baseline LP-OLS and LP-SSIV results, together with the control function approach to address spillovers. [Section 1.6](#) presents robustness checks, including narrative monetary shocks and difference-in-differences estimates. [Section 1.7](#) discusses the baseline results and benchmarks them against historical crises. [Section 1.8](#) examines the state-dependent heterogeneous output cost of sovereign default, and [Section 1.9](#) concludes.

1.2 MOTIVATING THE SHIFT-SHARE REPRESENTATION OF SOVEREIGN DEFAULT

This section presents a simple probabilistic framework that motivates the shift-share instrument used in the empirical analysis. Consider a country that borrows in different foreign currencies and services its external obligations using domestic fiscal resources. Since debt repayment is denominated in foreign currencies, exchange rate fluctuations can alter the local-currency cost of debt service and, consequently, sovereign default risk.

Section 1.10 builds on Romer (2019) and Cole and Kehoe (2000) to show that a contractionary foreign monetary shock raises default probability through two channels. First, exchange rate depreciation increases the local-currency value of foreign-currency denominated liabilities, thereby immediately raising repayment burdens and contemporaneous default risk. Second, under adaptive expectations, investors interpret higher current default probability as informative about future repayment risk, which increases perceived default likelihood in subsequent periods and raises future rollover interest rates, further worsening future repayment conditions and default risk. Through this repeated interaction between debt-service costs and rollover premia, a foreign monetary tightening can generate a self-fulfilling default equilibrium.

The framework has three implications for the empirical design. First, Section 1.10.5 shows that the pricing equation can be interpreted either as a binary default probability or as an expected return conditional on repayment, providing a natural link to the empirical analysis using sovereign arrears as partial default measures in Section 1.6.3.

Second, Section 1.11 establishes the existence of an interior steady-state equilibrium, and then Section 1.12 shows that this probabilistic framework admits multiple equilibria that connect naturally to a *local average treatment effect* (LATE) interpretation. The instrumental variable strategy therefore identifies the causal effect of sovereign default on output loss for countries whose default probability increases in response to foreign monetary tightening (i.e., the “compliers”).¹

Third, and most importantly for the baseline empirical specification in Section 1.5, this framework delivers a closed-form default probability that admits a shift-share representation in response to foreign monetary shocks. Motivated by the extreme exchange rate depreciations frequently observed in developing countries, Section 1.13 and Section 1.14 show that, under Fréchet distributed exchange rate depreciations, the linear approximation

¹Consequently, the causal estimates in this paper do not speak to episodes such as the 2010 Greece’s debt crisis, which occurred during a period of monetary easing in advanced economies following the 2008 Global Financial Crisis.

of latent default likelihood can be expressed as foreign interest rate changes weighted by currency-specific exposure shares. This result provides the theoretical motivation for interacting currency-denomination shares with foreign monetary shocks to instrument for sovereign default decisions.

1.2.1 Fréchet-Distributed Exchange Rates and the Shift-share Representation of Default Probability

The empirical shift-share instrumental variable (SSIV) design exploits differential exposure—measured by the currency denomination of external debt—to exogenous foreign monetary shocks. Building on the seminal [Eaton and Kortum \(2002\)](#) framework, I show how foreign monetary shocks generate a shift-share representation of sovereign default probability through currency-specific repayment obligations driven by foreign monetary tightening. The key prediction is straightforward: default probability responds more strongly to foreign monetary tightening when a larger share of a country’s foreign-currency liabilities is denominated in the currency that appreciates following the monetary tightening. Consequently, countries with greater exposure to that particular currency experience larger increases in debt-service costs and, therefore, larger increases in default risk. This prediction is consistent with the first-stage evidence presented in [Section 1.5.1](#).

Assume perfect competition in each currency market where a continuum of fund providers lends currency k at the same currency-specific gross interest rate R_t^{ik} , determined at time t and due at $t + 1$. Let $\mathcal{E}_{t+1}^{ik}$ denote the gross depreciation factor of debtor i ’s currency against foreign currency k between t and $t + 1$, where $\mathcal{E}_{t+1}^{ik} > 1$ indicates a depreciation of debtor i ’s currency against k in time $t + 1$.

Motivated by extreme depreciation episodes frequently observed in developing countries, assume that the gross exchange rate depreciations $(\mathcal{E}_{t+1}^{i1}, \mathcal{E}_{t+1}^{i2}, \dots, \mathcal{E}_{t+1}^{iK})$ are random variables drawn independently (but not necessarily identically) across currencies from Fréchet distributions with a common debtor-specific shape parameter θ^i but currency-specific location

parameters $\bar{e}^k$; that is, for all k , $\mathcal{E}_{t+1}^{ik} \sim \text{Fréchet}(\bar{e}^k, \theta^i)$.

For tractability, assume initially that debtor i borrows equal amounts in each currency, and debtor i 's fiscal capacity $\mathcal{T}_{t+1}^i$ also follows a Fréchet distribution with the same shape parameter θ^i and debtor-specific location parameter $\bar{t}^i$.² The corresponding cumulative distribution functions are:

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

Using the max-stable property of the Fréchet distribution, debtor i defaults whenever its fiscal capacity cannot cover the largest currency-specific repayment obligation:

$$\pi_{t+1|t}^i = \mathbb{P}\left(\mathcal{T}_{t+1}^i \leq \max_k \{\mathcal{R}_t^{ik} \mathcal{E}_{t+1}^{ik}\}\right) \quad (1.1)$$

Section 1.13 derives the following closed-form solution for the default probability:

$$\pi_{t+1|t}^i = \frac{\sum_k \bar{e}^k (R_t^{ik})^{\theta^i}}{\bar{t}^i + \sum_k \bar{e}^k (R_t^{ik})^{\theta^i}} \quad (1.2)$$

Intuitively, default risk increases with aggregate foreign-currency debt-service obligations, $\sum_k \bar{e}^k (R_t^{ik})^{\theta^i}$, and decreases with fiscal capacity, $\bar{t}^i$. It also depends on the Fréchet shape parameter θ^i , which governs the tail-thickness and hence the likelihood of extreme depreciation realizations, and the currency-specific location parameters $\bar{e}^k$, which govern average depreciation levels.

Assume debtor i has strictly positive fiscal capacity (i.e., $\bar{t}^i > 0$) and consider a

²These assumptions are intended to provide a tractable representation of sovereign default probability in environments characterized by large tail events. Emerging market currencies frequently experience extreme depreciation episodes (e.g., the Thai baht, Indonesian rupiah, and Malaysian ringgit during the 1997 Asian Financial Crisis, or the Mexican peso during the 1994–1995 Tequila Crisis). At the same time, their fiscal capacity can be highly volatile because of commodity-price fluctuations, capital-flow reversals, and external financing conditions, which can generate windfall revenues and expenditure booms but later undermine repayment capacity during currency crises (e.g., oil exporters such as Venezuela and Nigeria during the 1970s oil crisis). A closer examination of the interaction between fiscal capacity, exchange rates, and currency-share stability is left for future work.

single contractionary monetary shock to foreign currency \$ (i.e., $\hat{R}_t^{\$} > 0$ and $\hat{R}_t^{ik} = 0$ for all $k \neq \$$). Linearizing Equation (1.2) around the steady state delivers the shift-share relationship motivating the empirical specification in Section 1.4:

$$\hat{\pi}_{t+1|t}^i = \theta^i (1 - \bar{\pi}^i) \left(\frac{\bar{e}^{\$} (\bar{R}^{i\$})^{\theta^i}}{\sum_k \bar{e}^k (\bar{R}^{ik})^{\theta^i}} \right) \hat{R}_t^{i\$} \quad (1.3)$$

where $\hat{\pi}_{t+1|t}^i \equiv \frac{d\pi_{t+1|t}^i}{\bar{\pi}^i}$ and $\hat{R}_t^{i\$} \equiv \frac{dR_t^{i\$}}{\bar{R}^{i\$}}$ denote percentage deviations from steady-state values.³

Intuitively, the term $\left(\frac{\bar{e}^{\$} (\bar{R}^{i\$})^{\theta^i}}{\sum_k \bar{e}^k (\bar{R}^{ik})^{\theta^i}} \right)$ represents the model-implied exposure of debtor i to monetary shocks originating from currency \$. Consequently, changes in default probability $\hat{\pi}_{t+1|t}^i$ are proportional to exogenous foreign monetary shocks $\hat{R}_t^{i\$}$ scaled by predetermined debtor i 's exposure to the foreign currency appreciated after monetary tightening.

This expression provides the theoretical link between sovereign default probability and the shift-share instrument used in the empirical analysis. The identifying variation comes from foreign monetary shocks interacted with predetermined currency-denomination shares. For this exposure measure to be informative, currency shares should be sufficiently stable over short horizons so that variation in default probability is driven primarily by exogenous foreign monetary shocks rather than contemporaneous portfolio reallocation across currencies. This condition follows naturally from the institutional features of sovereign borrowing: repayment currency is typically pre-specified in advance in sovereign debt contracts, and sovereign debt often has long maturity. As a result, currency exposure is largely predetermined when shocks are realized and cannot be immediately adjusted ex post.⁴

Consistent with this view, Figure 1.6 plots the time trends of the currency shares for six major emerging markets over the 1970–2020 period. The currency-denomination

³Since default probabilities and interest rates are easier to interpret in percentage points, Section 1.14 also derives an expression for that case. Define $\tilde{\pi}_{t+1|t}^i \equiv \pi_{t+1|t}^i - \bar{\pi}^i$ and $\tilde{R}_t^{i\$} \equiv R_t^{i\$} - \bar{R}^{i\$}$, both expressed as percentage point deviations from steady-state values. Then we have $\tilde{\pi}_{t+1|t}^i = \left(\frac{\theta^i \bar{\pi}^i (1 - \bar{\pi}^i)}{\bar{R}^{i\$}} \right) \left(\frac{\bar{e}^{\$} (\bar{R}^{i\$})^{\theta^i}}{\sum_k \bar{e}^k (\bar{R}^{ik})^{\theta^i}} \right) \tilde{R}_t^{i\$}$.

⁴Also, creditors holding dollar-denominated claims would generally not regard peso repayment as an equivalent substitute without compensation for exchange rate risk.

shares are relatively sticky and persistent over time—they do not exhibit immediate large-magnitude adjustments following foreign monetary tightening. [Section 1.15](#) also discusses the determinants and sensitivity of currency shares to foreign monetary tightening from a theory perspective. Together, these results motivate the use of the shift-share instrument to causally estimate the output cost of sovereign default.

1.3 DATA DESCRIPTION

Panel (a) of [Table 1.2](#) reports sample counts of the dataset, which consists of 135 developing countries over the 1970–2010 period. Since the World Bank reports the currency denomination of external debt only for developing countries, the sample does not include any advanced economies. The analysis is conducted at the annual frequency using an unbalanced panel with approximately 3,898 country-year observations over 41 years. Throughout the sample period, there are 101 sovereign default episodes involving 61 distinct defaulting countries. Prior to 2001, the currency denomination data distinguish seven major currencies; following the introduction of the euro, the Deutsche mark and French franc, together with other legacy Eurozone currencies, are consolidated into a single euro category. In total, the sample contains 286 currency-year shock observations. [Table 1.9](#) provides the full list of countries in sample.

1.3.1 Currency Shares and Monetary Shocks

1.3.1.1 Data Sources and Description

For the endogenous currency shares, I retrieved the data on the currency compositions of long-term public and publicly guaranteed debt for developing countries from the *International Debt Statistics* (IDS) published by the World Bank Group⁵. This country-year-currency

⁵Due 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

specific variable covers external debt 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 francs⁶.

For the exogenous monetary shifts, I retrieve long-term interest rate series (typically government bond yields) for the six advanced economies over the 1970–2010 period from the *JST Macroeconomy Database* (Jordà, Schularick and Taylor, 2017). In the baseline specification, interest rate changes serve as the monetary shift component of the shift-share instrument. As an alternative measure, Section 1.5.6 constructs residualized interest rate changes by removing the predictable components from observable domestic macroeconomic variables in those advanced economies. In addition, Section 1.6.2 uses narrative monetary shock series for the United States and the United Kingdom from Cloyne and Hurtgen (2016).

1.3.1.2 Summary Statistics and Shift-share Diagnostics

Panels (b)–(d) of Table 1.2 report summary statistics and preliminary shift-share diagnostics following Borusyak, Hull and Jaravel (2025). First, Panel (b) summarizes the currency share of external debt. Although U.S.-dollar-denominated debt accounts for the largest share of external borrowing, there remains substantial cross-country and time variation in exposure to other advanced-economy currencies. Notably, before 2001, the 90th percentiles—computed across all country-year observations with non-negative shares (including zeros) for a given

from the *Global Development Finance* (the predecessor of the IDS) using their 2006 and 2010 CD-ROMs, along with their corresponding publications from earlier years. See Table 1.10 for more details.

⁶Data on currency shares of short-term debt (i.e., obligations with maturities under one year) are not available. Since 2001, 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 debt 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 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 1.7 for the overall structure of the dataset.

currency—of the Deutsche mark and French franc shares are 13.3% and 21% respectively; after 2001, the corresponding 90th percentile of the euro share rises to 45.2%. These patterns indicate meaningful heterogeneity in countries' exposure to foreign monetary shocks.

Consistent with this observation, [Figure 1.6](#) plots the evolution of currency shares for representative emerging markets over time. Two patterns stand out: first, although the U.S. dollar borrowing remains dominant, the Japanese yen also serves as an important funding currency for some Asian borrowers, implying heterogeneous exposure to foreign monetary shocks. Second, currency shares evolve gradually and do not respond immediately to U.S. monetary tightening episodes, illustrated by the grey dashed lines. This evidence supports the interpretation of currency denomination shares as predetermined exposure measures to foreign monetary shocks.

Panel (c) reports concentration diagnostics for the structure of the currency shares. I compute effective sample sizes based on the inverse of Herfindahl (HHI) index at the currency-year, year, and currency levels. The effective sample size is approximately 80 at the currency-year level and 39 at the year level, indicating that identifying variation is broadly distributed across both dimensions. By contrast, the effective sample size at the currency level is 2.2, reflecting the dominant role of the U.S. dollar and euro-area currencies in international borrowing. Although this concentration is an inherent feature of the historical data rather than a modeling choice, it motivates the use of complementary empirical approaches, such as the difference-in-differences specifications presented later in the paper. Overall, the diagnostics suggest that the identifying variation is not driven by a small number of currency-year observations, while also highlighting the importance of validating the baseline results using alternative empirical strategies.

Panel (d) reports summary statistics for the monetary shock measures. Across the advanced economies in the sample, the average of the annual long-term interest rate changes is around -0.1% pt, with an interquartile range of roughly 1% pt. Column (4) reports AR(1) persistence estimates obtained by regressing each monetary shock measure on its first lag

(with country fixed effects and clustered standard error at the country level). The coefficient estimate of 0.218 indicates that interest rate changes are only moderately persistent. By contrast, the residualized interest rate changes discussed in [Section 1.5.6](#) exhibit little serial correlation, suggesting that residualization removes much of the predictable component of foreign interest rate movements.

1.3.2 Binary Default Indicators

For binary sovereign default classifications, I use the *Sovereign Default Database* compiled by [Kuvshinov and Zimmermann \(2019\)](#), who have integrated several country-year varying default classifications from the literature, including [Beers and Chambers \(2006\)](#), [Beim and Calomiris \(2001\)](#), [Laeven and Valencia \(2020\)](#), [Reinhart and Rogoff \(2011\)](#), and [Reinhart and Trebesch \(2016\)](#). This comprehensive database allows robustness checks across alternative sovereign default classifications.

In the baseline LP-SSIV specification, I use the first-year 0/1 sovereign default indicators based on the Standard and Poor’s classification reported in [Kuvshinov and Zimmermann \(2019\)](#).⁷ Broadly speaking, sovereign default events are relatively rare. Over the 1970–2010 sample period, there are 101 sovereign default episodes across 61 distinct countries, suggesting that some countries may have repeatedly defaulted over time.

Such repeated treatment episodes may result in negative weighting problems in staggered-treatment settings. Therefore, I additionally use in-default indicators, which also identify the ending year of each default episode, from the same dataset for additional robustness checks using difference-in-differences estimators ([Borusyak and Hull, 2024](#); [Dube et al., 2025](#)).

⁷Since Panel (a) of [Table 1.2](#) shows that sovereign debt can be denominated in different currencies, a natural concern is whether countries selectively default on debt denominated in some currencies but not others. In practice, such selective defaults are relatively uncommon. Within a given legal jurisdiction, sovereign bonds are often issued under the “pari passu” principle (Latin for “equal footing”), implying equal treatment of creditors and no contractual seniority across currency denominations ([Aguar and Amador, 2021](#); [Wright, 2014](#); [Schumacher, Trebesch and Enderlein, 2021](#)).

1.3.3 Continuous Partial 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, Mateos-Planas and Ríos-Rull, 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 arrears data—late or missed debt service payments—from the International Debt Statistics as a measure of partial default. [Figure 1.8](#) 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 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 1.6.3](#)).

1.3.4 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 world 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, Reinhart and Rogoff \(2019\)](#).

[Table 1.10](#) reports summary statistics for all variables used in the empirical analysis, and [Table 1.11](#) 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. Therefore, all regression specifications include up to two lags of macroeconomic controls as defined in [Section 1.4.2](#).

1.4 EMPIRICAL STRATEGY

1.4.1 Horizon-dependent Directions of OLS Bias

The omitted variable bias is a central challenge in using OLS regressions to establish causal relationships, and understanding the direction of OLS bias can provide insight into the underlying mechanisms. In a macroeconomic setting with dynamic responses across horizons, however, the direction of omitted variable bias may itself be *horizon-dependent*, rather than characterized by a single *static* cross-sectional bias term.

Treat output loss magnitude as positive such that larger values correspond to more severe recessions and consider two thought experiments. First, in the short run, sovereign default may alleviate debt overhang, facilitate debt restructuring under more favorable terms, and reduce the need for immediate fiscal austerity. These benefits are positively associated with the decision to default yet negatively related to output loss magnitude. Consequently, omitting such short-run debt-relief channels would induce a downward bias in OLS estimates, causing OLS to underestimate the true causal effect of sovereign default on output losses.

Second, countries experiencing severe domestic distress—such as financial crises, wars, or institutional breakdowns—are both more likely to default on external obligations and more likely to suffer larger output losses.⁸ Omitting these domestic factors would induce an upward bias in OLS estimates because they are positively associated with both default probability and output loss magnitude. Importantly, unlike the short-run debt-relief channels, these macroeconomic disruptions tend to be highly persistent. Their effects may continue to dominate at longer horizons, thereby causing OLS estimates to overstate the long-run costs of sovereign default.

Since both mechanisms plausibly operate simultaneously, the net direction of bias becomes an empirical question. More importantly, the dominant source of omitted variable bias may differ across horizons, implying that OLS estimates may underestimate the causal effects in the short run while overestimating them in the long run.

1.4.2 Baseline Controls and Common Specifications

The central comparison in this paper is between the local projection–ordinary least squares (LP-OLS) and the local projection–shift-share instrumental variable (LP-SSIV) estimates of the cumulative output costs of sovereign default. All LP-OLS and LP-SSIV regressions—including both the first-stage and 2SLS regressions—are estimated using (1) the same 3,898 observations (unless otherwise noted due to data availability), (2) country fixed effects, (3) a common set of baseline controls denoted by $\mathbf{X}_{it}$, and (4) standard errors clustered at the country level.⁹

For each of the 135 developing countries in the sample (indexed by i), the baseline control set $\mathbf{X}_{it}$ includes up to two lags of the following control variables: (i) regression-specific controls, including the treatment variable (e.g., the first-year default indicator or arrears-

⁸Rodrik (1999) argues that weak institutions, including social fragmentation and corruption, often interact with external shocks to generate growth collapses in developing countries. Such institutional fragility may therefore constitute an important source of omitted-variable bias when estimating the causal effect of sovereign default on output loss.

⁹See Section 1.5.4 for a discussion of shock-level clustering.

based partial default measures), one lead and one lag of the shift-share instrument to satisfy the lead-lag exogeneity (Stock and Watson, 2018), and the “incomplete share” defined as one minus the sum of the six major currency-denomination shares (Borusyak, Hull and Jaravel, 2022); (ii) continuous macroeconomic controls, including the first differences of log real GDP per capita, the log GDP deflator, the log nominal exchange rate, and log world GDP, as well as the levels of the debt-to-GDP ratio and the Chinn and Ito (2006) capital openness index; and (iii) binary indicators for banking crises, currency crises, political crises (defined as wars, coups, or political transitions), and democracy (Acemoglu et al., 2019).

1.4.3 Baseline LP-OLS Setup

I begin with the baseline 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 $h = 0, 1, \dots, 6$. 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.

Henceforth, the LP-OLS specification is given as:

$$y_{it+h} - y_{it-1} = \alpha_{ih} + \beta_h D_{it} + \gamma_h \mathbf{X}_{it} + v_{it+h} \quad h = 0, 1, \dots, 6 \quad (1.4)$$

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 .

1.4.4 Baseline LP-SSIV Setup

To estimate the causal effect of sovereign default on output loss, I use the LP-SSIV approach. To be specific, I first define the “interest rate exposure”, ΔIRE_{it} , for country i in year t as

follows:¹⁰

$$\Delta \text{IRE}_{it} \equiv \sum_k \text{currencyshare}_{it}^k \times \Delta \text{interestrate}_t^k \quad (1.5)$$

where $\text{currencyshare}_{it}^k$ represents the share of country i 's external debt denominated in advanced economy k 's currency in year t . $\Delta \text{interestrate}_t^k$ denotes the change in the long-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).¹¹

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

$$D_{it} = a_i + b\Delta \text{IRE}_{it-\ell} + g\mathbf{X}_{it} + \eta_{it} \quad (1.6)$$

Henceforth, the LP-SSIV specification is given as:

$$y_{it+h} - y_{it-1} = \alpha_{ih} + \beta_h \hat{D}_{it} + \gamma_h \mathbf{X}_{it} + v_{it+h} \quad h = 0, 1, \dots, 6 \quad (1.7)$$

1.5 BASELINE RESULTS

1.5.1 First-stage Results

Panel (a) of [Table 1.3](#) reports the first-stage relationship between the first-year default binary indicator D_{it} and lagged values of the “interest rate exposure” $\Delta \text{IRE}_{it-\ell}$. Columns (1)–(3) show that the first-stage coefficient estimates on ΔIRE_{it-1} , ΔIRE_{it-2} , and ΔIRE_{it-3} are all positive, consistent with the hypothesis that foreign monetary tightening raises sovereign default probability more for countries with greater exposure to debt denominated in the tightening country’s appreciated currency. However, only the coefficient on the second lag,

¹⁰Following [Autor, Dorn and Hanson \(2013\)](#), I use “exposure” to denote share-weighted interest rate shocks constructed with pre-determined currency shares.

¹¹[Table 1.17–Table 1.21](#) 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, as well as time-invariant country-specific shares such as the average currency shares of the first-available shares. See [Section 1.5.2](#) for detailed discussions.

ΔIRE_{it-2} , is statistically significant, suggesting that the effect of foreign monetary tightening on sovereign default risk materializes with a delay.

Column (4) further shows that the second lag ΔIRE_{it-2} is the strongest predictor of sovereign default among the three lags considered. Given its first-stage strength and to satisfy the lead-lag exogeneity requirement, I therefore use ΔIRE_{it-2} as the primary instrument for the first-year binary default indicator, while including both ΔIRE_{it-1} and ΔIRE_{it-3} as controls in all regressions. To interpret the magnitude, 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 default two years later.¹²

The delayed response implied by the first-stage estimates is also visible in historical sovereign default experience. For example, Mexico’s 1982 sovereign default occurred roughly two to three years after the onset of the Volcker tightening cycle, suggesting that foreign monetary tightening may generate prolonged financial stress that accumulates gradually before ultimately resulting in default.

To examine this delayed timing pattern more systematically, [Figure 1.10](#) plots sovereign default waves alongside major U.S. monetary tightening episodes identified using the narrative monetary shock classifications of [Romer and Romer \(2023\)](#). This figure shows that sovereign default waves frequently follow major U.S. tightening episodes and tend to emerge with a lag. Complementing the figure, [Table 1.12](#) reports the share of total sovereign default episodes

¹²Since 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 provide a more natural 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 1.14](#) reports logit estimates using the same baseline controls and country fixed effects as in Panel (a) of [Table 1.3](#) to complement the 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 $\exp(0.815) - 1 = 126\%$ higher odds-ratios of sovereign default. Alternatively, when ΔIRE_{it-2} increases from 0 to 1, the default probability increases by $\frac{\exp(-3.042+0.815)}{1+\exp(-3.042+0.815)} - \frac{\exp(-3.042)}{1+\exp(-3.042)} = 5.18\text{pt}$ (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 the 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.

(101 defaults) occurring within various horizons following a U.S. monetary tightening episode over the 1970–2010 sample period (41 years). The result shows a substantial concentration of sovereign defaults episodes following U.S. monetary tightenings. While only 12.2% of years coincide with a narrative U.S. tightening episode, these years account for 15.8% of all sovereign defaults. Expanding the window to two years after a tightening episode increases year coverage to 31.7%, but these periods account for 52.5% of all defaults in the sample. Column (5) further shows that the largest increase in excess default concentration occurs approximately two years after the initial U.S. monetary tightening shock. Nevertheless, since narrative tightening episodes are relatively infrequent, the evidence in [Figure 1.10](#) and [Table 1.12](#) should be interpreted as descriptive evidence on delayed response rather than as a causal estimate of the transmission mechanism.

[Table 1.15](#) and [Table 1.16](#) also provide additional empirical evidence on mechanisms that may rationalize this delayed transmission. [Table 1.15](#) shows that countries with larger shares of variable-rate debt (e.g., LIBOR- or U.S. prime-linked sovereign debt) are more sensitive to interest rate exposure shocks. The coefficient estimates on the interaction terms $\Delta \text{IRE}_{it-\ell} \times \text{sharevarrate}_{it-1}$ are generally positive and statistically significant, indicating that holding $\Delta \text{IRE}_{it-\ell}$ constant, a larger variable-rate share amplifies the effect of foreign monetary tightening on default risk. In contrast, [Table 1.16](#) shows that countries with longer average debt maturities are less vulnerable to these shocks $\Delta \text{IRE}_{it-\ell}$, as shown by negative coefficient estimates on the interaction term $\Delta \text{IRE}_{it-\ell} \times \text{Maturity}_{it-1}$, implying that conditional on $\Delta \text{IRE}_{it-\ell}$, longer maturities attenuate the effect of foreign interest rate shock on sovereign default risk. Taken together, these results are consistent with a delayed transmission mechanism: variable-rate liabilities are repriced relatively quickly following foreign monetary tightening, whereas longer debt maturities slow the pass-through of higher borrowing costs into sovereign repayment stress.

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

Columns (1) and (2) in [Table 1.3](#) 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—decline in real GDP per capita—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, the LP-OLS results alone cannot establish the direction of causality due to potential omitted variable bias.

Column (2) of [Table 1.3](#) presents the baseline LP-SSIV results; Panel (a) of [Figure 1.1](#) visualizes the LP-SSIV estimates alongside their LP-OLS counterparts. The LP-SSIV causal estimate for year 0 is -8% but not statistically significant. It becomes more negative and reaches -18.5% by year 1, significant at the 5% level. It remains between -15% and -18% in year 2–3 before reverting to zero by year 6. Although the LP-SSIV estimates are generally larger in magnitude than the LP-OLS estimates, they are less persistent: the impulse response fades to zero within six years, and the joint significance test is not significant, suggesting that sovereign default may not necessarily entail a long-lasting negative impact. The difference between the LP-OLS and LP-SSIV estimates arises from the source of variations that each empirical design exploits: the LP-SSIV leverages default variations driven by exogenous foreign interest rate movements, whereas the LP-OLS mixes these with endogenous domestic shocks (e.g., recessions), complicating causal interpretation.

The multiple weak instrument tests in [Table 1.3](#) reinforce the validity of the shift-share instrument ([Anderson and Rubin, 1949](#); [Kleibergen and Paap, 2006](#); [Andrews, Stock and Sun, 2019](#)). 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.

1.5.3 Alternative Shift-share Instrument Specifications

In the Online Appendix, I examine the robustness of the baseline results to alternative shift-share instrument specifications. First, [Table 1.17](#) follows [Autor, Dorn and Hanson \(2013\)](#) and constructs the shift-share instrument using lagged currency shares. Second, [Table 1.18](#) replaces long-term interest rates with short-term policy rates from the JST Macrohistory Database. Third, [Table 1.19](#) presents a leave-one-out specification that excludes U.S.-dollar-denominated debt. A related concern is that the baseline shift-share instrument relies on time-varying currency shares, whereas many other shift-share applications use predetermined or time-invariant shares. To address this issue, [Table 1.20](#) constructs the instrument using country-specific average currency shares over the 1970–2010 sample period. On the other hand, [Table 1.21](#) uses the first available currency shares for each country (see [Table 1.9](#) for details). Last but not least, [Table 1.22](#) examines whether the results are driven by the 1980s Latin American Debt Crisis.

Across specifications, the estimated dynamic causal responses closely resemble the baseline results. There are two exceptions. First, the leave-one-out specification excluding U.S.-dollar-denominated debt yields a weak first stage (Kleibergen-Paap F-statistic is 3.91), although the estimated effects remain stable over the first two horizons (see [Table 1.19](#)). Second, while excluding observations associated with the 1980s Latin American Debt Crisis remain broadly unchanged, column (4) of [Table 1.22](#), which excludes all observations during the 1982–1989 period, produces a weak first stage. Apart from these cases, the estimated dynamic effects remain highly consistent with the baseline results in terms of their timing, persistence, and magnitude.

1.5.4 Robustness to Shift-share Identification Assumptions

Recent theoretical contributions to the shift-share literature emphasize that identification may arise from either exogenous shares or exogenous shifts. In this paper, I interpret the shift-share instrument primarily as an exogenous-shift design: conditional on predetermined currency-denomination shares, foreign interest rate changes provide plausibly quasi-random monetary shocks to sovereign borrowers. Building on this exogenous-shift framework, [Table 1.4](#) examines the robustness of the baseline LP-SSIV estimates using recent advances in the shift-share literature.

First, the major currencies used to construct the baseline shift-share instrument may not fully capture the evolving trends of currency composition of sovereign external debt. Over time, developing countries may issue local-currency debt, borrow in other emerging currencies (e.g., Renminbi), or restructure existing debt obligations. In addition, institutional developments such as the introduction of the euro mechanically altered the classification of currency-denominated debt by aggregating multiple Eurozone currencies into a single denomination. Therefore, the incomplete share control may capture economically meaningful variation in sovereign borrowing patterns omitted from the baseline specification.

To address this concern, column (2) follows [Borusyak, Hull and Jaravel \(2022, 2025\)](#) and includes year fixed effects interacted with the incomplete share, defined as one minus the sum of the six major currency shares. The estimated effects remain economically large and qualitatively similar to the baseline results. The first-year output loss increases from -8.0% in the baseline specification to -11.6% , while cumulative output losses reach approximately -22% at $h = 1$ and peak at nearly -30% by $h = 3$. Relative to the baseline estimates, however, the dynamic response becomes more persistent, remaining close to -20% at longer horizons rather than gradually returning toward zero. Although the Kleibergen-Paap statistic falls from 21.1 to 10.5, the instrument remains above conventional weak-instrument thresholds, and the qualitative conclusions remain unchanged.

One possible explanation is that debt restructurings and negotiations may alter the effective currency composition of outstanding obligations. During restructuring episodes, debt may be partially forgiven or renegotiated under different contractual terms. At the same time, debt-relief packages provided by international organizations may allow the borrowing countries to choose the currency denomination but are often recorded under broader categories such as “multiple currencies”, which will be included in the incomplete share in the baseline setting. Since debt-relief initiatives frequently occur in clusters (e.g., following the 1980s Latin American Debt Crisis), the incomplete share may capture the mitigating effects of sovereign distress after debt restructuring. If this channel matters but is omitted from the baseline specification, the LP-SSIV estimates may partially attenuate the true output cost of sovereign default.¹³

Columns (3)–(5) of [Table 1.4](#) implement the re-centered instrumental variable (re-centered IV) approach proposed by [Borusyak and Hull \(2023\)](#). The motivation is that neither the currency shares nor the observed interest rate changes are necessarily as-good-as-random. For example, developing countries may systematically select into borrowing in particular currencies because of geographic proximity or historical colonial ties. At the same time, monetary policies across advanced economies are often influenced by common global factors and exhibit substantial co-movement. As a result, both the exposure shares and the monetary shifts may contain components correlated with omitted determinants of sovereign default outcomes.

Following [Borusyak and Hull \(2023\)](#), I construct simulated shift-share instruments that preserve the observed currency shares while replacing realized interest rate changes with randomly reshuffled counterparts at the pooled currency-year, currency, and year levels. To be specific, for each permutation scheme, I generate 100 random permutations of interest

¹³Future work further examines the robustness of the results to alternative aggregation methods for Eurozone currencies during the transition from 1970–2000 to 2001–2010. Nevertheless, local-currency or Renminbi-denominated debt remain relatively limited over the 1970–2010 sample period and are therefore less likely to explain the observed differences.

rate changes and construct the corresponding simulated shift-share instruments:

$$\begin{aligned}\Delta\text{IRE}_{it}^{\text{Pooled},(d)} &= \sum_k \text{currencyshare}_{it}^k \times \text{Perm}(\Delta\text{interestrate}_t^k|(d)) \\ \Delta\text{IRE}_{it}^{\text{Currency},(d)} &= \sum_k \text{currencyshare}_{it}^k \times \text{Perm}(\Delta\text{interestrate}_t^k|k = K, (d)) \\ \Delta\text{IRE}_{it}^{\text{Year},(d)} &= \sum_k \text{currencyshare}_{it}^k \times \text{Perm}(\Delta\text{interestrate}_t^k|t = T, (d))\end{aligned}$$

where $(d = 1, \dots, 100)$ indexes the permutation draw. The operator $\text{Perm}(\Delta\text{interestrate}_t^k, (d))$ denotes a random permutation of interest rate changes across all currency-year observations, $\text{Perm}(\Delta\text{interestrate}_t^k|k = K, (d))$ denotes a permutation across years within a given advanced economy K , and $\text{Perm}(\Delta\text{interestrate}_t^j|t = T, (d))$ denotes a permutation across advanced economies within a given year T . In all cases, the observed currency shares remain unchanged. For each permutation scheme, I use the average of the 100 simulated shift-share instruments as an additional control variable.

Column (3) includes the pooled and currency-level re-centered IV controls, which account for the average monetary policy characteristics of advanced economies over the 1970–2010 period. The resulting estimates remain remarkably similar to the baseline specification. The first-year output loss increases modestly in magnitude from -8% to -9.1% , while the peak cumulative output loss reaches approximately -22.6% at $h = 3$, compared with -18.5% in the baseline specification. Although the recovery appears somewhat slower, the overall dynamic pattern remains highly comparable. Moreover, the Kleibergen-Paap statistic increases slightly from 21.1 to 21.8, indicating that instrument strength is unaffected.

The picture changes when year-level re-centered IV controls are introduced. Column (4) includes the year-level re-centered IV, while Column (5) includes all simulated instruments simultaneously. Although the first two horizons remain broadly comparable to the baseline specification, the coefficient estimates become substantially larger at longer horizons and the first stage weakens considerably. Including the year-level re-centered IV reduces the

Kleibergen-Paap F-statistic to approximately 10, while including all simulated shift-share instruments lowers it further to 7.67, below the conventional threshold associated with weak-instrument concerns.

This pattern is not unexpected. Unlike the pooled and currency-level permutations, the year-level permutation absorbs a substantial portion of the aggregate time-series variation underlying the shift component of the instrument. Since identification in the baseline LP-SSIV specification relies heavily on variation in foreign interest rate movements over time, controlling for year-level simulated instruments removes a large fraction of the variation used to identify the causal effect. Therefore, the resulting estimates become unstable and should be interpreted with caution.

More broadly, these results highlight the importance of global interest rate cycles for identification. As an additional robustness exercise (not reported in the main text), I include year fixed effects directly in the baseline LP-SSIV specification. Doing so similarly weakens the first stage and renders most coefficient estimates statistically insignificant. One interpretation is that sovereign defaults become substantially more likely during episodes of global monetary tightening across advanced economies. By construction, year fixed effects absorb this common variation and therefore remove much of the identifying variation underlying the shift-share instrument.

Rather than undermining the baseline findings, the year-level re-centered IV exercise helps clarify the source of identifying variation in the baseline LP-SSIV design. It also motivates the complementary LP-DiD specifications presented in [Section 1.6.4](#), which explicitly include year fixed effects while relying on a different identifying assumption—parallel trends. Reassuringly, the LP-DiD estimates lead to similar qualitative conclusions regarding the output costs of sovereign default.

Last but not least, another concern in the shift-share literature relates to statistical inference. [Adão, Kolesár and Morales \(2019\)](#) and [Borusyak, Hull and Jaravel \(2022\)](#) emphasize that conventional standard errors may understate uncertainty because observations

with similar exposure shares are mechanically correlated through common shocks. I explored implementing AKM-style inference for the present setting (not reported in the main text). However, unlike Autor, Dorn and Hanson (2013), which features a comparable number of treatment units and industry-level shocks (both around 700), the shift-share instrument in this paper is constructed from fewer than ten currency-level shocks but more than 130 sovereign borrowers. Consequently, the asymptotic framework underlying AKM-style inference is less directly applicable in this small-shifter environment. In practice, shock-level standard error adjustments without country-level clustering produced substantially narrower confidence intervals. For this reason, all baseline specifications report country-clustered standard errors throughout.

1.5.5 Addressing Exclusion Restriction: Control Function Approach

1.5.5.1 Control Function Framework

To address potential violations of the exclusion restriction assumption, I follow the method outlined by Jordà, Schularick and Taylor (2020), who integrates the control function framework developed by Conley, Hansen and Rossi (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 1 *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:*

$$\widehat{\beta}^{IV} \xrightarrow{p} \beta + \frac{\phi}{b} \tag{1.8}$$

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, $\mathbb{E}[z^{ND} v^{ND}] = 0$. 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,

$$y - \lambda \hat{\phi}^{ND} z = \beta D + \left(\phi - \lambda \hat{\phi}^{ND} \right) z + v \quad (1.9)$$

will yield a consistent estimator for the true treatment effect β

$$\hat{\beta}_{CF}^{IV}(\lambda) = \frac{\frac{1}{N} \sum_{i=1}^N \left((y_i - \lambda \hat{\phi}^{ND} z_i) z_i \right)}{\frac{1}{N} \sum_{i=1}^N D_i z_i} \xrightarrow{p} \beta \quad (1.10)$$

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

Proof. The first part of the proof is detailed in [Jordà, Schularick and Taylor \(2020\)](#). See [Section 1.16](#) for the second part of the proof. ■

1.5.5.2 Estimating and Interpreting Relative Spillover Ratios

To empirically implement the above idea, denote $\text{Defaulters}_i = 1$ as a binary indicator equal to one if country i has *ever* experienced a sovereign default episode (i.e., it is time-invariant), and consider the following OLS regression:

$$y_{it+h} - y_{it-1} = \alpha_{ih} + \phi \Delta \text{IRE}_{it-\ell} + \psi (\Delta \text{IRE}_{it-\ell} \times \text{Defaulters}_i) + \epsilon_{it+h}$$

In the above specification, ϕ captures the spillover effects of $\Delta \text{IRE}_{it-\ell}$ on *never-defaulting* countries, while ψ captures the *additional* spillover effects among defaulting countries. Therefore, the total spillover effects for defaulting countries are given by $\phi + \psi$. Using these estimates, define a relative spillover ratio λ (defaulting relative to never-defaulting

countries) as

$$\lambda = \frac{\phi + \psi}{\phi} \quad (1.11)$$

In the baseline specification of Table 1.3, the identifying assumption effectively imposes $\lambda = 0$, implying that $\Delta \text{IRE}_{it-\ell}$ has *no* spillover effects on defaulting countries and affects output losses entirely through the sovereign default channel. Table 1.5 relaxes this assumption and estimates the relative spillover parameter directly.

First, consider horizon $h = 0$. The estimated relative spillover ratio equals $\lambda = \frac{-0.35+0.20}{-0.35} = 0.43 < 1$, implying that defaulting countries experience smaller spillover effects than never-defaulting countries. At horizon $h = 1$, the estimated relative spillover ratio becomes $\lambda = \frac{-0.36+(-0.15)}{-0.36} = 1.42 > 1$, suggesting that defaulting countries experience larger spillover effects than never-defaulting countries.

Beginning from $h = 2$, however, the directions of spillover effects diverge between defaulting and never-defaulting countries. For example, at $h = 2$, the spillover effect experienced by never-defaulting countries equals $0.28 > 0$, whereas that for defaulting countries equals $0.28 + (-0.95) = -0.67 < 0$. As a result, the relative spillover ratio becomes negative, and similar patterns persist at later horizons.

What do these results imply, and what values of the relative spillover ratio should be considered reasonable? The estimates suggest that differential exposure to foreign monetary shocks generates economically meaningful spillover effects. In particular, later horizons show defaulting and never-defaulting countries experience opposite spillover responses: spillover effects for never-defaulting countries gradually shift from negative to positive, suggesting that although contractionary foreign monetary shocks are initially recessionary, they may eventually become expansionary through trade channels—currency depreciation may boost exports and partially cushion these economies from external shocks.

In contrast, defaulting countries appear to be more vulnerable to differential exposure to foreign monetary shocks, with spillover effects remaining negative throughout later horizons. Since these negative effects dominate at longer horizons, I view the estimate $\lambda = 1.42$ as

a more economically relevant benchmark, while treating the estimate at $\lambda = 0.43$ as a conservative lower-bound specification for robustness check. Henceforth, I consider rounded calibration values of $\lambda = 0.5$ and $\lambda = 1.5$, which approximate these estimated values while avoiding undue emphasis on exact point estimates. The negative spillover ratios at later horizons are more difficult to interpret structurally and therefore are not used as benchmark parameter values.

1.5.5.3 Spillover-Corrected Cost of Sovereign Default

Panel (b) of [Figure 1.1](#) presents sovereign default cost estimates under alternative assumptions regarding the relative spillover parameter λ . The baseline case of $\lambda = 0$ —that is, no spillover effects on defaulting countries—collapses back to the benchmark LP-SSIV estimates. The green shaded region represents the range of spillover-corrected estimates under empirically plausible values of $\lambda \in [0.5, 1.5]$.

Although the spillover-corrected estimates broadly resemble the baseline LP-SSIV results, they imply substantially more persistent output losses. Under the conservative specification of $\lambda = 0.5$, where defaulting countries experience smaller spillovers than never-defaulting countries, output losses remain around -10% by year six. In contrast, under $\lambda = 1.5$, where defaulting countries experience larger spillovers, output losses approach -35% by year six.

This persistence is economically intuitive. Foreign monetary shocks may generate positive spillover effects through trade channels: higher foreign interest rates depreciate local currencies, boost exports, and stimulate economic activity among non-defaulting economies. Consequently, even modest positive spillovers imply that the underlying costs of sovereign default may be more severe and persistent than suggested by the baseline estimates.

1.5.5.4 Comparable Country Set

A potential concern with the baseline control-function approach in [Section 1.5.5.3](#) is that defaulting and never-defaulting countries may not constitute an appropriate comparison group. In particular, defaulting countries may differ systematically from never-defaulting countries in terms of indebtedness, policy credibility, or institutional quality. If so, the estimated spillover ratios may partly capture differences in underlying country characteristics rather than group-specific instrument spillovers.

To address this concern, I first restrict the analysis to a comparable-country sample constructed using propensity-score matching based on average *country-level* characteristics over the 1970–2010 sample period. The matching variables include average real GDP per capita growth, inflation, exchange rate changes, debt-to-GDP ratios, capital-account openness, banking, currency, and political crises, democracy indicators, and incomplete currency shares. The resulting sample contains 92 countries—55 defaulting countries and 37 never-defaulting countries. Within this sample, each defaulting country can be matched to at least one never-defaulting counterpart with similar observable characteristics, vice versa. I then re-estimate the spillover parameters and construct the corresponding spillover-corrected LP-SSIV range using the matched sample.

[Figure 1.2](#) reports the results. Panel (a) reproduces the baseline spillover-corrected estimates from the full sample, corresponding to Panel (b) of [Figure 1.1](#). Panel (b) presents the range of the spillover-corrected LP-SSIV estimates for the comparable-country sample. Relative to the baseline, the estimated spillover ratios decline substantially, with the identified range narrowing from $\lambda \in [0.5, 1.5]$ in the full sample to $\lambda \in [0.36, 0.81]$ in the matched-country sample. While the lower bound (0.36) remains close to the baseline estimate (0.43), the upper bound (0.81) declines markedly from its counterpart (1.42), suggesting that the defaulting countries experience relatively smaller spillovers from the shift-share instrument once the comparison is restricted to observably similar countries. As a result, the range of the spillover-corrected LP-SSIV estimates becomes considerably tighter while continuing to imply

sizeable output losses following sovereign default. The estimated losses remain economically large, reaching approximately -20% at horizons $h = 2$ and $h = 3$ before returning to 0 at $h = 4$. Although the identified range somewhat diverges at longer horizons ($h = 5$ and $h = 6$), this likely reflects the reduced sample size and the lower precision of long-horizon estimates. Importantly, the magnitude, timing, and sign of the estimated output losses during the first four years following default remain largely unchanged relative to the baseline specification.

A related concern is that comparability should be assessed using pre-default characteristics measured at the *country-year-level* rather than at the country level. Accordingly, Panel (c) restricts the analysis to a propensity-score-matched country-year sample using the same set of observable characteristics measured annually. This procedure retains 3,786 of the 3,898 observations in the baseline sample (97.1%). Unlike the country-level matching exercise, the estimated spillover ratios remain close to their baseline values, producing a spillover-corrected range that closely resembles that in Panel (a), although the identified range becomes somewhat wider at longer horizons.

Taken together, the overall pattern of output losses remains remarkably stable across all specifications. These results suggest that the spillover-corrected LP-SSIV estimates are not driven by comparisons between fundamentally different defaulting and never-defaulting countries. Restricting the analysis to observably comparable countries yields qualitatively similar estimates and, if anything, produces a tighter identified range. Moreover, matching at the country-year level leaves the baseline conclusions largely unchanged. Collectively, these findings strengthen the interpretation that the estimated spillover-corrected output losses reflect the causal effects of sovereign default rather than systematic differences in underlying country characteristics.

1.5.6 Residualized Interest Rate Changes

One concern with the baseline specification in [Table 1.3](#) is that observed interest rate changes in advanced economies may not fully represent exogenous monetary shocks—central banks

adjust monetary policy in response to domestic economic conditions, implying that observed interest rate changes partly reflect endogenous responses to home-country macroeconomic conditions, which may indirectly affect developing countries through channels beyond monetary policy.¹⁴ In addition, interest rates across advanced economies often move together over time and may coincide with periods of broader global financial conditions, such as the disinflationary episodes of the early 1980s or the synchronized monetary easing following the 2008 Global Financial Crisis. Consequently, the baseline shift-share instrument based on observed interest rate changes may partly capture broader global macroeconomic conditions rather than purely unexpected monetary shocks.

To address these concerns, I follow [Jordà, Singh and Taylor \(2024\)](#) to construct residualized interest rate changes that isolate the unexplained components by advanced economies’ domestic conditions. In addition, I include log changes in the CRB global commodity price index—a common approach for addressing the “price puzzle” in the monetary literature—to account for common global time-varying shocks. This adjustment is particularly relevant because global commodity cycles were closely linked to major macroeconomic events during the 1970s and early 1980s: windfall oil revenues, global inflationary episodes, and subsequent monetary tightening in advanced economies coincided with the Latin American debt crisis, the largest cluster of sovereign default episodes in the sample. Controlling commodity price movements therefore helps separate unexplained monetary shocks from broader global macroeconomic conditions that simultaneously influence both advanced economies’ monetary policy decisions and developing countries’ economic performance.¹⁵

¹⁴For example, a strong U.S. economy may increase import demand and create more export opportunities for developing countries.

¹⁵One additional concern relates to sovereign default contagion. Although sovereign defaults were highly clustered during the 1980s Latin American debt crisis, these episodes may not necessarily reflect contagion in the traditional sense, where financial distress in one country directly triggers crises in another country through investor panic or capital flight. Instead, many of these defaulting countries had accumulated substantial foreign-currency debt and were simultaneously exposed to common shocks of global monetary tightening following the Volcker disinflation. In contrast, later sovereign default episodes appear considerably more localized. For example, Argentina’s 2001 default did not trigger widespread defaults among neighboring economies such as Brazil.

To be specific, consider the following regressions:

$$\Delta \text{interestrate}_t^k = \alpha^k + \beta \mathbf{X}_{kt} + \gamma \Delta \ln(\text{CommodityPrice}_t) + \xi_{kt}^{\text{Resid}} \quad (1.12)$$

where α^k denotes advanced economies' country fixed effects, and $\mathbf{X}_{kt}$ is a vector of control variables in advanced economy k , including up to two lags of changes in log real GDP per capita, real consumption per capita, real investment per capita, CPI, short-term interest rates, long-term interest rates, and the capital-account-to-output ratio.

Using $\{\xi_{kt}^{\text{Resid}}\}$ as the monetary shock series, I reconstruct the residualized “interest rate exposure” measure, $\Delta \text{IRE}_{it}^{\text{Resid}}$, as follows:

$$\Delta \text{IRE}_{it}^{\text{Resid}} \equiv \sum_k \text{currencyshare}_{it}^k \times \xi_{kt}^{\text{Resid}}$$

The only difference from Equation (1.5) is replacing observed interest rate changes, $\Delta \text{interestrate}_t^k$, with their residualized counterpart ξ_{kt}^{Resid} . The first-stage and second-stage specifications, analogous to Equation (1.6) and Equation (1.7), are then given as:

$$\begin{aligned} D_{it} &= a_i + b \Delta \text{IRE}_{it-\ell}^{\text{Resid}} + g \mathbf{X}_{it} + \eta_{it} \\ y_{it+h} - y_{it-1} &= \alpha_{ih} + \beta_h \hat{D}_{it} + \gamma_h \mathbf{X}_{it} + v_{it+h} \quad h = 0, 1, \dots, 6 \end{aligned}$$

Table 1.6 presents the residualized estimates on the cost of sovereign defaults and compares them with the baseline results in Table 1.3. Panel (a) reports the first-stage estimates. Consistent with the baseline specification, the second lag of the residualized exposure measure, $\Delta \text{IRE}_{it-2}^{\text{Resid}}$, remains the strongest predictor of sovereign default. Moreover, both the first-stage coefficient and the Kleibergen–Paap statistic remain comparable to the baseline estimates, suggesting that residualization does not materially weaken instrument relevance.

Panel (b) reports the second-stage estimates. Relative to the baseline specification, the initial output decline becomes smaller in magnitude: real GDP per capita falls by approximately 5.5% on impact compared with roughly 8% in the baseline specification. The cumulative output loss at horizon $h = 1$ reaches approximately -16.5% and becomes statistically significant, although smaller in magnitude than the baseline estimate of -18.5% . One interpretation is that the baseline specification partly captures broader global disturbances—such as synchronized global monetary tightening or subsequent global slowdowns—that amplify the measured output loss of sovereign default. Residualization may therefore remove parts of interest rate variation associated with these broader macroeconomic episodes.

At longer horizons, however, the residualized estimates display a noticeably different pattern. While the baseline estimates in [Table 1.3](#) gradually converge toward zero within the six-year horizon, the residualized estimates remain substantially negative even at later horizons, though not statistically significant. In particular, estimated output losses remain close to 20% through horizon six, and the coefficient equality tests indicate statistically significant differences between the baseline and residualized estimates at later horizons $h = 5$ and $h = 6$.

What explains this divergence remains less clear. Interestingly, the residualized estimates display patterns similar to the spillover-corrected LP-SSIV estimates in [Section 1.5.5](#). One possible interpretation is that residualization partially removes common global time trends embedded within observed interest rate co-movements across advanced economies periods of widespread economic distress in developing countries. During episodes of global monetary tightening, developing economies often jointly experience broad and simultaneous macroeconomic adjustments, including exchange rate depreciations, sudden stops, and synchronized global slowdowns. To the extent that these common responses are correlated with observed interest rate changes, residualization may purge portions of these shared adjustment mechanisms and alter the underlying identifying variation, thereby generating more

persistent dynamic responses.¹⁶

This interpretation should be viewed cautiously. The residualized specification does not necessarily provide a “better” estimate than the baseline model. Rather, I interpret the baseline estimates as policy-relevant *equilibrium* estimates that allow endogenous adjustment mechanisms. In contrast, I view the spillover-corrected LP-SSIV and residualized specifications as approximating a more *structural* estimate on the cost of sovereign default by further purging confounding channels embedded in broader global macroeconomic conditions. Reassuringly, regardless of whether persistence differs across specifications, the central conclusion remains unchanged: sovereign defaults continue to generate economically large output losses in the short run.

1.6 ALTERNATIVE MEASURES AND EMPIRICAL APPROACHES

This section evaluates the robustness of the baseline LP-SSIV estimates in [Section 1.5](#) along two dimensions. First, it considers alternative measures of sovereign distress and foreign monetary shocks. To be specific, [Section 1.6.1](#) examines whether the main findings are sensitive to other binary sovereign default classifications proposed in the literature. [Section 1.6.2](#) replaces observed interest rate changes with narrative monetary shock series as the source of exogenous shifts. [Section 1.6.3](#) uses continuous arrears-based partial default measures as alternative treatment variables.

Second, [Section 1.6.4](#) applies modern difference-in-differences (DiD) estimators, which rely on a different identification strategy based on the parallel trends assumption. An advantage of the DiD approach is that it can explicitly include year fixed effects, helping address concerns that common global time trends drive both foreign interest rates and

¹⁶To further address concerns regarding common global trends, [Section 1.6.4](#) explicitly introduces year fixed effects within the difference-in-differences framework.

sovereign default decisions. In contrast, the LP-SSIV framework is better suited for tracing heterogeneous effects across policy regimes, as explored in [Section 1.8](#), but cannot absorb global time trends as directly as DiD estimators. The consistency of the findings across these two empirical strategies with complementary strengths further strengthens the credibility of the baseline results.

1.6.1 Different Binary Default Classifications

Panel (c) in [Figure 1.1](#) 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 \(2001\)](#). 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.

[Table 1.23](#) compares the baseline LP-SSIV estimates based on the S&P first-year default indicator with alternative sovereign distress classifications from [Laeven and Valencia \(2020\)](#) in detail. In addition to sovereign debt crises, they separately identify sovereign debt restructuring episodes, allowing the construction of a broader debt distress indicator that combines both events. Columns (1)–(4) show that the debt crisis classification of [Laeven and Valencia \(2020\)](#) yields estimates that are quantitatively similar to the baseline results, although the implied output losses are somewhat larger in magnitude. Likewise, expanding the treatment definition to include debt restructuring episodes produces slightly larger losses but leaves the overall dynamic pattern largely unchanged. Across all classifications, output declines sharply following the onset of sovereign distress, reaches its trough after approximately two to three years later, and gradually returns toward zero thereafter. The

corresponding Anderson-Rubin p-values and Kleibergen-Paap statistics remain comparable across specifications. Overall, these findings suggest that the baseline results are robust to broader definitions of sovereign debt distress.¹⁷

1.6.2 Narrative Monetary Shocks

The baseline LP-SSIV specification in [Section 1.5.2](#) uses observed interest rate changes in advanced economies to capture foreign monetary tightening. A potential concern, however, is that realized interest rate changes may reflect broader macroeconomic conditions abroad rather than purely exogenous monetary policy surprises. To address this concern, I construct alternative shift-share instruments using the narrative monetary shock series developed by [Romer and Romer \(2004\)](#) for the Federal Reserve and [Cloyne and Hurtgen \(2016\)](#) for the Bank of England. To be specific, I define

$$\text{IRE}_{it}^{\text{USRR}} \equiv \text{currencyshare}_{it}^{\text{USD\$}} \times \text{USRRshock}_t$$

$$\text{IRE}_{it}^{\text{UKCH}} \equiv \text{currencyshare}_{it}^{\text{GBR£}} \times \text{UKCHshock}_t$$

Since the original narrative monetary shock series are available at the monthly frequency, they must be aggregated to annual observations. [Figure 1.9](#) shows that the U.K. monetary shock series is substantially more volatile than its U.S. counterpart, especially during the 1975–1985 period, and exhibits a more pronounced decline in volatility over time. Aggregating monthly shocks using annual averages or sums may therefore attenuate monetary policy surprises by offsetting opposite-sign monthly surprises within a given year. To preserve the most economically meaningful policy surprise, I define the annual monetary shock as the largest absolute monthly shock within each year while retaining its sign. The resulting annual

¹⁷An alternative approach would be to use sovereign bond spreads to construct market-implied sovereign distress indicators. This paper does not pursue that approach because spread-based measures require a subjective choice of distress thresholds and are difficult to implement consistently across the historical developing country sample, especially given the limited availability of liquid sovereign bond markets and spread data during earlier decades.

series are then standardized separately for the United States and the United Kingdom over the full sample, yielding USRRshock_t and UKCHshock_t .

Using these measures, the first-stage specification becomes

$$D_{it} = \mathbf{a}_i + \mathbf{b} \cdot (\text{IRE}_{it}^{\text{USRR}}, \text{IRE}_{it}^{\text{UKCH}}) + g\mathbf{X}_{it} + \eta_{it}$$

Panel (a) of [Table 1.7](#) reports the corresponding first-stage estimates. Both narrative-shock shift-share instruments exhibit positive and statistically significant relationships with the first-year binary default indicator. This result is noteworthy for two reasons. First, unlike realized interest rate changes, narrative monetary shocks are designed to isolate unexpected monetary policy surprises and are therefore less likely to capture predictable macroeconomic developments abroad. Second, the positive and statistically significant first-stage coefficient estimate on $\text{IRE}_{it}^{\text{UKCH}}$ suggests that, in addition to the dominant role of U.S. monetary policy, monetary tightening originating from other major creditor countries can also affect sovereign default risk for those with greater exposure to debt denominated in the corresponding currency (i.e., British pounds).¹⁸

The 2SLS specification remains identical to [Equation \(1.7\)](#):

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

Panel (b) of [Table 1.7](#) presents the corresponding LP-SSIV estimates using narrative monetary shocks. Although the coefficient estimates are imprecisely estimated and generally fail conventional weak-instrument diagnostics, their magnitudes and dynamic patterns remain

¹⁸Although Panel (a) of [Table 1.7](#) shows a smaller first-stage coefficient estimate on $\text{IRE}_{it}^{\text{USRR}}$ than on $\text{IRE}_{it}^{\text{UKCH}}$ (0.024 versus 0.053), [Table 1.10](#) indicates that $\text{IRE}_{it}^{\text{USRR}}$ has a much higher standard deviation than $\text{IRE}_{it}^{\text{UKCH}}$ (0.437 versus 0.095). To make magnitudes comparable, I scale the first-stage coefficient estimates by the dispersion of the corresponding regressors: a one-standard-deviation increase in $\text{IRE}_{it}^{\text{USRR}}$ is associated with $0.024 \times 0.437 = 0.010$ increase in default probability, whereas a one standard deviation increase in $\text{IRE}_{it}^{\text{UKCH}}$ is associated with a $0.053 \times 0.095 = 0.00535$ increase in default probability. After standardization, the implied first-stage effect is larger for U.S. monetary shocks, consistent with the dominant role of U.S. monetary policy in international financial markets.

remarkably similar to the baseline results reported in [Table 1.3](#). To be specific, using both IRE_{it}^{USRR} and IRE_{it}^{UKCH} as joint instruments yields an estimated first-year output loss of -8.2% , nearly identical to the baseline counterpart (-8%). The cumulative output losses subsequently deepen to roughly -21% after two years before gradually dissipating within four years, closely mirroring both the peak magnitude and the overall shape of the baseline impulse response.

Taken together, these findings suggest that the baseline LP-SSIV results are not solely driven by predictable components of foreign monetary policy. Although the narrative shift-share instruments are substantially weaker and therefore generate less precise estimates, they nevertheless imply a similar dynamic response of output losses following a sovereign default. The consistency of the estimates across two conceptually distinct measures of foreign monetary tightening provides additional support for the baseline interpretation.

1.6.3 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. Denote $arrears_{it} \in [0, \infty]$ as the late or missed debt service payments of country i to the world aggregate measured in current U.S. dollars. As shown in [Figure 1.8](#), 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 $arrears_{it}$ as follows: $\Lambda_1(arrears_{it}) = \ln(arrears_{it})$, $\Lambda_2(arrears_{it}) = \ln(1 + arrears_{it})$, and $\Lambda_3(arrears_{it}, debt_{it}) = \ln\left(1 + 100 \times \frac{arrears_{it}}{debt_{it}}\right)$.¹⁹ The first two are common in the firm entry-exit literature to handle zeros and yield interpretable

¹⁹The analysis uses $\ln\left(1 + 100 \times \frac{arrears}{debt}\right)$ instead of $\ln\left(1 + \frac{arrears}{debt}\right)$ because for small arrears-to-debt ratios, $\ln\left(1 + \frac{arrears}{debt}\right) \approx \frac{arrears}{debt}$, so a one-unit increase represents a 100-percentage-point (100%pt) change in the arrears-to-debt ratio. However, [Figure 1.8](#) 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—i.e., $\ln\left(1 + 100 \times \frac{arrears}{debt}\right)$ —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](#)).

elasticities; the third follows [Chen and Roth \(2024\)](#)’s guidance for anchoring the interpretation at an economically meaningful variable. Using these variables as proxies for partial default, the baseline LP-OLS and LP-SSIV specifications remain unchanged (as in [Equation \(1.6\)](#)–[Equation \(1.7\)](#)):

$$\begin{aligned}\Lambda_{it} &= a_i + b\Delta\text{IRE}_{it-\ell} + g\mathbf{X}_{it} + \eta_{it} \\ y_{it+h} - y_{it-1} &= \alpha_{ih} + \beta_h \hat{\Lambda}_{it} + \gamma_h \mathbf{X}_{it} + v_{it+h} \quad h = 0, 1, \dots, 6\end{aligned}$$

[Table 1.8](#) 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 $\Delta \ln\left(1 + 100 \times \frac{\text{arrears}}{\text{debt}}\right)$ are all statistically significant with the instrument ΔIRE_{it-3} instead of ΔIRE_{it-2} as in [Table 1.3](#). Some signs differ because of transformation, and interpretation 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 of coefficient estimates from either $\Delta \ln(\text{arrears})$ or $\Delta \ln(1 + \text{arrears})$ should be treated with caution because these log transformations do not separate the contributions from the extensive and intensive margins. [Chen and Roth \(2024\)](#) recommend anchoring log changes at economically meaningful value. Accordingly, I use the arrears-to-debt ratio and anchor the change at the threshold when it increases from 0 to an infinitesimal positive value, which I interpret as the threshold point at which a country enters initial default. The first-stage coefficient estimate evaluated at this anchor $\Delta \ln\left(1 + 100 \times \frac{\text{arrears}}{\text{debt}}\right)$ is 0.021, meaning a one-unit increase in ΔIRE_{it-3} is

associated with a change of arrears-to-debt ratio from 0 to 0.00021 (i.e., 0.021 percentage points increase).

Panel (b) of [Table 1.8](#) 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 $\Delta \ln(1 + 100 \times \frac{\text{arrears}}{\text{debt}})$ or baseline binary-default results as shown in [Table 1.3](#). 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 arrears-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 1.3](#)—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.

1.6.4 Difference-in-Differences 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 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 (Roth et al., 2023; Callaway and Sant’Anna, 2021; de Chaisemartin and DHaultfoeuille, 2020; Sun and Abraham, 2021; Borusyak and Hull, 2024). 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 1.4, 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 1.1. Accordingly, I restrict the sample to countries that have not defaulted in the previous six years and apply the following LP-DiD specification:

$$y_{it+h} - y_{it-1} = \beta_h^{\text{LP-DiD}} \Delta D_{it}^{\text{in}} + \delta_t^h + \mathbf{X}_{it} + \epsilon_{it}^h \quad h = -6, \dots, 0, \dots, 12 \quad (1.13)$$

restricting the estimation sample to a “clean control set”:²⁰

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

In terms of specification, several changes are made from Equation (1.7): 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 D_{it}^{in} is the S&P in-default indicators, which equals to 1 for the entire period when a country remains in a default episode.

Panel (a) of Figure 1.3 shows LP-DiD estimates closely align with the spillover-corrected IV estimates in Panel (b) of Figure 1.1—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 1.3 further validate the robustness of the main results.

While the LP-DiD estimator is my preferred specification because of its transparency and theoretical 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.²¹ Panel (b) of Figure 1.3 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.

²⁰Intuitively, the “clean control set” restrict samples to (i) treated in period t , $D_{it}^{\text{in}} = 1, D_{it-1} = 0$ (ii) treated in both period t and $t - 1$, $D_{it}^{\text{in}} = 1, D_{it-1} = 1$, (iii) untreated in both period t and $t - 1$, $D_{it}^{\text{in}} = 0, D_{it-1} = 0$. In other words, it excludes the negative weights $D_{it}^{\text{in}} = 0, D_{it-1} = 1$ by design.

²¹Since 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 1.11 re-plots these DiD estimates from Figure 1.3 alongside the Callaway and Sant’Anna (2021) estimates for direct comparison.

1.7 DISCUSSION OF BASELINE RESULTS

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

Table 1.3 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 Panel (b) of Table 1.1) consistent with observed default episodes, and how large are they relative to other crises? Panel (a) of Table 1.1 shows the cumulative output loss ($y_{t+h} - y_{t-1}$) with its counterparts from representative sovereign default episodes. Note that entries in Panels (a), (c), and (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 1.8.

Panel (c) of Table 1.1 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 1.1](#)), 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 1.1](#))—consistent with the difference-in-differences estimates in [Figure 1.3](#)—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.

1.7.2 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 ([Aguiar and Gopinath, 2006](#); [Yue, 2010](#)), they align well with the approximately -10% output drops calibrated in

more recent endogenous default models (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 (2009) and De Paoli, Hoggarth and Saporta (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, Graf von Luckner and Reinhart (2024), not only because we address similar questions, but also because we reach comparable conclusions despite using different methodologies and samples. Farah-Yacoub, Graf von Luckner and Reinhart (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 1.1) or difference-in-differences (Figure 1.3) aligns well with the conclusion in Farah-Yacoub, Graf von Luckner and Reinhart (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.

1.8 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 $h = 0, 1, \dots, 6$, I run the following state-dependent LP-SSIV regression:

$$y_{it+h} - y_{it-1} = I_{it} \times [\alpha_{Aih} + \beta_{Ah}\hat{D}_{it} + \gamma_{Ah}\mathbf{X}_{it}] + (1 - I_{it}) \times [\alpha_{Bih} + \beta_{Bh}\hat{D}_{it} + \gamma_{Bh}\mathbf{X}_{it}] + v_{it+h} \quad (1.14)$$

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 1.1](#) 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.²² In contrast, defaults under a floating exchange rate regime generally yield insignificant coefficients. This

²²Following the classification scheme developed by [Ilzetzki, Reinhart and Rogoff \(2019\)](#), 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 having floating exchange rate regimes. Countries with code 6, which denotes dual exchange rate markets, are excluded from the sample.

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 1.1](#) 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 debt with international creditors.

Third, Panel (f) of [Figure 1.1](#) 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.

Last but not least, Panel (g) in [Figure 1.1](#) 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 1.1](#) 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.

1.9 CONCLUDING REMARKS

How much does it cost if a country defaults on its external debt? Creditors in sovereign debt markets face “limited commitment”—sovereign debt contracts are difficult to enforce, and concerns about moral hazard have therefore long been central to the sovereign debt literature. Yet sovereign defaults occur only occasionally in practice, raising a fundamental question: why do countries repay their debt in the first place? A leading explanation—especially in quantitative sovereign default models that seek to match both the observed moments and the high levels of debt typically accumulated prior to default—is that defaulting countries suffer direct output losses. Despite its central role in the literature, however, this is an ad hoc assumption, as credible causal evidence on the magnitude and persistence of default-induced recessions remains limited.

In this paper, I introduce a novel LP-SSIV approach to estimate the causal effect of sovereign default on output, exploiting variation in developing countries’ endogenous currency denominations of external debt together with advanced economies’ quasi-random monetary policies, proxied by both interest rate changes and narrative monetary shocks. I find that sovereign default reduces real GDP per capita by -8% in the first year. The cumulative output losses peak at -18.5% in the second year, remain substantial through the fourth year, and gradually dissipate by the sixth year. I interpret these baseline LP-SSIV estimates as capturing policy-relevant *equilibrium* effects, reflecting the fact that exchange

rate depreciation following default generates positive spillovers through improved export competitiveness, thereby helping defaulting countries recover from recession.

By contrast, the spillover-corrected LP-SSIV estimates based on control-function bounds, the residualized LP-SSIV specifications, and the difference-in-differences estimators yield similar short-run but considerably more persistent output losses, ranging from -10% to -30% even a decade after default. I interpret these estimates as approximating the underlying *structural* effects of sovereign default, as they remove the remaining mitigating influence of exchange rate spillovers while also accounting for common global business-cycle fluctuations.

Last but not least, this paper provides empirical support for the monetary trilemma (Obstfeld, Shambaugh and Taylor, 2005). Although developing countries may not be able to fully shield themselves from quasi-random monetary policy shocks originating from abroad, conventional policy choices—such as transitioning from pegged to a floating exchange rate regimes and avoiding excessive accumulation of external debt, especially short-term debt—can substantially mitigate the economic costs of sovereign default.

Table 1.1: Cumulative output losses after sovereign defaults and other financial crises.

Cumulative decline in real GDP per capita (%): $y_{t+h} - y_{t-1}$							
Horizons (h)	$h = 0$	$h = 1$	$h = 2$	$h = 3$	$h = 4$	$h = 5$	$h = 6$
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
(a) 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
(b) Main results: LP-SSIV causal estimates							
Baseline LP-OLS	-2.7	-3.8	-4.8	-4.8	-5.3	-5.2	-5.2
Baseline LP-SSIV	-8.0	-18.5	-15.8	-18.2	-9.5	-5.7	-0.5
Residualized LP-SSIV	-5.5	-16.5	-12.8	-14.1	-17.0	-23.1	-18.1
(c) Representative currency and 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
(d) Benchmark—2008 Global Financial Crisis							
USA2009	-3.5	-1.7	-0.9	0.7	1.8	3.3	5.3

Notes: This table compares the magnitude and persistence of cumulative output losses following sovereign defaults with those observed during major financial crises. Entries report cumulative changes in log real GDP per capita, measured as $y_{t+h} - y_{t-1}$, where y_{t+h} denotes log real GDP per capita h years after the event and y_{t-1} denotes its value in the year preceding the event. The real GDP per capita data series, retrieved from the *World Development Indicators*, are in constant prices and local currency units. Panel (a) reports observed output paths following four representative sovereign default episodes: Argentina’s 2001–2002 default (ARG2001), Russia’s 1998 default (RUS1998), Ecuador’s 1999 default (ECU1999), and Greece’s 2010 debt crisis (GRC2010), which did not constitute a formal sovereign default. Panel (b) reports the main results: the baseline LP-OLS and LP-SSIV results from Table 1.3 and the residualized LP-SSIV estimates from Table 1.6 (see Section 1.5 and Section 1.5.6 for details). Panel (c) reports output losses associated with two prominent currency and financial crises in developing countries: the Mexican Tequila Crisis of 1994–1995 (MEX1995) and the Asian Financial Crisis centered on Thailand in 1997–1998 (THA1998). Panel (d) reports the output path of the United States during the 2008–2009 Global Financial Crisis (USA2009), which serves as a benchmark for comparison.

Table 1.2: Summary statistics and diagnostics at the shock- and share-levels.

(a) Observation counts				
(1) Country	(2) Defaults	(3) Unique defaulters	(4) Currency	(5) Currency-year
135	101	61	7 [†]	286
(b) Share-level summary statistics				
	(1) Mean	(2) S.D.	(3) P10	(4) P90
U.S. dollar	0.465	0.225	0.314	0.770
Japanese yen	0.055	0.095	0.066	0.166
Pound sterling	0.043	0.114	0.029	0.113
Swiss franc	0.011	0.045	0.006	0.024
Deutsche mark (pre-2001)	0.053	0.079	0.074	0.133
French franc (pre-2001)	0.069	0.131	0.069	0.210
Euro*	0.160	0.194	0.161	0.452
Incomplete share**	0.293	0.191	0.248	0.560
(c) Share-level diagnostics				
	(1) Currency-year	(2) Year	(3) Currency	
Effective sample size (1/HHI of s_t^k shares)	79.76	39.30	2.20	
Largest s_t^k share	0.023	0.033	0.658	
(d) Shock-level summary statistics				
	(1) Mean	(2) S.D.	(3) IQR	(4) Persistence
$\Delta \text{interestrate}_t^k$	-0.100	0.975	1.072	0.218*** (0.027)
ξ_{kt}^{Resid}	-0.025	0.837	1.009	-0.018 (0.039)

Notes: This table presents summary statistics and diagnostics of the shock- and share-level data. Panel (a) reports sample counts. Panel (b) summarizes country-year-currency shares of long-term public and publicly guaranteed external debt for 135 developing countries over 1970–2010. “Euro*” aggregates legacy ERM currencies (e.g., Deutsche mark, French franc, Italian lira, etc.) after 2001, while “Incomplete share**” is defined as one minus the sum of major-currency shares. Panel (c) reports concentration diagnostics of currency shares, including the effective sample size (inverse Herfindahl index) and the largest observed currency share. To be specific, given observed debtor-year-currency shares, ω_{it}^k , I first aggregate across countries to obtain $\omega_t^k \equiv \sum_i \omega_{it}^k$. I then normalize these currency-year exposures, that is, $s_t^k \equiv \omega_t^k / \sum_{k,t} \omega_t^k$ such that $\sum_{k,t} s_t^k = 1$. The currency-year effective sample size is defined as the inverse Herfindahl (HHI) index, $1 / \sum_{k,t} (s_t^k)^2$. To compute the currency-level inverse HHI index, I further aggregate $s^k \equiv \sum_t s_t^k$ and calculate $1 / \sum_k (s^k)^2$. The year-level inverse HHI index is computed similarly. Panel (d) reports summary statistics of long-term interest rate changes $\Delta \text{InterestRate}_{kt}$ and their residualized counterpart ξ_{kt}^{Resid} constructed by residualizing advanced-economy interest rate changes as detailed in [Section 1.5.6](#). Column (4) reports AR(1) persistence coefficients on its first lag with country fixed effects; standard errors, clustered at country-level, are reported in parentheses. All statistics are unweighted. † Before 2001, 7 major currencies are observed; since 2001, the Deutsche mark and French franc are incorporated into the Euro category, resulting in 6 major currencies.

Table 1.3: Baseline LP-OLS and LP-SSIV estimates on default cost with ΔIRE_{it-2}
(a) First-stage between $\Delta\text{IRE}_{it-\ell}$ with S&P first-year default indicators

Dependent variable: S&P first-year binary (0/1) default indicator				
	(1)	(2)	(3)	(4)
ΔIRE_{it-1}	0.007 (0.006)			0.009 (0.006)
ΔIRE_{it-2}		0.027*** (0.006)		0.026*** (0.006)
ΔIRE_{it-3}			0.010* (0.006)	0.010* (0.006)

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

Notes: This table reports the first-stage estimates and the baseline LP-OLS and LP-SSIV estimates of sovereign default costs using the second lag of interest rate exposure, ΔIRE_{it-2} , as the instrument. It also reports p-values from the joint significance test across all horizons (i.e., $\beta_0 = \dots = \beta_6 = 0$), the Kleibergen-Paap weak instrument (F-statistic), the Anderson-Rubin weak instrument test (F-statistic), and Baseline SSIV = Residualized SSIV coefficient equality tests (stacking method) at each horizon h . All columns include (i) country fixed effects, (ii) the baseline control set $\mathbf{X}_{it}$ defined in Section 1.4.2, and (iii) the same 3,898 observations. Robust standard errors clustered at the country level are reported in parentheses. Statistical significance is indicated as *** p<0.01, ** p<0.05, * p<0.1.

Table 1.4: Robustness checks for the baseline LP-SSIV estimates.

Horizons (h)	LP-SSIV output response ($y_{t+h} - y_{t-1}$)				
	Baseline	Year-specific Incomplete shares	Recentered-IVs as controls		
	(1)	(2)	(3)	(4)	(5)
$h = 0$	-7.99 (5.28)	-11.57 (8.07)	-9.10* (5.36)	-13.40 (9.60)	-14.16 (9.98)
$h = 1$	-18.48** (8.08)	-22.09* (13.30)	-21.24** (8.38)	-21.67 (14.38)	-24.61 (15.70)
$h = 2$	-15.78* (8.60)	-18.74 (13.61)	-17.34* (8.97)	-31.71* (16.51)	-29.25 (18.55)
$h = 3$	-18.25* (9.52)	-29.25* (15.48)	-22.63** (10.48)	-58.14** (22.71)	-49.62** (24.07)
$h = 4$	-9.47 (10.72)	-16.69 (17.16)	-19.95* (11.82)	-79.61*** (26.35)	-72.09** (28.33)
$h = 5$	-5.70 (10.55)	-16.63 (17.40)	-17.34 (12.17)	-64.97*** (24.08)	-62.59** (26.73)
$h = 6$	-0.52 (11.17)	-21.90 (19.03)	-13.08 (13.09)	-36.08* (21.25)	-34.18 (23.09)
KP Weak IV Instrument Observations	21.14	10.47	21.82 ΔIRE_{it-2} 3,898	10.72	7.67
Incomplete shares × Year FEs	Y				
<u>Simulated IVs</u>					
Pooled	Y				Y
Currency	Y				Y
Year				Y	Y

Notes: This table presents multiple robustness checks for the baseline LP-SSIV causal estimates of the cost of sovereign default. Column (1) reproduces the baseline LP-SSIV estimates reported in Table 1.3. Column (2) presents the robustness check that additionally includes year-specific incomplete shares, following Borusyak, Hull and Jaravel (2022). Columns (3)–(5) present robustness checks including re-centered IVs as additional controls, following Borusyak and Hull (2023). Column (3) includes re-centered IVs constructed from pooled reshuffling of interest rate changes across all currency-year observations, as well as currency-level reshuffling that holds currency fixed while reshuffling interest rate changes over the 1970–2020 period. Column (4) includes re-centered IVs constructed from year-level reshuffling that holds year fixed while reshuffling interest rate changes across currencies. Column (5) includes all simulated IVs as controls. Columns (3)–(5) use interest rate changes across 19 advanced economies, rather than the six major currencies, over the 1970–2020 period from the JST Macrohistory database to better capture variation in interest rate movements. This table also reports the p-values of the Kleibergen-Paap weak instrument test (F-statistic) for each specification. All columns include (i) country fixed effects, (ii) the baseline control set $\mathbf{X}_{it}$ defined in Section 1.4.2, and (iii) the same 3,898 observations. Robust standard errors clustered at the country level are reported in parentheses. Statistical significance is indicated as *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table 1.5: Estimating spillover ratio for defaulting relative to never-defaulting countries

Dependent variable: Cumulative output loss ($y_{it+h} - y_{it-1}$)			
Independent variable: ΔIRE_{it-2}			
	Spillovers to never-defaulters	Additional spillovers to defaulters	Relative spillover ratio
Horizons (h)	ϕ	ψ	$\lambda \equiv \frac{\phi + \psi}{\phi}$
	(1)	(2)	(3)
$h = 0$	-0.35* (0.21)	0.20 (0.25)	0.43
$h = 1$	-0.36 (0.31)	-0.15 (0.41)	1.42
$h = 2$	0.28 (0.42)	-0.95* (0.52)	-2.39
$h = 3$	0.24 (0.49)	-0.98 (0.60)	-3.08
$h = 4$	0.15 (0.50)	-0.54 (0.59)	-2.60
$h = 5$	0.34 (0.47)	-0.68 (0.58)	-1.00
$h = 6$	0.71 (0.53)	-1.01 (0.65)	-0.42
Observations		3,898	
Country FE		Y	
Baseline controls		Y	

Notes: This table estimates the relative spillover ratios λ directly from the data. The regression specification is given as $y_{it+h} - y_{it-1} = \alpha_{ih} + \phi\Delta\text{IRE}_{it-\ell} + \psi(\Delta\text{IRE}_{it-\ell} \times \text{Defaulters}_i) + \epsilon_{it+h}$, where Defaulters_i is a time-invariant indicator equal to one if country i has ever experienced a sovereign default episode. The coefficient ϕ captures spillover effects for never-defaulting countries, while ψ captures additional spillover effects among defaulting countries. The relative spillover ratio is defined as $\lambda \equiv (\phi + \psi)/\phi$, measuring spillovers for defaulting countries relative to never-defaulting countries. Negative values of λ may arise when spillover effects have opposite signs across the two groups. The sample period is 1970–2010. The dependent variable is the long-difference of log real GDP per capita, $y_{t+h} - y_{t-1}$, and the independent variable is the second lag of “interest rate exposure,” ΔIRE_{it-2} . All columns include (i) country fixed effects, (ii) the baseline control set $\mathbf{X}_{it}$ defined in [Section 1.4.2](#), and (iii) the same 3,898 observations. Robust standard errors clustered at the country level are reported in parentheses. Statistical significance is indicated as *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table 1.6: LP-SSIV results with residualized interest rate changes as shocks
(a) First-stage between $\Delta \text{IRE}_{it-\ell}^{\text{Resid}}$ with S&P first-year default indicators

Dependent variable: S&P first-year binary (0/1) default indicator				
	(1)	(2)	(3)	(4)
$\Delta \text{IRE}_{it-1}^{\text{Resid}}$	0.007 (0.007)			0.013* (0.007)
$\Delta \text{IRE}_{it-2}^{\text{Resid}}$		0.028*** (0.007)		0.031*** (0.007)
$\Delta \text{IRE}_{it-3}^{\text{Resid}}$			0.010 (0.006)	0.014** (0.006)

(b) 2SLS Estimates on the cost of sovereign default cost with $\Delta \text{IRE}_{it-2}^{\text{Resid}}$

Horizons h	Output response ($y_{it+h} - y_{it-1}$)		AR test	Baseline = Resid
	LP-SSIV	Residualized LP-SSIV	p-value	p-value
	(1)	(2)	(3)	(4)
$h = 0$	-7.99 (5.28)	-5.50 (4.81)	0.26	0.21
$h = 1$	-18.48** (8.08)	-16.45** (7.84)	0.03**	0.52
$h = 2$	-15.78* (8.60)	-12.81 (9.00)	0.16	0.45
$h = 3$	-18.25* (9.52)	-14.14 (10.48)	0.18	0.37
$h = 4$	-9.47 (10.72)	-16.98 (12.84)	0.18	0.15
$h = 5$	-5.70 (10.55)	-23.07 (14.40)	0.09*	0.01**
$h = 6$	-0.52 (11.17)	-18.11 (14.95)	0.21	0.02**
Joint significance	0.104	0.045**		
Instrument	ΔIRE_{it-2}	$\Delta \text{IRE}_{it-2}^{\text{Resid}}$		
KP weak IV		20.92		
Observations		3,898		

Notes: This table compares the baseline LP-SSIV estimates in Table 1.3 with the residualized LP-SSIV specification. It also reports p-values from the joint significance test across all horizons (i.e., $\beta_0 = \dots = \beta_6 = 0$), the Kleibergen-Paap weak instrument (F-statistic), Anderson-Rubin weak instrument test (F-statistic) and Baseline SSIV=Residualized SSIV coefficient equality tests (stacking method) at each horizon h . All columns include (i) country fixed effects, (ii) the baseline control set $\mathbf{X}_{it}$ defined in Section 1.4.2, and (iii) the same 3,898 observations. Robust standard errors clustered at the country level are reported in parentheses. Statistical significance is indicated as *** p<0.01, ** p<0.05, * p<0.1.

Table 1.7: Default cost estimates with narrative monetary shocks IRE_{it}^{USRR} and IRE_{it}^{UKCH}
(a) First-stage between $IRE_{it-\ell}^{USRR}$ and $IRE_{it-\ell}^{UKCH}$ with S&P first-year default indicators

Dependent variable: S&P first-year binary (0/1) default indicator				
	(1)		(2)	(3)
IRE_{it}^{USRR}	0.017* (0.009)			0.022** (0.010)
IRE_{it}^{USRR}	0.039** (0.017)			0.042** (0.018)
IRE_{it-1}^{USRR}			0.015 (0.012)	0.019 (0.012)
IRE_{it-1}^{UCKH}			0.024 (0.017)	0.044** (0.019)
(b) 2SLS Estimates on the cost of sovereign default cost with IRE_{it}^{USRR} and IRE_{it}^{UKCH}				
	Output response ($y_{it+h} - y_{it-1}$)		AR test	OLS=IV
Horizons h	LP-OLS	Narrative LP-SSIV	p-value	p-value
	(1)	(2)	(3)	(4)
$h = 0$	-2.54*** (0.66)	-8.15 (9.64)	0.68	0.56
$h = 1$	-3.42*** (1.04)	-20.55 (17.72)	0.34	0.35
-				
$h = 2$	-4.07*** (1.17)	-8.16 (20.00)	0.63	0.87
$h = 3$	-3.85*** (1.34)	-2.96 (21.55)	0.76	0.94
Joint significance	0.00	0.58		
Instrument	IRE_{it}^{USRR} and IRE_{it}^{UKCH}			
KP weak IV		4.29		
Observations		3,235		

Notes: This table presents the LP-SSIV coefficient estimates of sovereign default costs using shift-share instruments constructed from the narrative monetary shocks of [Romer and Romer \(2004\)](#) and [Cloyne and Hurtgen \(2016\)](#). It also reports p-values from the joint significance test, the Kleibergen-Paap and the Anderson-Rubin weak instrument test, and OLS=IV coefficient equality tests as in [Table 1.3](#). Annual narrative monetary shocks are defined as the largest-magnitude monthly shock (sign retained) within each year, standardized within each country, and then interact with endogenous currency denomination shares denominated in U.S. dollars and British pounds. The sample period (1975–2007) differs from that in [Table 1.3](#) (1970–2010) because of data constraints, resulting in different LP-OLS estimates and a different number of observations relative to [Table 1.3](#). All specifications include (i) country fixed effects, (ii) the baseline control set $\mathbf{X}_{it}$ defined in [Section 1.4.2](#), with the exception of one lead and one lag of both IRE_{it}^{USRR} and IRE_{it}^{UKCH} and incomplete share controls excluding U.S. dollar and British pound debt, and (iii) the same 3,235 observations. Robust standard errors clustered at the country level are reported in parentheses. Statistical significance is indicated as *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table 1.8: Default cost estimates with sovereign arrears and $\Delta \text{IRE}_{it-\ell}$ (a): First-stage between sovereign arrears and $\Delta \text{IRE}_{it-\ell}$

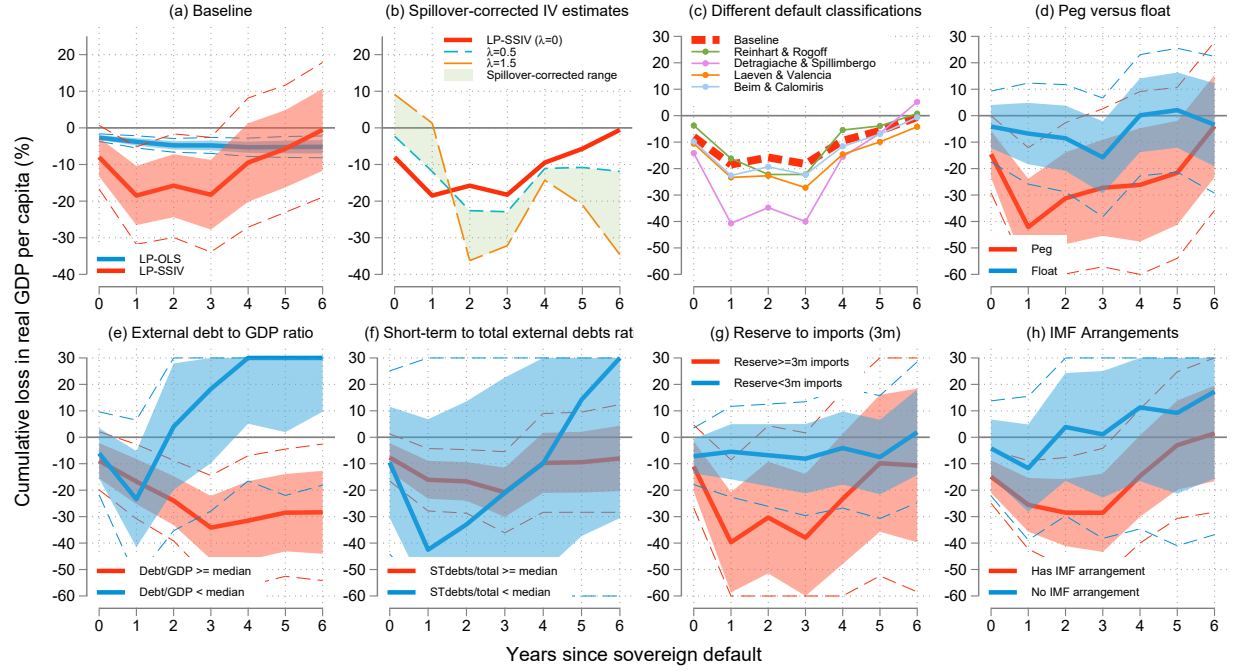
Dependent variable	$\Delta \ln(\text{arrears})$ (1)	$\Delta \ln(1 + \text{arrears})$ (2)	$\Delta \ln(1 + 100 \times \frac{\text{arrears}}{\text{debt}})$ (3)
ΔIRE_{it-1}	0.046 (0.040)	-0.004 (0.024)	-0.006 (0.011)
ΔIRE_{it-2}	0.070* (0.041)	0.002 (0.030)	-0.003 (0.015)
ΔIRE_{it-3}	0.126*** (0.044)	0.053**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

(b) 2SLS Estimates on the cost of sovereign default cost with IRE_{it-3}

Dependent variable: cumulative loss in real GDP per capita (%)						
Independent variable	$\Delta \ln(\text{arrears})$		$\Delta \ln(1 + \text{arrears})$		$\Delta \ln(1 + 100 \times \frac{\text{arrears}}{\text{debt}})$	
	LP-SSIV (1)	AR (2)	LP-SSIV (3)	AR (4)	LP-SSIV (5)	AR (6)
$h = 0$	-1.77 (1.41)	0.18	-3.16 (2.74)	0.18	-8.03 (7.17)	0.17
$h = 1$	-3.02 (2.34)	0.17	-4.04 (3.98)	0.28	-10.42 (10.19)	0.26
$h = 2$	-5.25* (2.96)	0.04**	-8.62 (5.28)	0.05**	-21.84 (13.71)	0.04**
$h = 3$	-4.58 (3.40)	0.14	-7.36 (5.57)	0.14	-18.67 (14.05)	0.13
$h = 4$	-4.02 (3.77)	0.25	-6.15 (5.98)	0.28	-15.66 (14.92)	0.27
$h = 5$	-1.25 (3.88)	0.75	1.98 (6.61)	0.76	4.51 (16.40)	0.78
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	

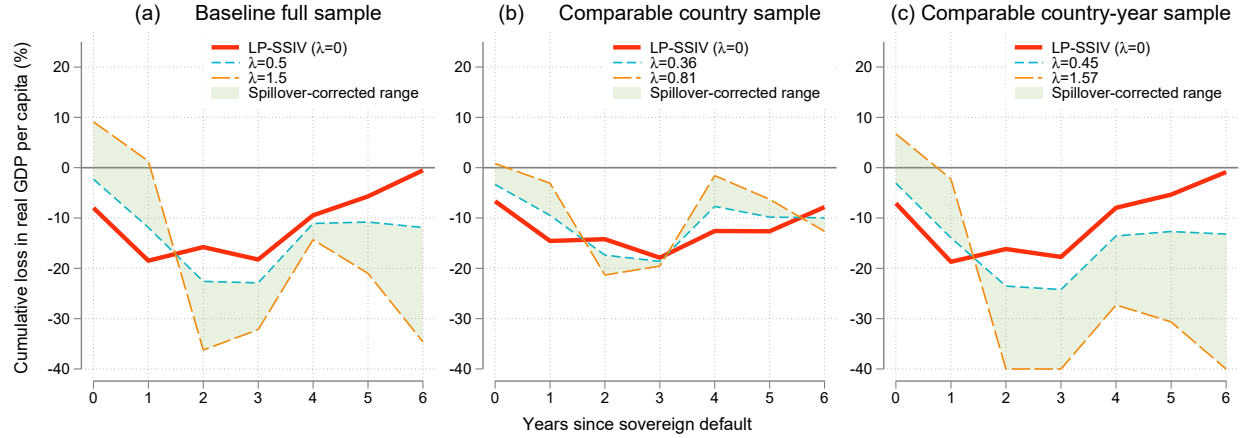
Notes: This table presents the first-stage and LP-SSIV estimates of the cost of sovereign default using continuous default intensity measures. It also reports p-values of relevant tests as in Table 1.3. The IDS arrears data refer to late repayments in the public and publicly guaranteed (PPG) sector of debtor countries, summing both principal and interest arrears and admitting both zero and strictly positive values. To distinguish missing entries from real zeros, I compare values across different versions of the IDS database (see Table A1). The sample period is 1970–2020, compared to the 1970–2010 in Table 1.3, and the number of observations in columns (1)–(2) differs from the remaining specifications because $\Delta \ln(\text{arrears})$ excludes zero values, whereas the other two transformations admit zeros. All columns include (i) country fixed effects, (ii) the baseline control set $\mathbf{X}_{it}$ defined in Section 1.4.2, with the exception of including ΔIRE_{it-1} and ΔIRE_{it-2} to account for lead-lag exogeneity. Robust standard errors clustered at the country level are reported in parentheses. Statistical significance is indicated as *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Figure 1.1: Baseline results, different default classifications, and state-dependent cost.



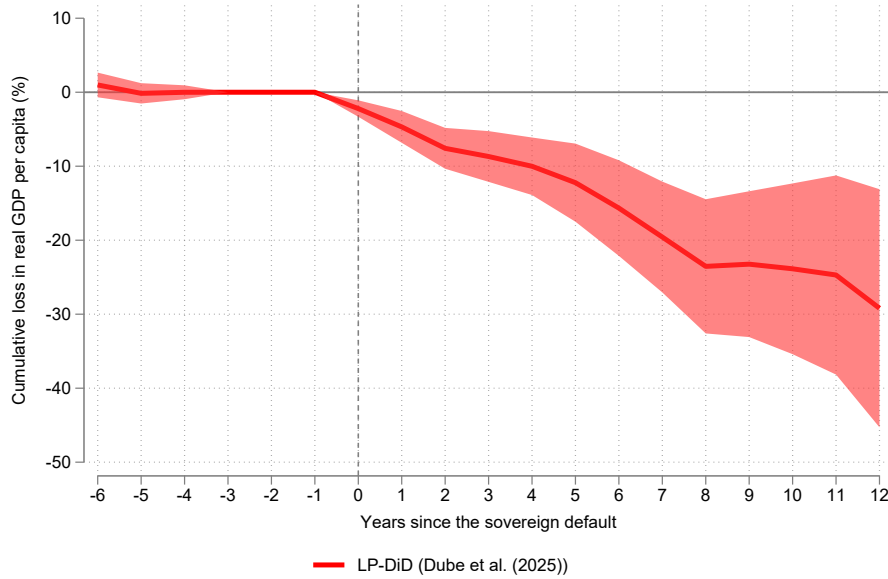
Notes: This figure presents the baseline and state-dependent LP-SSIV estimates of the cost of sovereign default. The solid lines report the point estimates; the shaded areas are one-standard-deviation (68%) confidence intervals; and dashed lines are the 90% confidence intervals. Panels (a) and (b) use an output loss scale from -40% to 20% , while Panels (c) and (d) use a scale from -60% to 30% . Panel (a) shows the baseline LP-OLS and LP-SSIV coefficient estimates reported in Table 1.3. Panel (b) provide a credible range on the spillover-corrected IV estimates using the control-function approach. The solid red line reproduces the baseline LP-SSIV estimate under the assumption of no spillovers to defaulting countries (i.e., $\lambda = 0$). The shaded green region spans the plausible range of spillover-corrected IV estimates under alternative spillover ratios $\lambda > 0$, where defaulting countries are assumed to experience λ times the spillover effect of never-defaulting countries. A value of $\lambda = 1$ implies equal spillover exposure across defaulting and never-defaulting countries, while $\lambda > 1$ implies larger spillover effects for defaulting countries relative to 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 Ilzetzi, Reinhart and Rogoff (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 1.3. Robust standard errors are clustered at the country level to address serial correlation.

Figure 1.2: Robustness of spillover-corrected IV estimates using comparable subsamples.

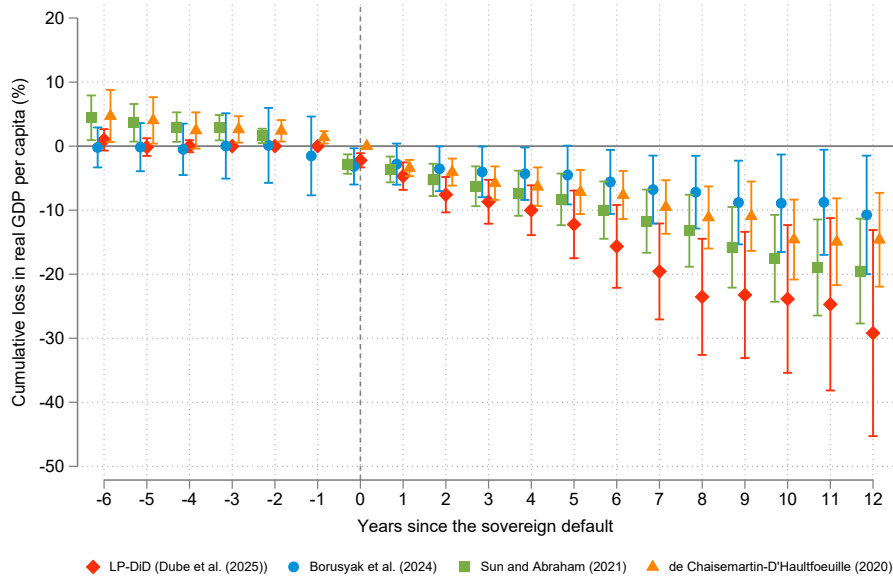


Notes: This figure examines the robustness of the control-function approach in providing a credible range of spillover-corrected IV estimates across alternative comparison samples. Panel (a) reproduces the baseline spillover-corrected estimates reported in Figure 1.1 using the full sample (3,898 country-year observations). Panel (b) restricts the analysis to a comparable-country sample constructed using propensity-score matching at the *country* level. Countries are matched based on similar observable characteristics, including average real GDP per capita growth, inflation, exchange rate changes, the share of external debt to GDP, capital-account openness, banking, currency, and political crises, democracy indicators, and incomplete shares. The resulting comparable-country-set contains 92 countries, including 55 defaulting and 37 never-defaulting countries. Panel (c) further restricts the analysis to a comparable-country-year sample constructed using propensity-score matching at the *country-year* level based on the same set of observable characteristics measured annually, retaining 3,786 of 3,898 observations (97.1% of the baseline sample). In each panel, both the spillover ratio λ , and the OLS spillover coefficient for never-defaulting countries, $\hat{\phi}^{\text{ND}}$, are re-estimated within the corresponding subsample. The solid red line reports the baseline LP-SSIV estimate under the assumption of no spillovers to defaulting countries (i.e., $\lambda = 0$). The shaded green region spans the plausible range of spillover-corrected IV estimates under alternative values of $\lambda > 0$, where defaulting countries are assumed to experience λ times the spillover effect of never-defaulting countries. A value of $\lambda = 1$ implies equal spillover exposure across defaulting and never-defaulting countries, while $\lambda > 1$ implies larger spillover effects for defaulting countries relative to never-defaulting countries.

Figure 1.3: Estimating default cost with difference-in-differences estimators.
(a) Local projection–difference-in-differences (LP-DiD)



(b) Alternative difference-in-differences estimators



Notes: This figure plots the estimated cost of sovereign default using various difference-in-differences (DiD) estimators that address staggered treatment in the literature. All DiD specifications include both country and year fixed effects, as well as the baseline controls specified in Table 1.3. 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 1.11 for additional results with Callaway and Sant’Anna (2021) and the main text for reference on these DiD estimators.

1.10 A PROBABILISTIC SOVEREIGN DEFAULT MODEL

The Online Appendix begins by presenting a simple probabilistic sovereign default model that illustrates the feedback between default probabilities and rollover interest rates induced by foreign monetary tightening. [Section 1.11](#) imposes log-normal distributions to derive a closed-form expression for the steady-state default probability. [Section 1.12](#) shows how the multiple-equilibria property of this probabilistic model naturally links to a *local average treatment effect* (LATE) interpretation. [Section 1.13](#) and [Section 1.14](#) assume Fréchet-distributed exchange rate depreciation to derive a shift-share representation of sovereign default probability based on the interaction between currency-denomination shares and foreign monetary shocks. [Section 1.15](#) further explores the stability of currency-denomination shares under the same distributional structure.

The framework is intentionally stylized. Its purpose is not to provide a new theory of sovereign default, but rather to clarify three economic mechanisms underlying the empirical design: foreign-currency debt exposure, rollover-rate amplification, and the local average treatment effect interpretation of the instrument.

1.10.1 Environment, Timing, and Debt Contract

The model consists of three discrete periods t , $t + 1$, and $t + 2$, together with an initial steady-state period t_0 . Normalizing the funding requirement in 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} denotes the known nominal exchange rate (pesos per dollar) at t . The loan is denominated in currency k and carries a gross interest factor R_t^{ik} , determined at t and due at $t + 1$.²³

When the stochastic counterpart E_{t+1}^{ik} is realized in period $t + 1$, the peso-equivalent

²³While there is growing interest in local-currency-denominated debt among emerging markets, this paper primarily examines foreign-currency external debt and therefore abstracts from domestic-currency borrowing. The “original sin” literature—the idea that developing countries cannot borrow internationally in their own currency—provides justification for temporarily setting aside analysis of local-currency-denominated debt ([Eichengreen and Hausmann, 1999](#)).

repayment obligation for this dollar-denominated debt becomes $\frac{1}{E_t^k} R_t^{ik} E_{t+1}^{ik} = R_t^{ik} \mathcal{E}_{t+1}^{ik}$, where $\mathcal{E}_{t+1}^{ik} \equiv \frac{E_{t+1}^{ik}}{E_t^k}$ denotes the gross depreciation factor. A value of $\mathcal{E}_{t+1}^{ik} > 1$ indicates 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 at $t + 2$, to reflect the default probability anticipated for period $t + 2$ revealed at $t + 1$.

1.10.2 Exchange Rate Determination

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

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

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

$$\mathcal{E}_{t+1}^{ik} \equiv \frac{E_{t+1}^{ik}}{E_t^{ik}} = \frac{\frac{M_{t+1}^i}{M_{t+1}^k}}{\frac{M_t^i}{M_t^k}} \times \frac{\bar{L}^k}{\bar{L}^i} = \frac{\mathcal{M}_{t+1}^i}{\mathcal{M}_{t+1}^k} \propto \frac{1}{\mathcal{M}_{t+1}^k} \quad (1.15)$$

where $\mathcal{M}_{t+1}^k \equiv \frac{M_{t+1}^k}{M_t^k}$ denotes the gross money supply factor of country k from t to $t + 1$. For tractability, I assume no domestic monetary response and set $\mathcal{M}_{t+1}^i = 1$. The key mechanism is that a contractionary foreign monetary shock (i.e., $\mathcal{M}_{t+1}^k < 1$, equivalently $\mathcal{M}_{t+1}^k - 1 = \frac{M_{t+1}^k - M_t^k}{M_t^k} = \% \Delta M_{t+1}^k < 0$) raises $\mathcal{E}_{t+1}^{ik}$ and therefore *depreciates* the peso, ceteris paribus.

1.10.3 Fiscal Capacity and Default Probability

Define $\mathcal{T}_{t+1}^i = \frac{T_{t+1}^i}{D_t^i}$ as the realized revenue-to-debt ratio of debtor i at $t + 1$. Assume that the debtor defaults whenever realized revenue falls short of the peso-equivalent foreign-currency

repayment obligation $R_t^{ik} \mathcal{E}_{t+1}^{ik}$:

$$\pi_{t+1}^{iG} \equiv \mathbb{P}(\mathcal{T}_{t+1}^i \leq R_t^{ik} \mathcal{E}_{t+1}^{ik}) \quad (1.16)$$

Therefore, conditional on the gross interest factor R_t^{ik} , a contractionary foreign monetary shock depreciates the peso, raises $\mathcal{E}_{t+1}^{ik}$ and increases the government default probability π_{t+1}^{iG} (i.e., the exchange rate channel).

1.10.4 Risk-Neutral Pricing of Rollover Debt

The previous subsection shows how a foreign monetary tightening increases the local-currency value of foreign-currency debt and therefore raises default probability. This subsection introduces risk-neutral investors who price rollover debt in response to higher *perceived* future default risk under adaptive expectations, generating an amplification channel through sovereign risk premia.

Given the government default condition in Equation (1.16), I close the model with a pricing equation that maps perceived default risk into rollover interest rates. Suppose that at $t + 1$, debtor i seeks to roll over the $R_t^{ik} \mathcal{E}_{t+1}^{ik}$ peso-equivalent of dollar-denominated debt service. Let the model-consistent expectation of next-period default risk be $\bar{\pi}_{t+1}^{iG} \equiv \mathbb{E}[\pi_{t+2}^{iG}] = \mathbb{E}[\pi_{t+2}^{iG} | \mathcal{F}_{t+1}]$. If creditors form anchored adaptive expectations, then after observing π_{t+1}^{iG} in period $t + 1$, their perceived default probability—formed at $t + 1$ —for period $t + 2$ is:

$$\pi_{t+2|t+1}^{iR} = \sigma \pi_{t+1}^{iG} + (1 - \sigma) \bar{\pi}_{t+1}^{iG} \quad \text{with} \quad \sigma \in [0, 1] \quad (1.17)$$

Thus, when $\sigma \rightarrow 1$, the model features purely adaptive expectations, whereas when $\sigma \rightarrow 0$, expectations become fully anchored to the steady-state prior.²⁴

Then risk-neutral creditors choose the gross rollover factor R_{t+1}^{ik} —set at $t + 1$ and due

²⁴If random variables $(\mathcal{T}_{t+1}^i, \mathcal{M}_{t+1}^k)$ are independently drawn across time, then rational expectations imply a constant steady-state anchor as $\mathbb{E}_{t+1}[\pi_{t+2}^{iG}] = \mathbb{E}[\pi_{t+2}^{iG}] \equiv \pi_0^i$. In that case, Equation (1.17) becomes $\pi_{t+2|t+1}^{iR} = \sigma \pi_{t+1}^{iG} + (1 - \sigma) \pi_0^i$.

at $t+2$ —such that expected return at $t+2$ equals the risk-free benchmark return in currency k , denoted as R_{t+1}^{fk} :

$$\pi_{t+2|t+1}^{iR}(0) + (1 - \pi_{t+2|t+1}^{iR})R_{t+1}^{ik} = R_{t+1}^{fk} \quad (1.18)$$

where $\pi_{t+2|t+1}^{iR}$ is an exogenous variable (input) determined by Equation (1.17) and R_{t+1}^{ik} is an endogenous variable (output). The pricing equation implies a sovereign risk premium $RP(\pi_{t+2|t+1}^{iR})$ as $R_{t+1}^{ik} = \frac{R_{t+1}^{fk}}{1 - \pi_{t+2|t+1}^{iR}} = R_{t+1}^{fk} \left(1 + RP(\pi_{t+2|t+1}^{iR})\right) > R_{t+1}^{fk}$, and we can show that:

$$RP(\pi_{t+2|t+1}^{iR}) = \frac{\pi_{t+2|t+1}^{iR}}{1 - \pi_{t+2|t+1}^{iR}} \quad \text{and} \quad RP'(\cdot) = \frac{1}{\left(1 - \pi_{t+2|t+1}^{iR}\right)^2} > 0 \quad (1.19)$$

so that higher perceived default risk leads to a higher rollover interest rate accordingly.

1.10.5 Extension: Partial Default

The specification in Section 1.10.4 extends naturally to the case of partial default: instead of assuming an all-or-nothing repayment outcome, suppose debtor i pledges to liquidate all tradable assets and repays creditors in dollars upon default. Reinterpret $\pi_{t+2|t+1}^{iR} \in [0, 1]$ as the expected share of loss given default at $t+2$, so that creditors expect to recover $1 - \pi_{t+2|t+1}^{iR}$ of the disbursed debt. Since repayments are pre-denominated in dollars, conditional on repayment, the spot exchange rate fluctuations matter to debtor but are irrelevant to creditors (since they are promised to be repaid back in dollars). Pricing therefore equates expected dollar repayment on the surviving fraction of rollover debt to the risk-free borrowing cost:

$$(1 - \pi_{t+2|t+1}^{iR})R_{t+1}^{ik} = R_{t+1}^{fk}$$

which is identical to the pricing equation in Equation (1.18). Therefore, both the binary default and partial default specifications imply the same qualitative result: higher perceived

default risk $\pi_{t+2|t+1}^{iR}$ increases the rollover interest rate R_{t+1}^{ik} (i.e., the pricing channel).

1.10.6 Dynamic Feedback Between Default Risk and Rollover Premiumia

To highlight the feedback mechanism between default probability $\pi_{t+h|t+h-1}^i$ and the associated risk premium $RP(\pi_{t+h|t+h-1}^i)$, consider a parsimonious setting in which the initial steady-state equilibrium $A = (\pi_0^i, R_0^{ik})$ lies in the interior of the state space at period t_0 .²⁵ Suppose a temporary foreign contractionary monetary shock $\mu_k > 0$ occurs in period t , such that $\% \Delta E_{t+1}^{ik} = \mu_k > 0$, so the peso *depreciates*.²⁶ Unlike the baseline setup, the shock is modeled as a single realization rather than a stochastic process, although the revenue-to-debt ratio $\mathcal{T}_{t+1}^i$ remains stochastic.

To illustrate the key mechanism, Panel (a) of [Figure 1.4](#) illustrates the partial equilibrium effect of a temporary foreign monetary tightening in period $t + 1$. Since the peso value of dollar-denominated obligations increases, $\pi^i > \pi_0^i$ for all R^{ik} , the government default probability rises for every given level of the interest factor, shifting the government default curve upward. Since the interest factor is predetermine in period t , the equilibrium moves from point A to point B , with $\pi_B^i > \pi_0^i$.

Suppose next that debtor i must roll over its entire debt obligation into period $t + 2$. Observing the higher default probability π_B^i due to peso depreciation, the risk-neutral global investors reprice the rollover debt and demand a risk premium by endogenously raising rollover interest factor from R_0^{ik} to R_C^{ik} with $R_C^{ik} > R_0^{ik}$.

As the rollover rate rises to R_C^{ik} , future debt-service obligations become larger, which further increases default risk in period $t + 3$ to $\pi_D^i > \pi_B^i > \pi_0^i$, and the rollover interest factor

²⁵[Section 1.11](#) established the existence of such steady-state under log-normal distributions. [Section 1.12](#) releases this assumption and show the interior and corner solutions connect to the local average treatment effect interpretation

²⁶From [Equation \(1.15\)](#), we obtain $\Delta \% E_{t+1}^{ik} \approx \ln \mathcal{E}_{t+1}^{ik} = -\ln \mathcal{M}_{t+1}^k \approx -\% \Delta M_{t+1}^k$. Define $\mu_k \equiv -\% \Delta M_{t+1}^k$, that is, $\mu_k > 0$ corresponds to a contractionary foreign monetary policy shock. For sufficiently small shocks, $\% \Delta E_{t+1}^{ik} = -\% \Delta M_{t+1}^k = \mu_k > 0$.

subsequently increases to R_E^{ik} , further raising the peso-value obligation. The higher debt burden again raises perceived default probability to π_F^i , which induces another increase in rollover premia, raising the rollover interest factor to R_G^{ik} and further worsening repayment prospects. Repeated iterations of this mechanism generate a feedback loop between sovereign default risk and rollover interest rate that may eventually push the economy toward a *self-fulfilling* default equilibrium.²⁷

In addition to increasing the local-currency value of foreign-currency liabilities, a contractionary foreign monetary shock may also increase the borrowing cost in currency k itself—even holding π^i fixed—by raising global interest rates and reducing market liquidity. Consequently, Panel (b) of [Figure 1.4](#) illustrates an amplification effect in which the government default curve shifts leftward while the investor pricing curve shifts rightward. As default risk rises because of peso depreciation, borrowing costs simultaneously increase, accelerating convergence toward the default equilibrium.

The same intuition can be formalized using comparative statics analysis under the general equilibrium system. Combining [Equation \(1.16\)](#), [Equation \(1.17\)](#), and [Equation \(1.18\)](#), we have

$$\begin{aligned} R_{t+1}^{ik} &= \frac{R_{t+1}^{fk}}{1 - \pi_{t+2|t+1}^{iR}} = \frac{\mathcal{R}(\mathcal{M}_{t+1}^k)}{1 - (\sigma\pi_{t+1}^{iG} + (1 - \sigma)\bar{\pi}_{t+1}^{iG})} \\ &= \frac{\mathcal{R}(\mathcal{M}_{t+1}^k)}{1 - \sigma F_T R_t^{ik} \mathcal{E}(\mathcal{M}_{t+1}^k) - (1 - \sigma)\pi_{t+1}^{iG}} \end{aligned}$$

where $F_T(\cdot)$ denotes the cumulative distribution function of the random variable $\mathcal{T}_{t+1}^i$ and

²⁷The same mechanism applies under a permanent tightening, although the transition path (denoted with primes) differs.

$F'_T(\cdot) > 0$ is the corresponding probability density function. Therefore,

$$\begin{aligned} \frac{\partial R_{t+1}^{ik}}{\partial \mathcal{M}_{t+1}^k} = & \left(\frac{1}{1 - \sigma F'_T(\cdot) - (1 - \sigma) \bar{\pi}_{t+1}^{iG}} \right) \underbrace{\frac{\partial R_{t+1}^{fk}}{\partial \mathcal{M}_{t+1}^k}}_{<0} \\ & + \left(\frac{\sigma F'_T(\cdot) R_t^{ik} R_{t+1}^{fk}}{(1 - \sigma F'_T(\cdot) - (1 - \sigma) \bar{\pi}_{t+1}^{iG})^2} \right) \underbrace{\frac{\partial \mathcal{E}_{t+1}^{ik}}{\partial \mathcal{M}_{t+1}^k}}_{<0} \end{aligned} \quad (1.20)$$

The above expression highlights two channels through which a contractionary foreign monetary shock raises the rollover rate. The first is a *direct* borrowing-cost channel operating through the elevated risk-free benchmark rate in currency k . The second is an indirect risk-premium channel operating through an increase in perceived default risk induced by exchange rate depreciation.

When $\sigma \rightarrow 0$, Equation (1.20) simplifies to

$$\frac{\partial R_{t+1}^{ik}}{\partial \mathcal{M}_{t+1}^k} = \left(\frac{1}{1 - \bar{\pi}_{t+1}^{iG}} \right) \underbrace{\frac{\partial R_{t+1}^{fk}}{\partial \mathcal{M}_{t+1}^k}}_{<0}$$

where $\pi_{t+1}^i \equiv \mathbb{E}_{t+1} [\pi_{t+2}^{iG}]$ denotes the fully anchored rational expectation prior formed at $t+1$ for period $t+2$. In this limiting case, the risk-premium channel becomes substantially weaker because perceived default risk responds little to the realized currency depreciation. More generally, the parameter σ governs the extent to which expectations formation—anchored by policy credibility—amplifies the effects of foreign monetary tightening through the risk-premium channel.

To characterize default probability in period $t+2$, rewrite Equation (1.16) as

$$\pi_{t+2|t+1}^{iG} \equiv \mathbb{P}(\mathcal{T}_{t+2}^i \leq R_{t+1}^{ik} \mathcal{E}_{t+2}^{ik})$$

Assume a one-period foreign contractionary monetary shock such that $\mathcal{E}_{t+1}^{ik} > 0$ and

$\mathcal{E}_{t+2}^{ik} = 1$. If $\mathcal{T}_t^i$ is independently and identically distributed over time, then

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

Differentiating the above equation with respect to $\mathcal{M}_{t+1}^j$ gives $\frac{\partial \pi_{t+2|t+1}^{iG}}{\partial \mathcal{M}_{t+1}^k} < 0$:

$$\begin{aligned} \frac{\partial \pi_{t+2|t+1}^{iG}}{\partial \mathcal{M}_{t+1}^k} &= F_T'(\cdot) \left(\underbrace{\frac{1}{1 - \sigma F_T(\cdot) - (1 - \sigma)\bar{\pi}_{t+1}^{iG}}}_{<0} \right) \frac{\partial R_{t+1}^{fk}}{\partial \mathcal{M}_{t+1}^k} \\ &\quad + F_T'(\cdot) \left(\frac{\sigma R_t^{ik} R_{t+1}^{fk} F_T'(\cdot)}{(1 - \sigma F_T(\cdot) - (1 - \sigma)\bar{\pi}_{t+1}^{iG})^2} \right) \underbrace{\frac{\partial \mathcal{E}_{t+1}^{ik}}{\partial \mathcal{M}_{t+1}^k}}_{<0} \end{aligned} \quad (1.21)$$

The above expression shows that a contractionary foreign monetary shock raises default probability through both higher borrowing costs and currency depreciation. The latter channel becomes weaker when expectations are more firmly anchored and sovereign credibility is stronger.

This probabilistic framework therefore generates a self-fulfilling feedback mechanism absent from standard binary default models. A temporary foreign monetary tightening increases the local-currency value of dollar-denominated debt; higher perceived default risk raises rollover interest rate; and higher rollover premia further increase the default likelihood in the next period.²⁸

1.10.7 Extension: Pegged Exchange Rates and Foreign Exchange Reserves

Two additional implications follow from the above model. First, Panel (c) of [Figure 1.4](#) illustrates that even under a pegged exchange rate regime, economies may still be pushed toward a default equilibrium since a foreign monetary tightening directly raises borrowing

²⁸The 2010 euro-zone debt crisis provides a vivid illustration: as markets revised Greece's default probability upward, bond yields surged, and the soaring risk premia made repayment increasingly untenable.

costs and thereby shifts the investor pricing curve rightward. Higher rollover interest rate then increases default risk even when exchange rate movements are constrained by policy.

Second, if interest-rate shocks can trigger self-fulfilling default dynamics, a natural question is why sovereign defaults have become less frequent in recent decades. One explanation is the substantial accumulation of foreign exchange reserves—which effectively augment the resources available to service foreign-currency liabilities—following the 1980s Latin American Debt Crisis and the 1997 Asian Financial Crisis.

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

As a result, Panel (d) of [Figure 1.4](#) shows that the accumulation of the foreign exchange reserve shift the government default curve downward and increase the likelihood of converging to a no-default equilibrium.

1.11 A CLOSED-FORM STEADY-STATE DEFAULT PROBABILITY

This section derives a closed-form expression for the steady-state default probability in the probabilistic sovereign default model developed in [Section 1.10](#). The goal is to show that, under log-normal assumptions on fiscal capacity and foreign monetary shocks, the model delivers an analytically tractable expression for sovereign default probability.

Assume that 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 independence assumption is strong, but it captures the exogeneity idea underlying the empirical design: monetary policy decisions in advanced economy k are assumed not to respond to confounding factors (e.g., wars, political instability, or financial crises) affecting debtor country i .

Proposition 2 *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:*

1. $\ln \mathcal{T}_{t+1}^i$ and $\ln \mathcal{M}_{t+1}^k$ are independent, so $\ln \mathcal{T}_{t+1}^i + \ln \mathcal{M}_{t+1}^k \sim \text{Normal}(\mu_{t_i} + \mu_{m_k}, \sigma_{t_i}^2 + \sigma_{m_k}^2)$
2. The default probability admits a closed-form expression

$$\begin{aligned} \pi_0^i &\equiv \mathbb{P}(\mathcal{T}_{t+1}^i \leq R_t^{ik} \mathcal{E}_{t+1}^{ik}) \\ &= \Phi \left(\underbrace{\frac{\ln \sigma_{ik}}{\Xi}}_{\text{EXR responsiveness}} + \underbrace{\frac{-\left[\ln \frac{T'_i}{D_i}\right]}{\Xi}}_{\text{fiscal capacity}} + \underbrace{\frac{\overbrace{[\ln \mathcal{M}_{t+1}^i]}^{\text{assume}=0} - [\ln \mathcal{M}_{t+1}^k]}{\Xi}}_{\text{monetary prudence}} + \frac{\ln R_t^{ik}}{\Xi} \right) \end{aligned} \quad (1.22)$$

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

To plot the government default curve with default probability on the vertical axis and

the rollover interest rate on the horizontal axis, it is useful to rewrite Equation (1.22) as

$$\pi^i = \Phi \left(\frac{\ln \bar{\sigma}^{ik}}{\Xi} + \frac{-\mathbb{E} \left[\ln \frac{T'_i}{D_i} \right]}{\Xi} + \frac{-\mathbb{E} [\ln \mathcal{M}^k]}{\Xi} + \frac{1}{\Xi} (R^{ik} - 1) \right) \quad \text{if } R^{ik} \approx 1$$

An interior steady state exists whenever the associated z value is finite. Corner solutions arise only when the numerator becomes sufficiently negative (the “never default” region with $\pi_0^i = 0$) or sufficiently positive (the “always default” region with $\pi_0^i = 1$). Under the log-normal distributions, the numerator can be decomposed into five economically interpretable components: (i) exchange rate responsiveness, $\ln \sigma_{ik}$; (ii) fiscal capacity, $-\mathbb{E} \left[\ln \frac{T'_i}{D_i} \right]$, (iii) monetary prudence, $\mathbb{E} [\ln \mathcal{M}_{t+1}^i] - [\ln \mathcal{M}_{t+1}^k]$; (iv) the interest burden, $\ln R_{t+1}^{ik}$; and (v) aggregate uncertainty, $\Xi \equiv \sqrt{\sigma_{ti}^2 + \sigma_{mk}^2}$.

1.11.1 Log-Normal Setup and Independence

The first part of the proposition follows from the fact that $\ln(\cdot)$ is an injective function on the strictly positive domain and from standard properties of the normal distribution. Fix an arbitrary debtor i , currency k , and period $t + 1$, and suppress time when clear for simplification. Assume $(\mathcal{T}^i, \mathcal{M}^k)$ are both strictly positive and mutually independent random variables, then:

Lemma 1 *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. The result follows from a standard change-of-variables argument using integration by substitution. Using that the fact that $\ln(\cdot)$ is an injective function at the *strictly positive* domain, it is sufficient to show that for any mutually *independent* random variables with $X, Y > 0$; $X \perp Y$, then we have $\ln X \perp \ln Y$. First, let $U \equiv \ln X, V \equiv \ln Y$, so we have $X = \exp(U), Y = \exp(V)$. Therefore, the Jacobian matrix of the transformation is given as:

$$J \equiv \begin{pmatrix} \frac{\partial X}{\partial U} & \frac{\partial X}{\partial V} \\ \frac{\partial Y}{\partial U} & \frac{\partial Y}{\partial V} \end{pmatrix} = \begin{pmatrix} \exp(U) & 0 \\ 0 & \exp(V) \end{pmatrix} \Rightarrow \det(J) = \exp(U) \exp(V)$$

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

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

Hence U and V are independent, implying $\ln X \perp \ln Y$. Applying the result to $(\mathcal{T}^i, \mathcal{M}^k)$ proves the claim. ■

To derive a closed-form solution for the steady-state default probability π_0^i , assume that $(\mathcal{T}^i, \mathcal{M}^k)$ are mutually independent log-normal random variables.

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

Proof. By definition of the log-normal distribution, we have $\ln \mathcal{T}^i \sim \text{Normal}(\mu_{ti}, \sigma_{ti}^2)$ and likewise $\ln \mathcal{M}^k \sim \text{Normal}(\mu_{mk}, \sigma_{mk}^2)$. 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 normal random variables is itself normally distributed,

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

■

The mutual independence assumption is considerably stronger than a zero-covariance restriction because it imposes restrictions on the entire joint probability density function. To be specific, $X \perp Y \Leftrightarrow f_{X \times Y}(x, y) = f_X(x)f_Y(y)$. Therefore, observing $\text{Cov}(\mathcal{T}'_i, \mathcal{M}'_j) = 0$ in

the data is a necessary—but not sufficient—condition to establish mutual independence.

1.11.2 Deriving the Closed-Form Steady-state Default Probability

The previous subsection establishes that $\mathcal{Z}^{ik} \equiv \ln \mathcal{T}^i + \ln \mathcal{M}^k$ is normally distributed. The next step standardizes this random variable and applies the sovereign default condition.

Lemma 3 *If $\mathcal{Z}^{ik} \equiv \ln \mathcal{T}^i + \ln \mathcal{M}^k \sim \text{Normal}(\mu_{ti} + \mu_{mk}, \sigma_{ti}^2 + \sigma_{mk}^2)$, then $\frac{\mathcal{Z}^{ik} - (\mu_{ti} + \mu_{mk})}{\sqrt{\sigma_{ti}^2 + \sigma_{mk}^2}} \sim \text{Normal}(0, 1)$. Hence,*

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

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

Last but not least, assume domestic money supply growth is constant and an incomplete passthrough of foreign monetary policy on exchange rate, that is, $\mathcal{E}^{ik} = \bar{\sigma}^{ik} \frac{\bar{\mathcal{M}}^i}{\bar{\mathcal{M}}^k}$, where $\bar{\sigma}^{ik}$ 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:

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

Notice that since $\ln \mathcal{M}_{t+1}^k = \ln M_{t+1}^k - \ln M_t^k \approx \% \Delta M_{t+1}^k$, foreign money growth maps naturally into foreign monetary shocks. Similarly, if $\frac{T^i}{D^i} \approx 1$, then $\ln \frac{T^i}{D^i} \approx \frac{T^i}{D^i} - 1 \propto (\mathcal{T}^i - D^i - 1)$, so the revenue-to-debt ratio provides a natural measure of fiscal capacity regarding whether, on average, debtor i 's revenue exceeds its debt obligation. Finally, if $R^{ik} \approx 1$, then $\ln R^{ik} \approx R^{ik} - 1 = r^{ik}$, where r^{ik} denotes the net interest rate.

1.11.3 Economic Interpretation

The probability expression in Equation (1.22) can be decomposed into four economically meaningful components.

The first term, $\ln \bar{\sigma}^{ik}$, captures sovereign i 's *exchange rate responsiveness* to foreign monetary tightening. A stronger exchange rate response magnifies the peso value of foreign-currency-denominated debt obligations and therefore raises default probability.

The second term, $-\mathbb{E} \left[\ln \frac{T^i}{D^i} \right]$, reflects sovereign i 's *fiscal capacity*. When ex-post revenue $\mathcal{T}_{t+1}^i$ systematically exceeds ex-ante required debt obligations D_t^i , this term becomes more negative and lowers default probability. Conversely, weak fiscal capacity raises default risk. Since this component depends on the entire joint distribution of the revenue-to-debt ratio rather than occasional revenue windfalls (e.g., commodity booms), reductions in long-run steady-state default risk require improvements in fiscal fundamentals, such as broadening the tax base or containing government expenditure.

The third term, $\mathbb{E} [\ln \mathcal{M}^i - \ln \mathcal{M}^k]$, captures sovereign i 's *monetary prudence*, measured by the expected gap between domestic and foreign money growth. Rapid domestic monetary expansion depreciates the peso, makes foreign-currency debt harder to service, and increases default probability. Conversely, more stable domestic monetary policy reduces the likelihood that exchange rate movements amplify foreign-currency debt burdens. 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. Its direct effect is intuitive: higher rollover interest rates raise default probability by increasing repayment obligations. More importantly, [Section 1.10](#) shows that default probabilities may themselves feed back into rollover rates through sovereign risk premia under adaptive expectations: even a small upward revision in perceived default risk can sharply raise short-term rollover costs and push the economy toward a self-fulfilling default equilibrium.

1.12 FROM MULTIPLE EQUILIBRIA TO A LATE INTERPRETATION

The probabilistic sovereign default model in [Section 1.10](#) provides a multiple-equilibria solution that naturally connects to a local average treatment effect (LATE) interpretation of the baseline LP-SSIV estimates. To be specific, a foreign monetary tightening does not affect all sovereign borrowers equally. Countries with strong fiscal fundamentals and high market credibility may remain far from defaulting even after adverse external shocks, whereas countries with persistently weak fundamentals may face a high default likelihood regardless of foreign monetary conditions. Consequently, the baseline LP-SSIV estimates should be interpreted as identifying the causal effect of sovereign default on output loss for countries whose default probability is shifted by foreign monetary shocks through the exchange rate and rollover premium channels described in [Section 1.10](#).

1.12.1 Interpreting Corner Solutions as Always-Takers and Never-Takers

The analysis in [Section 1.11](#) establishes the existence of an interior steady-state equilibrium with default probability satisfying $\pi_0^i \in (0, 1)$. More generally, however, the location of the steady-state equilibrium depends on underlying model parameters and may instead lie at a corner solution.

First, consider the “never default” region shown in Panel (a) of [Figure 1.5](#). In this case, the sovereign default and investor pricing curves intersect only along the horizontal axis both before and after a foreign monetary tightening. The equilibrium default probability therefore remains effectively zero throughout the range of foreign monetary shocks considered in the empirical analysis. These countries are analogous to *never-takers* in the LATE framework.

Second, consider the “always default” region shown in Panel (b) of [Figure 1.5](#). Here, the

sovereign default and investor pricing curves intersect along the vertical axis both before and after the shock, implying a default probability effectively equal to one regardless of foreign monetary conditions. These countries correspond to *always-takers*.

Because neither group changes treatment status in response to the foreign monetary tightening, neither contributes to identification of the causal effect in the baseline LP-SSIV specification.

1.12.2 Compliers and the local average treatment effect

The empirically relevant case is the interior equilibrium in which foreign monetary shocks shift the steady-state equilibrium. Depending on model parameters, changes in $\mathcal{M}_{t+1}^k$ may generate either a single tangency or multiple intersections between the sovereign default and investor pricing curves.

Point A in Panel (c) of [Figure 1.5](#) illustrates the *complier* region that underlies the instrumental-variable strategy developed in [Section 1.4](#). For these countries, a contractionary foreign monetary shock increases the domestic-currency value of foreign-currency liabilities and raises rollover interest rate through higher perceived default risk. The resulting feedback between default probabilities and borrowing costs generates the self-fulfilling adjustment mechanism described in [Section 1.10.6](#). Under the monotonicity condition, foreign monetary tightening increases default probability for this group. The baseline LP-SSIV estimates should therefore be interpreted as a local average treatment effect for countries whose default probability responds monotonically to foreign monetary tightening.

A limitation is that not all sovereign default episodes satisfy this monotonicity condition or are primarily driven by foreign monetary tightening. Point B in Panel (d) of [Figure 1.5](#) illustrates a case in which a temporary foreign monetary tightening initially raises default probability and rollover interest rates, but the economy subsequently reverts back toward its original equilibrium.²⁹ In such cases, foreign monetary tightening is not the dominant force

²⁹Illustrative examples include the Greece's debt crisis during 2010–2012, when many advanced economies

governing sovereign default dynamics, and the instrumental variable design does not cleanly identify a causal effect.

Recognizing these scope conditions clarifies the interpretation of the empirical estimates and motivates complementary identification strategies, such as the difference-in-differences approach discussed in [Section 1.6.4](#), as robustness checks.

approached the zero lower bound following the 2008 Global Financial Crisis, and Argentina’s 2001 sovereign default episode, which was driven primarily by Brazil’s devaluation and the constraints imposed by a rigid currency-board regime.

1.13 A FRÉCHET MODEL OF SOVEREIGN DEFAULT

This section adapts a Fréchet framework of [Eaton and Kortum \(2002\)](#), which uses Fréchet heterogeneity in productivity to characterize international trade patterns, to sovereign borrowing. The object of interest here is sovereign default risk arising from foreign-currency-denominated debt. The common Fréchet structure in exchange rate depreciation delivers a tractable aggregation formula on default probability that will later generate a shift-share representation of sovereign default probability.

Assume a continuum of loan suppliers $j \in [0, 1]$ and a countable set of borrowing currencies $k \in \mathbb{N}$. Within each currency k , perfect competition implies that all lenders charge debtor i the same gross interest factor $R_t^{ik} \geq 1$, determined at period t and due at $t + 1$. Suppose debtor i allocates borrowing across currencies according to ex-post debt-service costs, that is,

$$R_t^{i\$} \mathcal{E}_{t+1}^{i\$} \leq \min_{k \neq \$} \{ R_t^{ik} \mathcal{E}_{t+1}^{ik} \}$$

Fix an arbitrary debtor i and period $t + 1$. Suppose the gross depreciation factor against currency k , denoted $\mathcal{E}_{t+1}^{ik}$, is independently drawn across currencies from Fréchet distributions with the same debtor-specific shape parameter θ^i but different currency-specific location parameters $\bar{e}^k$. For model closure, assume debtor i 's fiscal capacity also follows a Fréchet distribution with the same shape parameter θ^i but a different debtor-specific location parameter $\bar{t}^i$.

For simplicity, suppress subscripts when no ambiguity arises. After relabeling

$$(\pi_{t+1}^i, \bar{e}^k, \theta^i, \bar{t}^i, R_t^{ik}, \mathcal{E}_{t+1}^{ik}, \mathcal{T}_{t+1}^i) \rightarrow (\pi, \bar{e}_k, \theta, \bar{t}, R_k, \mathcal{E}_k, \mathcal{T})$$

we have, for all $k \in \mathbb{N}$,

$$\mathcal{E}_k \sim \text{Fréchet}(\bar{e}_k, \theta) \quad \mathcal{T} \sim \text{Fréchet}(\bar{t}, \theta)$$

From Equation (1.1), a sovereign default occurs whenever

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

Let $u \equiv \tau^{-\theta}$, $du = d(\tau^{-\theta}) = -\theta \tau^{-\theta-1}$. When $\tau \rightarrow 0, u \rightarrow \infty$; when $\tau \rightarrow \infty, u \rightarrow 0$.

Also, define $\Phi \equiv \bar{t} + \sum_k \bar{e}_k R_k^\theta$. The above equation can be further simplified to:

$$\begin{aligned}
\pi &= 1 - (-\bar{t}) \int_\infty^0 \exp(-\Phi u) du \\
&= 1 - \bar{t} \int_0^\infty \exp(-\Phi u) du \\
&= 1 - \bar{t} \left[-\frac{1}{\Phi} \exp(-\Phi u) \right]_0^\infty \\
&= 1 - \frac{\bar{t}}{\bar{t} + \sum_k \bar{e}_k R_k^\theta} \\
&= \frac{\sum_k \bar{e}_k R_k^\theta}{\bar{t} + \sum_k \bar{e}_k R_k^\theta}
\end{aligned}$$

Therefore,

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

1.14 RELATIONSHIP BETWEEN DEFAULT PROBABILITY AND THE SHIFT-SHARE INSTRUMENT

This section builds upon the probabilistic sovereign default model with Fréchet distributed exchange rate depreciation outlined in [Section 1.13](#) to show how its default probability can be locally approximated with a shift-share representation. The objective is to connect the theoretical model developed in [Section 1.13](#) to the empirical shift-share instrument defined in [Equation \(1.5\)](#) of the main text.

Starting from [Equation \(1.2\)](#) in the main text,

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

To linearize the above expression, I first define $A_t^i \equiv \sum_k \bar{e}^k (R_t^{ik})^{\theta^i}$. Let $\bar{R}^{ik}$ denote the steady-state interest factor and define $\bar{A}^i \equiv \sum_k \bar{e}^k (\bar{R}^{ik})^{\theta^i}$, $\bar{\pi}^i = \frac{\bar{A}^i}{\bar{t}^i + \bar{A}^i}$, $\hat{\pi}_{t+1|t}^i \equiv \frac{d\pi_{t+1|t}^i}{\bar{\pi}^i}$, and $\hat{R}_t^{ik} \equiv \frac{dR_t^{ik}}{\bar{R}^{ik}}$ as percentage deviations from their respective steady-state values.

Taking a first-order approximation of $\pi_{t+1|t}^i = \frac{A_t^i}{\bar{t}^i + A_t^i}$ yields:

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

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

Combining the above equations gives

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

where the last step follows from $\bar{\pi}^i = \frac{\bar{A}^i}{\bar{t}^i + \bar{A}^i} \implies 1 - \bar{\pi}^i = \frac{\bar{t}^i}{\bar{t}^i + \bar{A}^i}$.

The resulting expression takes the form of a weighted average of currency-specific interest rate changes, where the weights are determined by steady-state currency exposures. This provides the key link between the theoretical model and the empirical shift-share specification.

To further illustrate the mechanism, suppose a single contractionary foreign monetary shock occurs in currency \$, that is, $\hat{R}_t^{i\$} > 0$ and $\forall k \neq \$, \hat{R}_t^{ik} = 0$. Then

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

The term in parentheses can be interpreted as debtor i 's steady-state exposure to currency \$. Consequently, the response of default probability to a currency-specific interest rate change depends both on the magnitude of the shock and on the debtor's pre-existing exposure to that particular currency \$.

To express the result in percentage-point deviations from steady state, define $\tilde{\pi}_{t+1|t}^i \equiv \pi_{t+1|t}^i - \bar{\pi}^i = d\pi_{t+1|t}^i$ and $\tilde{R}_t^{ik} \equiv R_t^{ik} - \bar{R}^{ik} = dR_t^{ik}$, where both variables are measured as percentage-point deviations from their steady-state values. Repeating the same first-order

approximation yields

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

Again, suppose a single contractionary monetary shock occurs in currency \$, that is, $\tilde{R}_t^{i\$} \equiv R_t^{i\$} - \bar{R}^{i\$} > 0$ and $\forall k \neq \$, \tilde{R}_t^{ik} = 0$. Then

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

This expression is also a shift-share representation of default probability: the effect of a currency-specific interest rate change on sovereign default risk equals the product of a shock component, $\tilde{R}_t^{i\$}$, and a predetermined exposure component $\left(\frac{\bar{e}^{\$} \left(\bar{R}^{i\$}\right)^{\theta^i}}{\sum_k \bar{e}^k \left(\bar{R}^{ik}\right)^{\theta^i}}\right)$, which, analogous to [Section 1.14](#), summarizes debtor i 's steady-state dependence on currency \$.

Thus, whether interest rate movements are expressed as percentage deviations or percentage-point deviations from steady state, the model delivers the same economic insight: sovereign default likelihood responds to foreign monetary shocks through the interaction between currency-specific shocks and predetermined currency exposures. This provides the theoretical foundation for the empirical shift-share instrument used in the main analysis.

1.15 CURRENCY SHARES IN THE FRÉCHET SOVEREIGN DEFAULT MODEL

This section illustrates conditions under which currency-denomination shares remain relatively stable over time. The key idea is that observed currency shares effectively summarize predetermined exposure to foreign monetary shocks. While the full portfolio-allocation problem may be complex, a simple probabilistic discrete-choice framework shows that currency-denomination shares depend on relative debt-service costs across currencies. This framework also clarifies why the currency composition of debt does not necessarily adjust immediately in response to temporary monetary shocks.

To characterize currency-denomination choice, Let $s(k)$ denote the total share of debt denominated in currency k from all creditors. For simplicity, suppress subscripts when no ambiguity arises. After relabeling

$$(\theta^i, R_t^{ik}, \mathcal{E}_{t+1}^{ik}, s^i(k)) \rightarrow (\theta, R^k, \mathcal{E}^k, s(k))$$

consider an arbitrary borrowing currency $\$$ and assume the debtor i chooses to borrow in currency $\$$ whenever its realized debt-service cost is weakly lower than that of all alternative borrowing currencies:

$$R^{\$} \mathcal{E}^{\$} \leq \min_{k \neq \$} \{R^k \mathcal{E}^k\}$$

Under repeated draws and a continuum of competitive lenders under perfect competition, the binary borrowing decision is a Bernoulli random variable with the probability density function:

$$\mathbf{1}^{ij}(\$) = \begin{cases} 1, & \text{with } \mathbb{P}(R^{\$} \mathcal{E}^{\$} \leq \min_{k \neq \$} \{R^k \mathcal{E}^k\}), \\ 0, & \text{with } 1 - \mathbb{P}(R^{\$} \mathcal{E}^{\$} \leq \min_{k \neq \$} \{R^k \mathcal{E}^k\}) \end{cases}$$

Assume that exchange rate depreciation realizations are independently and identically Fréchet distributed with the same debtor-specific location parameter $\bar{e}^i$ and same shape parameter θ . Under repeated draws and a continuum of price-taking lenders, the borrowing share $s(\$)$ is:

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

While the general expression must be computed numerically, a closed-form solution exists in the two-currency case.

1.15.1 A Two-Currency Closed Form

Suppose debtor i chooses only between U.S. dollar (\$) and British pound (£) borrowing. The borrowing share in the U.S. dollar, $s(\$)$, simplifies to:

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

Therefore, we have:

$$s(\$) = \frac{(R^{\mathcal{L}})^{\theta}}{(R^{\$})^{\theta} + (R^{\mathcal{L}})^{\theta}} \quad (1.23)$$

Moreover, comparative static analysis indicates that

$$\frac{\partial s(\$)}{\partial R^{\$}} < 0 \quad \frac{\partial s(\$)}{\partial R^{\mathcal{L}}} > 0$$

and, in particular,

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

where $r^{\mathcal{S}} \equiv R^{\mathcal{S}} - 1$ denotes the net interest rate.

The above expression shows that currency-share adjustment depends not only on relative borrowing costs, but also on how frequently extreme depreciation events happen. When U.S. dollar borrowing is cheaper than British pound borrowing, the U.S. dollar share increases with the Fréchet shape parameter θ . Since a larger θ corresponds to a thinner-tailed Fréchet distribution, currency-denomination shares become more responsive to systematic borrowing-cost differences when extreme depreciations are less likely. Conversely, when the distribution exhibits thicker tails (smaller θ)—that is, when extreme depreciations are more likely—currency shares become less sensitive to relative cost advantages.

1.15.2 Implications for the Shift-share Instrument

This framework yields three implications relevant to the shift-share IV empirical design in [Section 1.4](#).

First, if one currency repeatedly offers lower debt-service costs, borrowing tends to concentrate in that currency. Currency dominance therefore increases with persistent relative cost advantages and with the likelihood of extreme exchange rate depreciation captured by the Fréchet parameter θ .

Second, currency shares depend on both own and rival debt-service costs. As a result, a modest temporary foreign monetary tightening need not necessarily generate an immediate

change in currency composition over short horizons. Currency shares are therefore persistent and relatively sticky, and the shape parameter θ governs their responsiveness to relative borrowing cost advantages.

Third, repeated draws average out transitory exchange rate fluctuations. Consequently, short-run exchange-rate volatility need not translate immediately into changes in currency composition, allowing currency shares to remain relatively stable over time.

Most importantly, the closed-form currency-share expression in [Equation \(1.23\)](#) mirrors the same shift-share logic as the default probability representation in [Equation \(1.2\)](#). This correspondence provides a microfoundation for interacting predetermined currency exposure with foreign monetary shocks in the empirical design of [Section 1.4](#).

1.16 CONTROL FUNCTION APPROACH TO ADDRESS EXCLUSION RESTRICTION

Suppose the true model is given as:

$$y = \beta D + \phi z + v \quad (1.24)$$

where y is the dependent variable, D is a binary treatment indicator, and z is the instrument for D . Assume (y, D, z) are covariance-stationary so that standard OLS asymptotic inference applies.³⁰ If the treatment variable D is not truly exogenous, then $\mathbb{E}[Dv] \neq 0$. Assume z is an exogenous instrument for D but it violates the exclusion restriction— $\mathbb{E}[zv] = 0$ but $\phi \neq 0$. Estimating a two-stage least squares (2SLS) regression:

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

will yield an inconsistent estimator because

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

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

³⁰The 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 1.24](#) strongly rejects a unit root.

sample and estimate the 2SLS regression again. This allows us to assess how the IV estimate 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 $y^{\text{ND}} = \phi^{\text{ND}} z^{\text{ND}} + v^{\text{ND}}$. Assume the instrument remains exogenous in this subsample, that is, $\mathbb{E}[z^{\text{ND}} v^{\text{ND}}] = 0$. Then the OLS estimator in the never-default subsample satisfies:

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

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, $\phi = \lambda \phi^{\text{ND}}$ with $\lambda > 0$. Since $\hat{\phi}^{\text{ND}} \xrightarrow{p} \phi^{\text{ND}}$, we have $\lambda \hat{\phi}^{\text{ND}} \xrightarrow{p} \lambda \phi^{\text{ND}} = \phi$. Subtracting $\lambda \hat{\phi}^{\text{ND}} z$ from the dependent variable y in the full sample, we have:

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

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

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

since the term $\left(\phi - \lambda \hat{\phi}^{\text{ND}} \right) \rightarrow 0$ 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 :

$$\begin{aligned}
\hat{\beta}_{\text{CF}}^{\text{IV}}(\lambda) &= \frac{\text{Cov}\left(y - \lambda \hat{\phi}^{\text{ND}} z, z\right)}{\text{Cov}(D, z)} \\
&= \frac{\frac{1}{N} \sum_{i=1}^N \left((y_i - \lambda \hat{\phi}^{\text{ND}} z_i) z_i \right)}{\frac{1}{N} \sum_{i=1}^N D_i z_i} \\
&\xrightarrow{p} \frac{\mathbb{E}\left[(y - \lambda \hat{\phi}^{\text{ND}} z) z\right]}{\mathbb{E}[Dz]} = \beta
\end{aligned} \tag{1.25}$$

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

Table 1.9: Data sample, availability, and comparable-country-set in the control-function robustness check.

Country	Period	IDS source	Ever defaulted	Comparable	Country	Period	IDS source	Ever defaulted	Comparable
(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
AGO	1989–2010				LSO	1970–2010			
ALB	1991–2010				LTU	1995–2010	CD2010		
ARG	1970–2010		✓	✓	LVA	1995–2010	CD2010		
ARM	1993–2010				MAR	1970–2010		✓	✓
AZE	1993–2010				MDA	1995–2010		✓	✓
BDI	1970–2010			✓	MDG	1970–2010		✓	
BEN	1970–2010			✓	MDV	1995–2010			
BFA	1970–2010		✓	✓	MEX	1970–2010		✓	✓
BGD	1972–2010			✓	MKD	1993–2010			✓
BGR	1981–2010		✓	✓	MLI	1970–2010			✓
BIH	1999–2010				MMR	1970–2010		✓	✓
BLR	1993–2010				MNG	1992–2010			✓
BLZ	1980–2010		✓	✓	MOZ	1991–2010			
BOL	1970–2010		✓	✓	MRT	1970–2010		✓	✓
BRA	1970–2010		✓	✓	MUS	1976–2010			
BRB	1974–2005	CD2006		✓	MWI	1980–2010		✓	✓
BTN	1982–2010				MYS	1970–2010	CD2010		✓
BWA	1970–2010				NER	1970–2010		✓	✓
CAF	1970–2010		✓	✓	NGA	1970–2010		✓	✓
CHL	1970–2010	CD2010	✓	✓	NIC	1970–2010		✓	✓
CHN	1981–2010				NPL	1971–2010			✓
CIV	1970–2010		✓	✓	OMN	1976–2010	CD2006		✓
CMR	1970–2010		✓	✓	PAK	1970–2010		✓	
COD	1970–2010		✓		PAN	1970–2010		✓	✓
COG	1970–2010		✓	✓	PER	1970–2010		✓	✓
COL	1970–2010			✓	PHL	1970–2010		✓	✓
COM	1980–2010			✓	PNG	1970–2010			✓
CPV	1981–2010				POL	1990–2008	CD2010		
CRI	1970–2010		✓	✓	PRY	1970–2010		✓	✓
CZE	1993–2004	CD2006		✓	ROU	1990–2010			✓
DMA	1981–2010		✓	✓	RUS	1992–2010		✓	✓
DOM	1970–2010		✓	✓	RWA	1971–2010			✓
DZA	1970–2010		✓	✓	SDN	1970–2010		✓	✓
ECU	1970–2010		✓	✓	SEN	1970–2010		✓	✓
EGY	1970–2010			✓	SLB	1983–2010		✓	✓
ERI	1994–2005			✓	SLE	1970–2010		✓	✓
EST	1995–2005	CD2006			SLV	1970–2010			✓
ETH	1981–2010		✓	✓	SOM	1973–2010			✓

Continued on next page

Table 1.9: Data sample, availability, and comparable-country-set in the control-function robustness check—continued

Country	Period	IDS source	Ever defaulted	Comparable	Country	Period	IDS source	Ever defaulted	Comparable
(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
FJI	1970–2010			✓	SRB	1999–2010			
GAB	1970–2010		✓	✓	STP	2001–2010			
GEO	1992–2010			✓	SVK	1993–2006	CD2006		
GHA	1970–2010		✓		SWZ	1970–2010			
GIN	1986–2010		✓	✓	SYC	1980–2010	CD2010	✓	✓
GMB	1970–2010		✓	✓	SYR	1970–2004			✓
GNB	1974–2010		✓	✓	TCD	1970–2010			✓
GNQ	1984–2006	CD2006			TGO	1970–2010		✓	✓
GRD	1977–2010		✓	✓	THA	1970–2010			
GTM	1970–2010		✓	✓	TJK	1992–2010			✓
GUY	1970–2010		✓	✓	TKM	1993–2010			✓
HND	1970–2010		✓	✓	TON	1988–2010			
HRV	1995–2006	CD2006		✓	TTO	1970–2005	CD2006	✓	✓
HTI	1970–2010		✓	✓	TUN	1970–2010			
HUN	1991–2006	CD2006			TUR	1970–2010		✓	✓
IDN	1970–2010		✓	✓	TZA	1988–2010			
IND	1970–2010			✓	UGA	1982–2010			
IRN	1979–2010			✓	UKR	1992–2010		✓	✓
JAM	1970–2010		✓	✓	URY	1970–2010	CD2010	✓	✓
JOR	1976–2010		✓	✓	UZB	1992–2010			✓
KAZ	1992–2010				VCT	1975–2010			
KEN	1970–2010		✓	✓	VEN	1970–2008		✓	
KGZ	1992–2010			✓	VNM	1988–2010			
KHM	1993–2010				YEM	1990–2010			✓
KNA	1984–2010	CD2010			VUT	1984–2010			✓
LAO	1984–2010				WSM	1985–2010			
LBN	1988–2010				ZAF	1994–2010			✓
LBR	2000–2010			✓	ZMB	1970–2010		✓	✓
LCA	1984–2010			✓	ZWE	1970–2008		✓	

Notes: This table reports the country sample, sample period, and data availability of the *International Debt Statistics* (IDS) used in the analysis. Blank entries in the “IDS source” column indicate that data were obtained from the World Bank online data portal. “CD2010” and “CD2006” denote data retrieved from the 2010 and 2006 editions of the *Global Development Finance* CD-ROM (the predecessor to IDS) respectively. These CD-ROMs were accessed through interlibrary loan services. A checkmark ✓ indicates that a country belongs to the corresponding category. The “Ever defaulted” columns indicate whether a sample country has defaulted at least once over the 1970–2010 period. The “Comparable” columns indicate whether a country belongs to the comparable-country-sample used in the robustness check of the control-function specification based on propensity-score matching between defaulting and never-defaulting countries. The baseline control-function specification in panel (c) of [Figure 1.1](#) uses the full sample of 135 countries and partitions countries into defaulting and never-defaulting groups. The robustness exercise reported in [Figure 1.1](#) restricts the sample to countries with similar observable characteristics, including real GDP per capita growth, inflation, exchange rate changes, the share of external debt to GDP, capital-account openness, banking, currency, and political crises, democracy indicators, and incomplete shares. The resulting comparable-country-set contains 92 countries, including 55 defaulting and 37 never-defaulting countries.

Table 1.10: Descriptive statistics for variables used in the regression analysis.

	(1)	(2)	(3)	(4)	(5)	(6)
	Mean	SD	Min	Median	Max	#obs
<i>(a) Treatment-related variables</i>						
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
$\Delta \ln\left(1 + \frac{\text{arrear}}{\text{debt}}\right)$	-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
<i>(b) Continuous variables</i>						
$\Delta \ln(\text{rgdppc})$	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
<i>(c) Binary variables</i>						
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 1.11: Balance test between defaulters and never-defaulters

	(1) Defaulters	(2) Never-defaulters	(3) Difference	(4) p-value
<i>(a) Treatment-related variables</i>				
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
$\Delta \ln(1 + \frac{\text{arrear}}{\text{debt}})$	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
<i>(b) Continuous variables</i>				
$\Delta \ln(\text{rgdppc})$	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
<i>(c) Binary variables</i>				
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 and 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 1.12: Number of sovereign defaults relative to U.S. monetary tightening episodes from [Romer and Romer \(2023\)](#)

Years relative to U.S. tightening (1)	Number of defaults (2)	Share of defaults (101) (3)	Share of sample years (41) (4)	Excess concentration (Columns (3)–(4)) (5)
-1	9	8.91%	12.20%	−3.28%pp
0	16	15.84%	12.20%	+3.65%pp
1	32	31.68%	21.95%	+9.73%pp
2	53	52.48%	31.71%	+20.77%pp
3	58	57.43%	41.46%	+15.96%pp
4	61	60.40%	48.78%	+11.62%pp

Notes: This table presents a descriptive summary of the concentration of sovereign defaults around U.S. monetary tightening episodes—consisting of 5 contractionary shocks and 1 expansionary shock during the 1970–2010 sample period—identified in [Romer and Romer \(2023\)](#). The table reports the number and share of sovereign default observations occurring within different windows relative to U.S. monetary tightening episodes, together with the share of sample years covered by those windows. Column (5) reports an “excess concentration” measure, defined as the percentage point difference between the share of sovereign defaults and the share of years over the entire sample used in [Table 1.3](#). It suggests that sovereign defaults are disproportionately concentrated following U.S. monetary tightening episodes and tend to materialize with a lag, with the strongest concentration occurring within two years after a tightening episode. Results for easing episodes are not reported because only one narrative easing episode occurs during the sample period.

Table 1.13: Argentina’s bilateral “other-currency” share (selected years and creditors)

Counterpart (1)	Indicator (2)	1970 (3)	1980 (4)	1990 (5)	2000 (6)	2010 (7)	2020 (8)
<i>(a) “Other-currency” to all creditors (world aggregate)</i>							
WLD	d_{it}	5,893.2	27,322.5	62,477.6	150,062.9	126,642.4	275,474.9
	$s_{it}^{OTHC,all}$	4.7%	3.5%	1.0%	6.3%	1.6%	2.8%
<i>(b) “Other-currency” to selected advanced economies (bilateral)</i>							
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%
	$s_{ijt}^{OTHC,all}$				4.1%	0.2%	1.1%
AUS	d_{ijt}	0	—†	0	—†	0.6	0.1
	s_{ijt}^{OTHC}	—¶	—¶	—¶	—¶	100%	—§
	$s_{ijt}^{OTHC,all}$					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%
	$s_{ijt}^{OTHC,all}$	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%
	$s_{ijt}^{OTHC,all}$			0.1%	0.0%	0.0%	0.0%
<i>(c) “Other-currency” to selected emerging markets (bilateral)</i>							
BRA	d_{ijt}	0.4	17.4	151.6	46.2	457.4	239.3
	s_{ijt}^{OTHC}	0	0	—§	0	0	13.1%
	$s_{ijt}^{OTHC,all}$						0.0%
CHN	d_{ijt}	—†	—†	—†	—†	25	22,928
	s_{ijt}^{OTHC}	—¶	—¶	—¶	—¶	0	5.6%
	$s_{ijt}^{OTHC,all}$					0.0%	0.5%

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). $s_{it}^{OTHC,all}$ 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 $s_{ijt}^{OTHC,all}$ is the “bilateral” OTHC share between i and j against the total external debt of all creditors, that is $s_{ijt}^{OTHC,all} = \frac{s_{ijt}^{OTHC} \times d_{ijt}}{d_{it}}$. Inferred IDS’s coding convention: —†: data entries are missing (either blank or “.”); —¶: IDS codes OTHC (“other-currenc share”) as 100, but often then the bilateral external debt 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 1.14: Comparison between linear and logistic regression on the default probability.

Dependent variable: S&P first-year binary (0/1) default indicator			
	Linear (1)	Logistic (2)	Linear (same sample as (2)) (3)
ΔIRE_{it-1}	0.009 (0.006)	0.308 (0.191)	0.014 (0.010)
ΔIRE_{it-2}	0.026*** (0.006)	0.815*** (0.213)	0.040*** (0.009)
ΔIRE_{it-3}	0.010* (0.006)	0.166 (0.218)	0.014 (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 Panel (a) of [Table 1.3](#), and column (2) reports logit estimates with the same set of baseline controls and country fixed effects described in [Section 1.4.2](#). 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 1.15: Amplification of $\Delta\text{IRE}_{it-\ell}$ through floating rate liabilities.

Dependent variable: S&P first-year binary (0/1) default indicator				
	(1)	(2)	(3)	(4)
sharevarrate _{it-1}	0.0341 (0.0223)	0.0524** (0.0245)	0.0440* (0.0244)	0.0587** (0.0259)
ΔIRE_{it-1}	0.0143* (0.0084)			0.0143 (0.0090)
$\Delta\text{IRE}_{it-1} \times \text{sharevarrate}_{it-1}$	-0.0290 (0.0278)			-0.0202 (0.0277)
ΔIRE_{it-2}		0.0163** (0.0068)		0.0155** (0.0066)
$\Delta\text{IRE}_{it-2} \times \text{sharevarrate}_{it-1}$		0.0470** (0.0193)		0.0471** (0.0183)
ΔIRE_{it-3}			-0.0002 (0.0061)	0.0003 (0.0067)
$\Delta\text{IRE}_{it-3} \times \text{sharevarrate}_{it-1}$			0.0489** (0.0241)	0.0453* (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 $\Delta\text{IRE}_{it-\ell}$ 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—IBOR or the U.S. prime linked interest rates—out of total external debt stock. The main finding is that, conditional on $\Delta\text{IRE}_{it-\ell}$, 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 [Section 1.4.2](#). 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 1.16: Longer debt maturities mitigate first-stage effect on $\Delta\text{IRE}_{it-\ell}$.

Dependent variable: S&P first-year binary (0/1) default indicator				
	(1)	(2)	(3)	(4)
Maturity _{it-1}	-0.0010** (0.0004)	-0.0011*** (0.0004)	-0.0011*** (0.0004)	-0.0011*** (0.0004)
ΔIRE_{it-1}	-0.0050 (0.0147)			-0.0055 (0.0150)
$\Delta\text{IRE}_{it-1} \times \text{Maturity}_{it-1}$	0.0005 (0.0005)			0.0005 (0.0006)
ΔIRE_{it-2}		0.0551*** (0.0133)		0.0537*** (0.0132)
$\Delta\text{IRE}_{it-2} \times \text{Maturity}_{it-1}$		-0.0013*** (0.0005)		-0.0013** (0.0005)
ΔIRE_{it-3}			0.0308** (0.0152)	0.0239 (0.0148)
$\Delta\text{IRE}_{it-3} \times \text{Maturity}_{it-1}$			-0.0010* (0.0006)	-0.0007 (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 $\Delta\text{IRE}_{it-\ell}$ 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 $\Delta\text{IRE}_{it-\ell}$, 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 Section 1.4.2. 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 1.17: Robustness check using lagged shares in the shift-share instrument.(a): First-stage results with lagged currency shares ($\Delta\text{IREls}_{it-\ell}$)

Dependent variable: S&P first-year binary (0/1) default indicator				
	(1)	(2)	(3)	(4)
$\Delta\text{IREls}_{it-1}$	0.008 (0.006)			0.009 (0.006)
$\Delta\text{IREls}_{it-2}$		0.026*** (0.006)		0.025*** (0.005)
$\Delta\text{IREls}_{it-3}$			0.010* (0.006)	0.009 (0.006)

(b) Two-stage least square (2SLS) results with lagged currency shares ($\Delta\text{IREls}_{it-2}$)

Horizons (h)	Output response ($y_{it+h} - y_{it-1}$)		AR test	OLS=IV
	LP-OLS	LP-SSIV (lagged shares)	p-value	p-value
	(1)	(2)	(3)	(4)
$h = 0$	-2.65*** (0.63)	-5.98 (5.40)	0.27	0.55
$h = 1$	-3.79*** (1.01)	-16.11** (7.92)	0.04**	0.13
$h = 2$	-4.75*** (1.14)	-13.51 (8.42)	0.11	0.30
$h = 3$	-4.78*** (1.33)	-15.66* (9.36)	0.10*	0.25
$h = 4$	-5.27*** (1.51)	-5.76 (10.79)	0.60	0.96
$h = 5$	-5.19*** (1.74)	-1.27 (10.57)	0.91	0.72
$h = 6$	-5.15*** (1.81)	3.30 (11.40)	0.77	0.47
Joint significance	0.00	0.33		
Instrument		$\Delta\text{IREls}_{it-2}$		
KP weak IV		20.64		
Observations		3,898		

Notes: This table presents the robustness check of constructing the shift-share instrument with lagged currency shares. To be specific, I define $\Delta\text{IREls}_{it} \equiv \sum_j \text{currencyshare}_{it-1}^j \times \Delta\text{interestrates}_t^j$, which differs from Equation (1.5) in the lagged exposure term $\text{currencyshare}_{it-1}^j$ (instead of contemporaneous $\text{currencyshare}_{it}^j$ in the original ΔIRE_{it} specification). It also reports p-values from the joint significance test across all horizons (i.e., $\beta_0 = \dots = \beta_6 = 0$), the Kleibergen-Paap weak instrument (F-statistic), the Anderson-Rubin weak instrument test (F-statistic), and OLS=IV coefficient equality tests (stacking method) at each horizon h . All columns include (i) country fixed effects, (ii) the baseline control set $\mathbf{X}_{it}$ defined in Section 1.4.2, and (iii) the same 3,898 observations. Robust standard errors clustered at the country level are reported in parentheses. Statistical significance is indicated as *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table 1.18: Robustness check using short-term interest rates in shift-share instrument.
(a): First-stage results with short-term interest rates ($\Delta\text{IREstir}_{it}$)

Dependent variable: S&P first-year binary (0/1) default indicator				
	(1)	(2)	(3)	(4)
$\Delta\text{IREstir}_{it-1}$	0.000 (0.003)			-0.000 (0.004)
$\Delta\text{IREstir}_{it-2}$		0.013*** (0.002)		0.012*** (0.003)
$\Delta\text{IREstir}_{it-3}$			0.008*** (0.002)	0.006** (0.003)
(b) Two-stage least square (2SLS) results with short-term interest rates ($\Delta\text{IREstir}_{it-2}$)				
	Output response ($y_{it+h} - y_{it-1}$)		AR test	OLS=IV
Horizons (h)	LP-OLS (1)	LP-SSIV (ST interest rate) (2)	p-value (3)	p-value (4)
$h = 0$	-2.64*** (0.63)	-12.42* (6.75)	0.05*	0.15
$h = 1$	-3.71*** (1.01)	-23.08** (10.20)	0.02**	0.06*
$h = 2$	-4.62*** (1.13)	-31.09*** (10.23)	0.00***	0.01**
$h = 3$	-4.57*** (1.32)	-34.74*** (10.64)	0.00***	0.01***
$h = 4$	-5.17*** (1.50)	-33.36*** (10.64)	0.00***	0.01***
$h = 5$	-5.15*** (1.73)	-5.67 (9.00)	0.53	0.95
$h = 6$	-5.13*** (1.79)	12.82 (9.88)	0.16	0.07*
Joint significance	0.00	0.01***		
Instrument		$\Delta\text{IREstir}_{it-2}$		
KP weak IV		22.24		
Observations		3,898		

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 $\Delta\text{IREstir}_{it} \equiv \sum_j \text{currencyshare}_{it}^j \times \Delta\text{STinterestrate}_t^j$, which differs from Equation (1.5) 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). It also reports p-values from the joint significance test across all horizons (i.e., $\beta_0 = \dots = \beta_6 = 0$), the Kleibergen-Paap weak instrument (F-statistic), the Anderson-Rubin weak instrument test (F-statistic), and OLS=IV coefficient equality tests (stacking method) at each horizon h . All columns include (i) country fixed effects, (ii) the baseline control set $\mathbf{X}_{it}$ defined in Section 1.4.2, and (iii) the same 3,898 observations. Robust standard errors clustered at the country level are reported in parentheses. Statistical significance is indicated as *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table 1.19: Robustness check of excluding U.S. dollar debt in shift-share instrument (leave-one-out).

(a): First-stage results with excluding U.S. dollar debt ($\Delta\text{IREnoUSD}_{it}$)				
Dependent variable: S&P first-year binary (0/1) default indicator				
	(1)	(2)	(3)	(4)
$\Delta\text{IREnoUSD}_{it-1}$	0.022 (0.014)			0.021 (0.014)
$\Delta\text{IREnoUSD}_{it-2}$		0.025** (0.011)		0.018* (0.009)
$\Delta\text{IREnoUSD}_{it-3}$			0.017* (0.010)	0.016* (0.009)
(b) 2SLS results with excluding U.S. dollar debt ($\Delta\text{IREnoUSD}_{it-2}$)				
	Output response ($y_{it+h} - y_{it-1}$)		AR test	OLS=IV
Horizons (h)	LP-OLS	LP-SSIV (exclude USD)	p-value	p-value
	(1)	(2)	(3)	(4)
$h = 0$	-2.65*** (0.62)	-1.55 (15.38)	0.92	0.59
$h = 1$	-3.86*** (0.99)	-9.22 (26.05)	0.73	0.55
$h = 2$	-4.90*** (1.11)	22.36 (36.85)	0.49	0.31
Joint significance	0.00	0.01***		
Instrument	$\Delta\text{IREnoUSD}_{it-2}$			
KP weak IV	22.24			
Observations	3.91			

Notes: This table presents the robustness check that constructs the shift-share instrument while excluding the U.S. dollar denomination debt. I define $\Delta\text{IREnoUSD}_{it} \equiv \sum_{j \neq \text{US\$}} \text{currencyshare}_{it}^j \times \Delta\text{interestrates}_t^j$, which differs from Equation (1.5) by dropping the USD-weighted term $\text{currencyshare}_{it}^{\text{US\$}} \times \Delta\text{interestrates}_t^{\text{US\$}}$. To be specific, I define $\Delta\text{IREstir}_{it} \equiv \sum_j \text{currencyshare}_{it}^j \times \Delta\text{STinterestrates}_t^j$, which differs from Equation (1.5) in that it uses the short-term interest rates $\Delta\text{STinterestrates}_t^j$ (instead of long-term $\Delta\text{interestrates}_t^j$ in the original ΔIRE_{it} specification). It also reports p-values from the joint significance test across all horizons (i.e., $\beta_0 = \dots = \beta_6 = 0$), the Kleibergen-Paap weak instrument (F-statistic), the Anderson-Rubin weak instrument test (F-statistic), and OLS=IV coefficient equality tests (stacking method) at each horizon h . All columns include (i) country fixed effects, (ii) the baseline control set $\mathbf{X}_{it}$ defined in Section 1.4.2, and (iii) the same 3,898 observations. Robust standard errors clustered at the country level are reported in parentheses. Statistical significance is indicated as *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table 1.20: Robustness check using average currency shares in shift-share instrument.(a): First-stage results with average currency shares ($\Delta\text{IREavgshare}_{it}$)

Dependent variable: S&P first-year binary (0/1) default indicator				
	(1)	(2)	(3)	(4)
$\Delta\text{IREavgshare}_{it-1}$	0.007 (0.006)			0.008 (0.006)
$\Delta\text{IREavgshare}_{it-2}$		0.027*** (0.005)		0.026*** (0.005)
$\Delta\text{IREavgshare}_{it-3}$			0.011* (0.006)	0.010* (0.006)

(b) Two-stage least square (2SLS) results with average currency shares ($\Delta\text{IREavgshare}_{it-2}$)

Horizons (h)	Output response ($y_{it+h} - y_{it-1}$)		AR test	OLS=IV
	LP-OLS	LP-SSIV (average shares)	p-value	p-value
	(1)	(2)	(3)	(4)
$h = 0$	-2.64*** (0.63)	-5.73 (4.96)	0.25	0.55
$h = 1$	-3.80*** (1.01)	-16.90** (7.60)	0.02**	0.09*
$h = 2$	-4.76*** (1.14)	-16.57** (7.50)	0.02**	0.12
$h = 3$	-4.77*** (1.33)	-18.42** (8.72)	0.03**	0.12
$h = 4$	-5.28*** (1.51)	-9.93 (10.50)	0.34	0.66
$h = 5$	-5.20*** (1.74)	-7.49 (10.72)	0.48	0.83
$h = 6$	-5.15*** (1.81)	0.62 (11.72)	0.96	0.63
Joint significance	0.00	0.131		
Instrument	$\Delta\text{IREavgshare}_{it-2}$			
KP weak IV	23.26			
Observations	3,898			

Notes: This table presents the robustness check that constructs the shift-share instrument with average currency shares across all years for each country. To be specific, I define $\Delta\text{IREavgshare}_{it} \equiv \sum_j \text{avgshare}_i^j \times \Delta\text{interestrates}_t^j$, where $\text{avgshare}_i^j \equiv \frac{1}{T} \sum_t \text{currencyshare}_{it}^j$ is country-specific but time-invariant. It also reports p-values from the joint significance test across all horizons (i.e., $\beta_0 = \dots = \beta_6 = 0$), the Kleibergen-Paap weak instrument (F-statistic), the Anderson-Rubin weak instrument test (F-statistic), and OLS=IV coefficient equality tests (stacking method) at each horizon h . All columns include (i) country fixed effects, (ii) the baseline control set $\mathbf{X}_{it}$ defined in Section 1.4.2, except the incomplete shares—since the average shares are country-specific but time-invariant and hence will be absorbed by country fixed effects, and (iii) the same 3,898 observations. Robust standard errors clustered at the country level are reported in parentheses. Statistical significance is indicated as *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table 1.21: Robustness check using first-year currency shares in shift-share instrument.(a): First-stage results with first-year currency shares ($\Delta\text{IRE1styrshare}_{it}$)

Dependent variable: S&P first-year binary (0/1) default indicator				
	(1)	(2)	(3)	(4)
$\Delta\text{IRE1styrshare}_{it-1}$	0.008 (0.007)			0.008 (0.007)
$\Delta\text{IRE1styrshare}_{it-2}$		0.024*** (0.005)		0.023*** (0.005)
$\Delta\text{IRE1styrshare}_{it-3}$			0.009 (0.006)	0.007 (0.006)

(b) Two-stage least square (2SLS) results with first-year currency shares ($\Delta\text{IRE1styrshare}_{it-2}$)

Horizons (h)	Output response ($y_{it+h} - y_{it-1}$)		AR test	OLS=IV
	LP-OLS	LP-SSIV (first-year shares)	p-value	p-value
	(1)	(2)	(3)	(4)
$h = 0$	-2.64*** (0.63)	-4.09 (5.87)	0.48	0.81
$h = 1$	-3.80*** (1.00)	-12.64 (9.01)	0.15	0.34
$h = 2$	-4.81*** (1.13)	-11.48 (9.61)	0.23	0.49
$h = 3$	-4.82*** (1.32)	-3.71 (10.12)	0.72	0.91
$h = 4$	-5.34*** (1.51)	10.23 (11.91)	0.36	0.20
Instrument	$\Delta\text{IRE1styrshare}_{it-2}$			
KP weak IV	18.30			
Observations	3,898			

Notes: This table presents the robustness check that constructs the shift-share instrument with first-year currency shares across all years for each country. To be specific, I define $\Delta\text{IRE1styrshare}_{it} \equiv \sum_j 1\text{styrshare}_i^j \times \Delta\text{interestrates}_t^j$, where 1styrshare_i^j is the corresponding currency share values in the first year when the IDS data become available, and it is country-specific but time-invariant. It also reports p-values from the joint significance test across all horizons (i.e., $\beta_0 = \dots = \beta_6 = 0$), the Kleibergen-Paap weak instrument (F-statistic), the Anderson-Rubin weak instrument test (F-statistic), and OLS=IV coefficient equality tests (stacking method) at each horizon h . All columns include (i) country fixed effects, (ii) the baseline control set $\mathbf{X}_{it}$ defined in Section 1.4.2, except the incomplete shares—since the first-year shares are country-specific but time-invariant and hence will be absorbed by country fixed effects, and (iii) the same 3,898 observations. Robust standard errors clustered at the country level are reported in parentheses. Statistical significance is indicated as *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table 1.22: Robustness check: Excluding observations related to the 1980s Latin American Debt Crisis.

Dependent variable: cumulative loss in real GDP per capita (%) LP-SSIV				
Horizons (h)	Baseline	Drop LAC	Drop LAC in 1982–1989	Drop 1982–1989
	(1)	(2)	(3)	(4)
$h = 0$	-7.99 (5.28)	-6.49 (8.85)	-6.81 (7.69)	24.59 (30.16)
$h = 1$	-18.48** (8.08)	-24.28* (14.14)	-25.09** (12.57)	-49.91 (46.08)
$h = 2$	-15.78* (8.60)	-13.93 (13.13)	-17.57 (12.35)	-90.87 (71.78)
$h = 3$	-18.25* (9.52)	-12.96 (14.91)	-19.46 (13.68)	-73.45 (65.07)
$h = 4$	-9.47 (10.72)	-2.03 (17.65)	-4.40 (15.66)	63.93 (72.88)
$h = 5$	-5.70 (10.55)	6.40 (17.40)	4.92 (15.58)	130.40 (106.13)
$h = 6$	-0.52 (11.17)	21.24 (19.32)	15.68 (16.64)	140.53 (111.74)
Instrument	ΔIRE_{it-2}			
KP weak IV	21.14	10.33	12.49	2.20
Observations	3,898	2,867	3,676	3,177

Notes: This table examines whether the baseline LP-SSIV estimates are primarily driven by the 1980s Latin American Debt Crisis. Column (1) reproduces the baseline LP-SSIV estimates reported in Table 1.3. Column (2) excludes all 29 Latin American and Caribbean (LAC) countries from the baseline sample. Column (3) excludes observations for these LAC countries during the 1982–1989 debt crisis period only, while Column (4) excludes all country-year observations during the 1982–1989 period for all countries. This table additionally reports the Kleibergen–Paap weak instrument (F-statistic). All columns include (i) country fixed effects and (ii) the baseline control set $\mathbf{X}_{it}$ defined in Section 1.4.2. The number of observations differs across columns because of sample differences. Since Column (4) exhibits substantially weaker first-stage strength, its estimates should be interpreted with caution. Robust standard errors clustered at the country level are reported in parentheses. Statistical significance is indicated as *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table 1.23: Robustness check using alternative binary default and debt distress indicators from [Laeven and Valencia \(2020\)](#).

Dependent variable: cumulative loss in real GDP per capita (%)						
Independent variable	S&P (Baseline)		Laeven and Valencia (2020)			
			Debt crisis		Debt distress	
Horizons (h)	LP-SSIV (1)	AR test (2)	LP-SSIV (3)	AR test (4)	LP-SSIV (5)	AR test (6)
$h = 0$	-7.99 (5.28)	0.12	-9.86 (6.30)	0.11	-10.38 (6.85)	0.11
$h = 1$	-18.48** (8.08)	0.01**	-22.02** (9.24)	0.01**	-23.44** (10.31)	0.01**
$h = 2$	-15.78* (8.60)	0.06*	-21.73** (10.05)	0.03**	-22.69** (10.79)	0.03**
$h = 3$	-18.25* (9.52)	0.05*	-25.98** (11.40)	0.02**	-27.48** (12.50)	0.02**
$h = 4$	-9.47 (10.72)	0.38	-13.74 (12.82)	0.29	-14.57 (13.62)	0.28
$h = 5$	-5.70 (10.55)	0.59	-9.12 (12.54)	0.47	-9.49 (13.22)	0.48
$h = 6$	-0.52 (11.17)	0.96	-3.44 (13.34)	0.80	-3.65 (14.08)	0.80
Instrument	ΔIRE_{it-2}					
KP weak IV	21.14		32.21		16.31	
Observations	3,898		3,688		3,688	

Notes: This table presents the robustness check using alternative default and debt distress indicators from [Laeven and Valencia \(2020\)](#). To be specific, debt distress is defined as a country experiencing either a debt crisis or debt restructuring as classified in [Laeven and Valencia \(2020\)](#). Columns (1)–(2) reproduce the baseline LP-SSIV results as in [Table 1.3](#). Columns (3)–(6) use the debt crisis classifications from [Laeven and Valencia \(2020\)](#). Panel (c) of [Figure 1.1](#) visualizes the LP-SSIV estimates using the debt crisis indicators as reported in columns (3)–(4). Columns (5)–(6) combine the debt crisis and debt restructuring binary indicators into a broader debt distress measure. This table also reports the Kleiberg–Paap weak instrument (F-statistic) and p-values of the Anderson–Rubin weak instrument test (F-statistic) at each horizon h . All columns include (i) country fixed effects and (ii) the baseline control set $\mathbf{X}_{it}$ defined in [Section 1.4.2](#). The number of observations differs because the S&P and [Laeven and Valencia \(2020\)](#) classifications define default episodes differently. Robust standard errors clustered at the country level are reported in parentheses. Statistical significance is indicated as *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table 1.24: Persistence and unit-root tests for the U.S. dollar share and $\Delta\text{IRE}_{it-\ell}$

Country	U.S. dollar share			$\Delta\text{IRE}_{it-\ell}$			Country	U.S. dollar share			$\Delta\text{IRE}_{it-\ell}$		
	ρ (1)	DF (2)	p-value (3)	ρ (3)	DF (4)	p-value (4)		ρ (5)	DF (6)	p-value (7)	ρ (7)	DF (8)	p-value (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***		

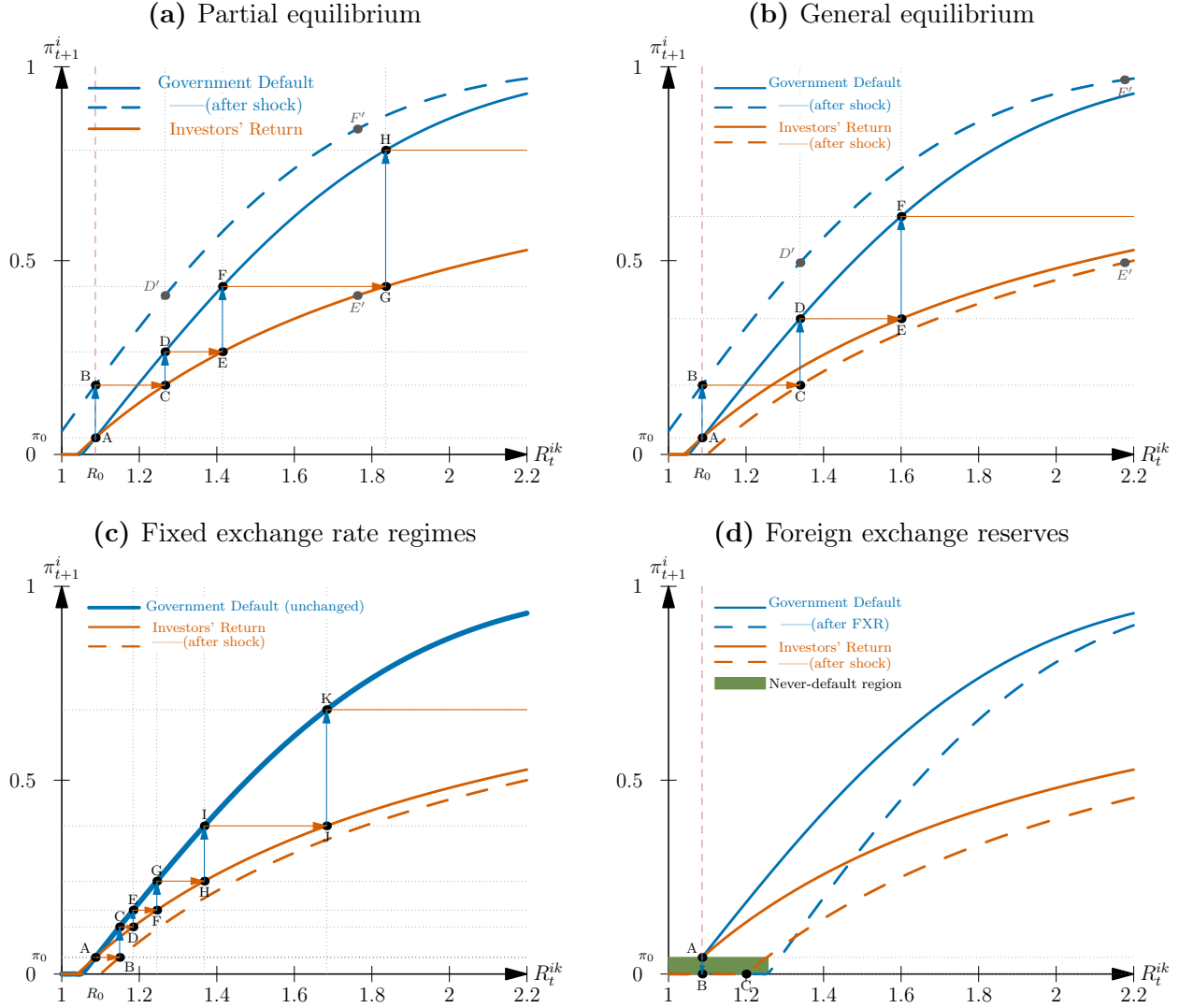
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Table 1.24: Persistence and unit-root tests for the U.S. dollar share and $\Delta\text{IRE}_{it-\ell}$ —continued

Country	U.S. dollar share			$\Delta\text{IRE}_{it-\ell}$			U.S. dollar share			$\Delta\text{IRE}_{it-\ell}$		
	ρ	DF	p-value	ρ	DF	p-value	ρ	DF	p-value	ρ	DF	p-value
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
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***	

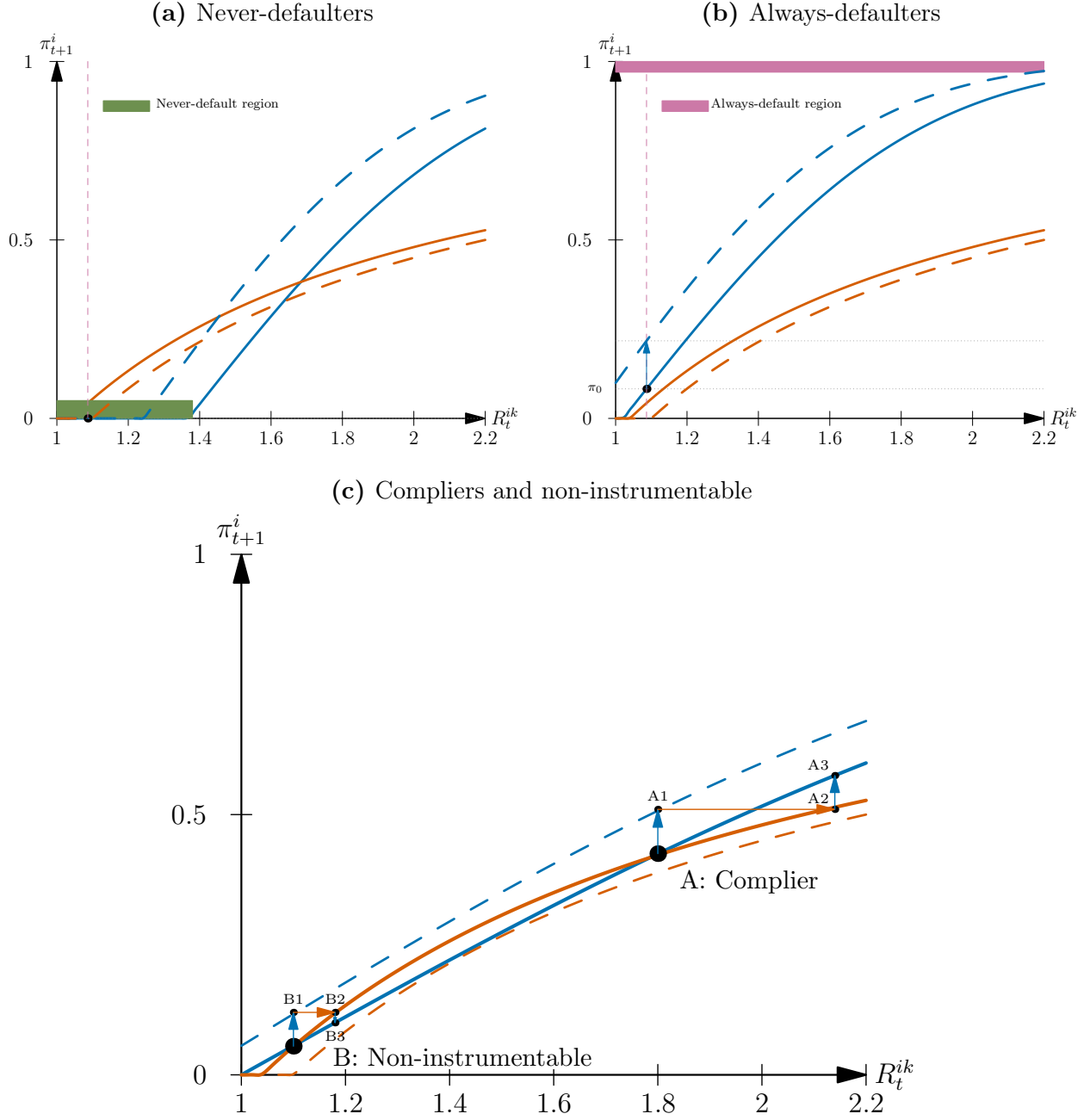
Notes: This table reports persistence and unit-root test results for both the U.S. dollar currency share and the shift-share instrument ΔIRE_{it} estimated separately for each country. The AR(1) coefficient ρ is derived from the OLS regression of each series on its first lag. The “DF p-value” is the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

Figure 1.4: Multiple equilibria in the probabilistic sovereign default model.



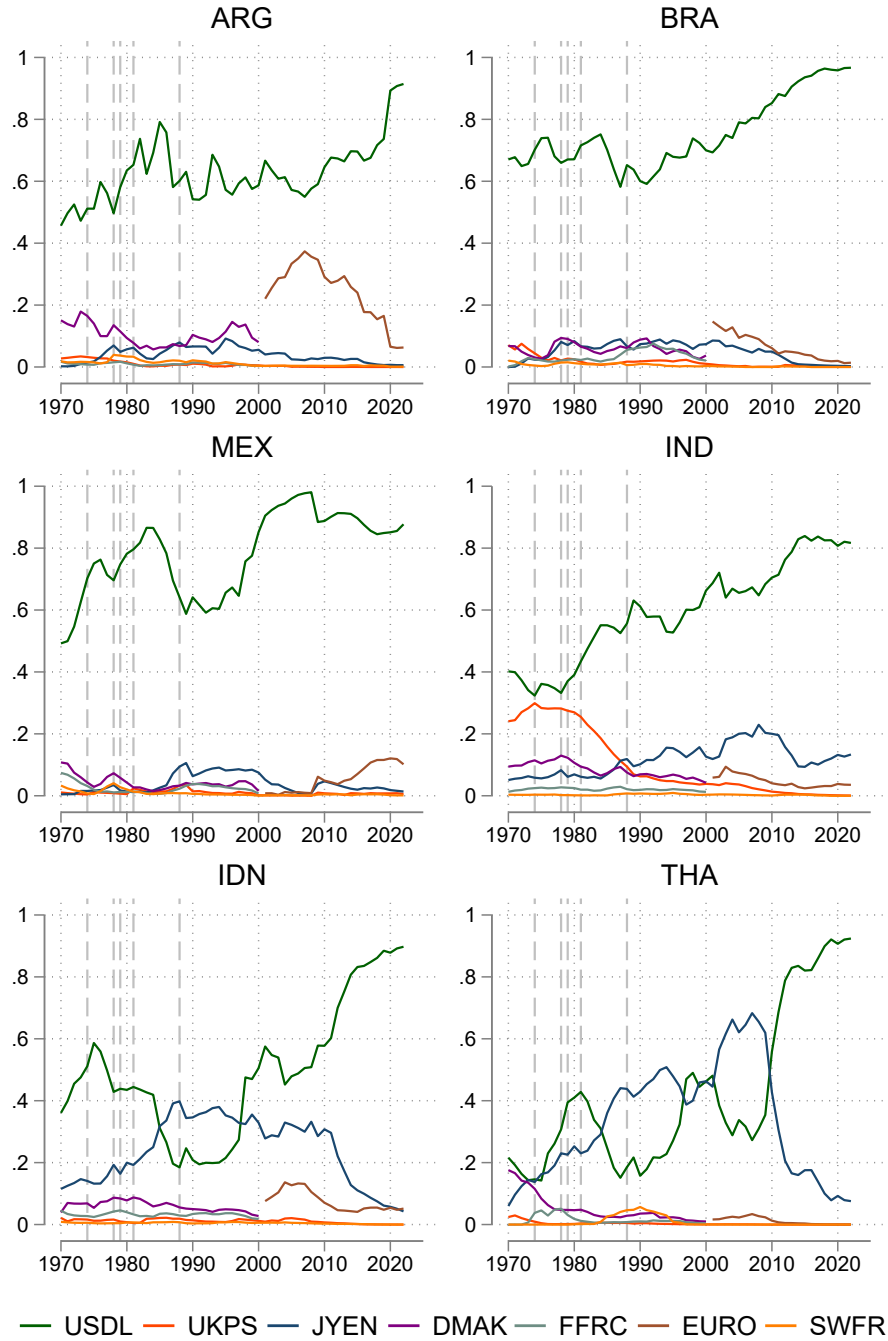
Notes: This figure illustrates multiple equilibria in the probabilistic sovereign default from Section 1.2. The vertical axis is the default probability; the horizontal axis is the gross rollover interest factor. “Shock” denotes a foreign monetary tightening; solid curves are pre-shock and dashed curves 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 1.5: Multiple equilibria and local average treatment effect (LATE).



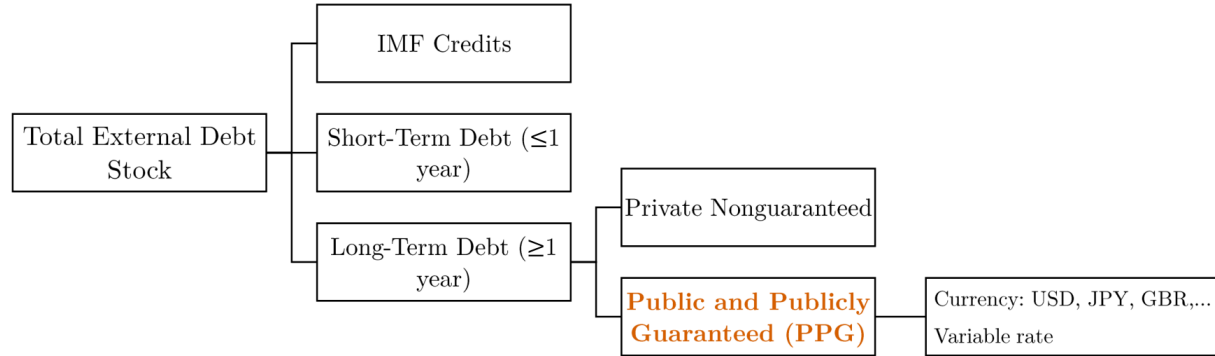
Notes: This figure classifies initial steady states in the probabilistic sovereign default model in Section 1.2 into four scenarios: (a) never-defaulters: intersections lie on the horizontal axis pre- and post-shock; (b) always-defaulters: intersections lie on the vertical axis pre- and post-shock; (c) 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 1.6: Overall trends in currency denomination for selected countries.



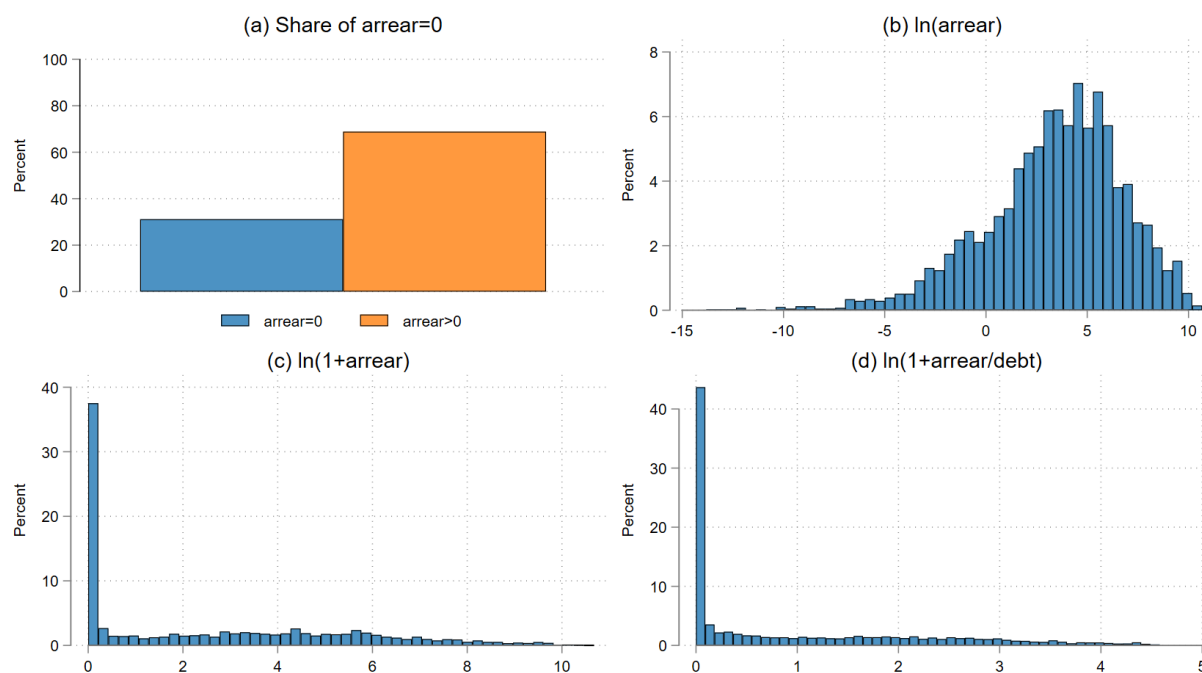
Notes: This figure visualizes the currency denomination trends of selected countries over time. The data are based on the public and publicly guaranteed debt from the World Bank's *International Debt Statistics* (IDS). The currencies shown are USDL (U.S. dollar), UKPS (pound sterling), JYEN (Japanese 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 1.7: Data Structure of the World Banks' International Debt Statistics



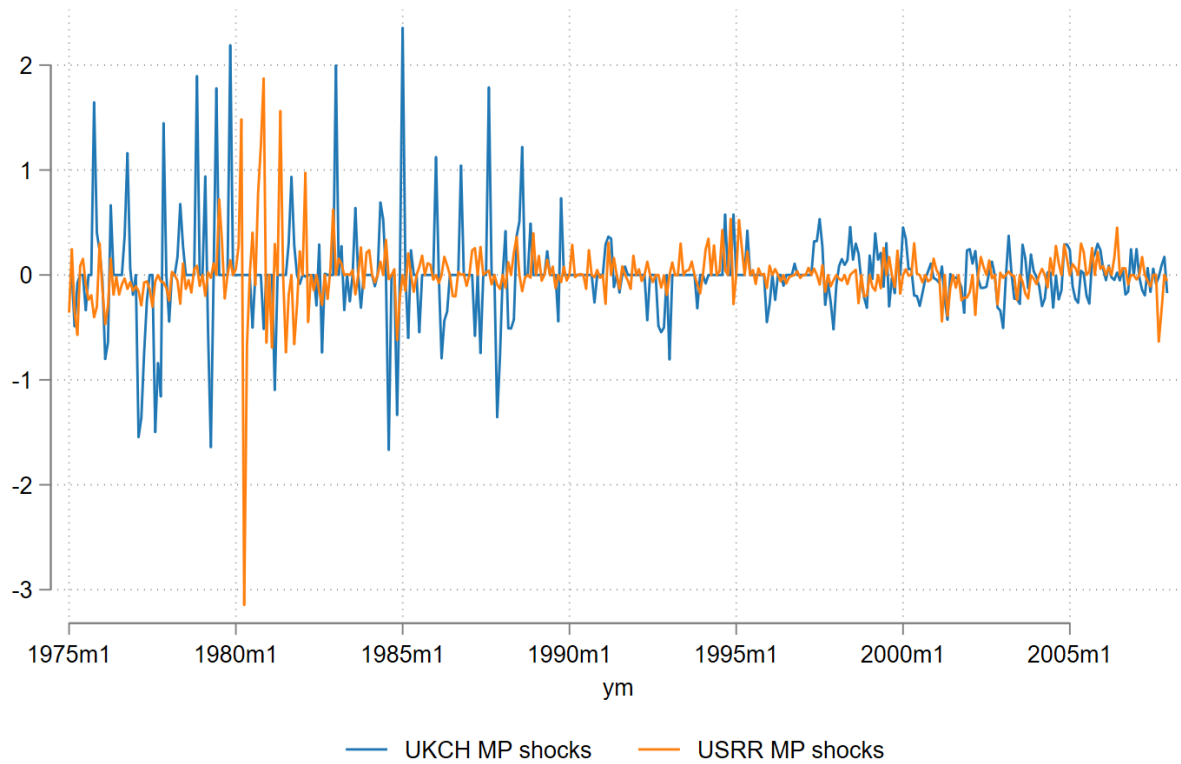
Notes: This figure summarizes the overall data structure of the World Bank's International Debt Statistics relevant to this paper. The currency denomination data are only available for long-term public and publicly guaranteed (PPG) debt. The currency denomination data are not available for short-term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

Figure 1.8: Distribution of sovereign arrears (late repayments) for 1970–2020



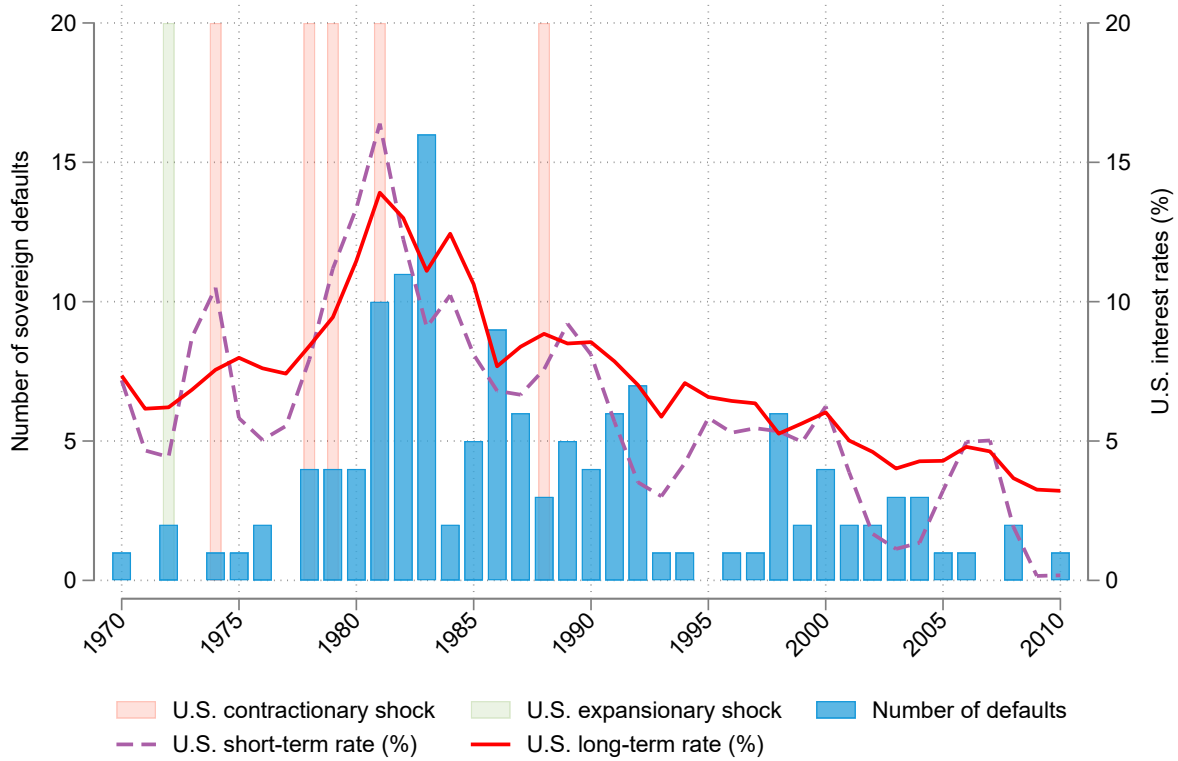
Notes: This figure shows the distribution of sovereign arrears from the World Bank’s International Debt Statistics (IDS) over the 1970–2020 period. The data include the sum of principal and interest arrears for each country vis-à-vis all creditors (i.e., the world). All monetary values are expressed 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 1.9: Narrative monetary shocks from Cloyne and Hurtgen (2016)



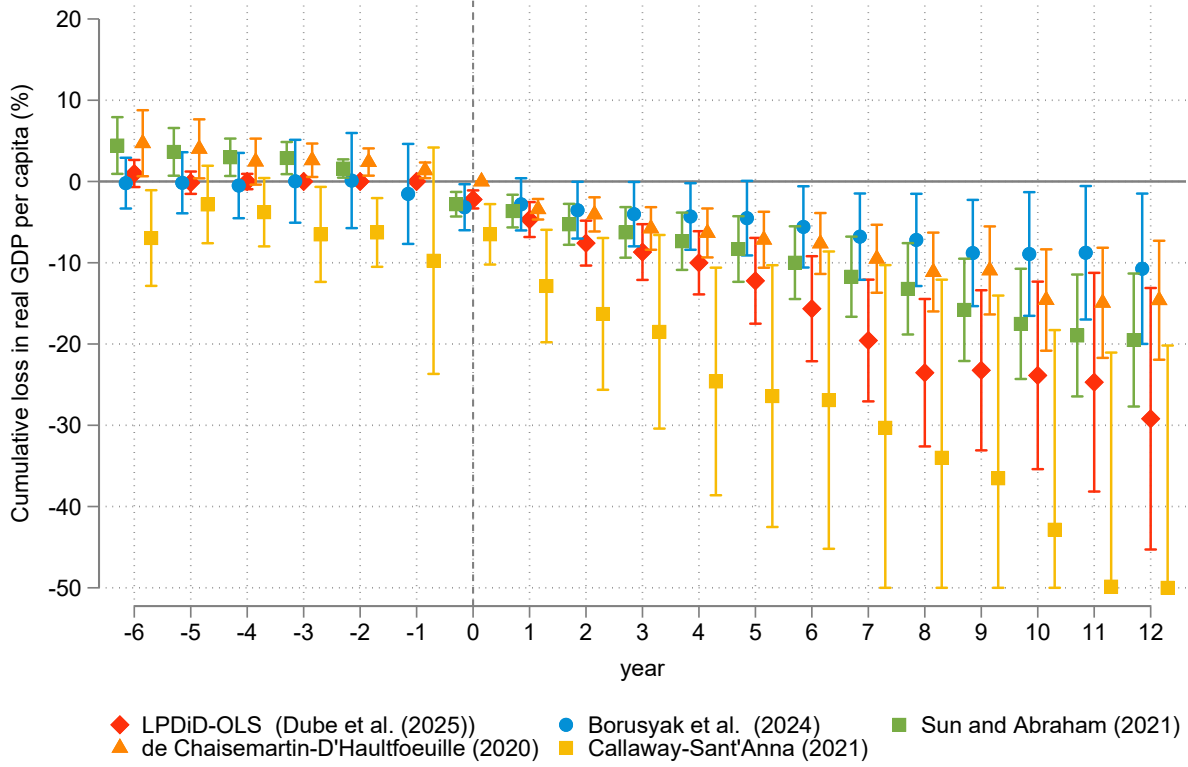
Notes: This figure plots monthly narrative monetary shocks series from Cloyne and Hurtgen (2016). Notice that the U.K. monetary shocks series is markedly more volatile and exhibits larger shocks in magnitude than its U.S. counterpart. Also, the within-year positives and negatives monthly monetary shocks often offset each other. To mitigate this attenuation, Table 1.7 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 1.10: Close relationship between U.S. monetary cycles and sovereign default waves.



Notes: This figure plots the annual number of sovereign defaults (left axis) together with U.S. short-term and long-term interest rates (right axis). Together with Table 1.12, it highlights a close relationship between U.S. monetary policy cycles and sovereign default waves, particularly during the late 1970s and early 1980s. The shaded red (green) areas indicate narrative U.S. contractionary (expansionary) monetary shocks identified by Romer and Romer (2023). While descriptive rather than causal, the strong co-movement underscores an important identification challenge for specifications with year fixed effects: global interest-rate cycles coincide with waves of sovereign default, implying that year fixed effects absorb much of the common variation in both sovereign default indicators and the shift-share instrument. By contrast, a difference-in-differences (DiD) framework can help isolate cross-country variation in sovereign default decisions while flexibly controlling for common time shocks through year fixed effects. Section 1.6.4 presents a DiD application that delivers qualitatively similar results. Sovereign default is measured using the binary S&P default indicator from Kuvshinov and Zimmermann (2019), while U.S. short-term and long-term interest rates are obtained from the *JST Macroeconomy Database*.

Figure 1.11: All difference-in-differenceS 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 1.3, it additionally reports results from Callaway and Sant'Anna (2021). All DiD specifications include country and year fixed effects, as well as the baseline controls specified in Table 1.3. 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.

Chapter 2

Evaluating Emerging Markets' Monetary Policy

2.1 INTRODUCTION

Are inflation-target regimes in emerging markets credible? Since the early 2000s, central banks around the world have increasingly adopted monetary frameworks centered on price stability, often with explicit numerical inflation targets. While advanced economies typically operate under clear legal mandates and strong institutional independence, emerging markets face structurally different constraints. They are more exposed to external shocks, rely more heavily on export-led growth, carry substantial foreign currency denominated debt, and often operate under political pressure. These features raise doubts about whether formal inflation targets meaningfully constrain policy in practice.

When difficult trade-offs arise, do central banks in emerging markets honor their stated commitment to price stability, or do they systematically deviate in favor of other objectives, such as supporting output or stabilizing the exchange rate? Such deviations matter because policy credibility shapes expectation formation, which in turn affects monetary transmission through both conventional tools (e.g., interest rates) and unconventional tools (e.g., forward guidance), especially near the zero lower bound.

This paper builds on the sufficient statistics framework of [Barnichon and Mesters \(2023\)](#) to evaluate the credibility of inflation-target regimes in emerging markets. They show that, given (i) empirical impulse responses of macroeconomic variables to a monetary policy shock and (ii) forecasts of future economic paths, one can construct a sufficient statistic—the “Optimal Policy Perturbation (OPP)” —that assesses whether the observed policy rate is locally optimal. Intuitively, if no small deviation in the policy rate can improve the central bank’s loss function under expected future outcomes, the observed policy rate is internally consistent and therefore credible.

The logic of the OPP framework is closely related to first-order optimality conditions. In a single-variable setting, optimality requires that the derivative of the objective function with respect to the policy instrument equals zero. In a multivariate setting, local optimality requires the gradient of the objective function to be zero. If a small perturbation to the policy rate could move the gradient closer to zero, then the observed policy rate is not locally optimal. The magnitude and sign of this perturbation summarize both the direction and strength of the policy adjustment that would further reduce the central bank’s welfare loss.

A central contribution of this paper is to exploit the *dual interpretation* of the OPP statistic. Under the assumption that policymakers are cautious and generally avoid unnecessary policy changes, observed policy rates can be treated as approximately optimal. I therefore adopt a revealed-preference approach to recover the implied policy priorities: assuming observed policy rates are optimal, I solve for the weight vector on output, inflation, and depreciation that rationalizes the observed policy rate at each point in time by minimizing the OPP deviation from zero. If a central bank claims to prioritize inflation stabilization, yet the implied policy weights place most emphasis on exchange rate stability, this constitutes strong empirical evidence against credibility.

This approach is particularly well suited to emerging markets, where policy objectives are rarely explicit or stable. Unlike the Federal Reserve’s dual mandate—often stylized as equal weights on inflation and employment—there is no clear benchmark loss function.

Historical episodes such as the 1997 Asian Financial Crisis and the 2013 Taper Tantrum suggest that exchange-rate stabilization is often an implicit objective. I therefore define monetary policy as credible if observed policy rates can be rationalized by forecasted outcomes under a stable weighting of output, inflation, and depreciation over time.

Operationally, I develop an algorithm that recovers the implicit policy weights consistent with the observed policy rate in each period. While period-by-period estimates are naturally noisy, the implied weights exhibit systematic and economically meaningful patterns when examined across time and across regimes—for example, before and after inflation-target adoption or during financial crisis episodes.

The OPP framework also admits a complementary interpretation. Instead of recovering the implicit policy weights, one can impose some fixed weight vectors representing externally imposed benchmarks, such as an inflation-target benchmark $(\frac{1}{2}, \frac{1}{2}, 0)$, an equal-weight benchmark $(\frac{1}{3}, \frac{1}{3}, \frac{1}{3})$, or a global-optimal benchmark defined as the weight vector that minimizes the sum of squared OPP deviations from zero across all periods. Under each benchmark, the OPP statistic identifies the direction and magnitude of the policy rate adjustment that would improve central bank’s welfare. When different benchmarks agree on the sign of the OPP, this provides robust evidence that observed policy was not locally optimal. When benchmarks disagree, this reveals the underlying trade-offs faced by the central bank. Reassuringly, the main credibility conclusions in this paper are robust across these alternative benchmarks.

The empirical implementation proceeds in two steps. First, I use a Bayesian autoregression (Bayesian VAR) to estimate impulse responses of output, inflation, and depreciation to a monetary policy shock for twelve inflation-target emerging economies. Unlike local projections, which estimate each horizon independently, the Bayesian VAR delivers joint posterior distributions of impulse response estimates across horizons, allowing coherent inference for OPP statistics over time. Moreover, since high-frequency monetary policy shocks are largely unavailable for emerging markets, I use a recursive (Cholesky) identification scheme that orders the policy rate last.

Second, I construct expected future paths using forecasts from the Economist Intelligence Unit (EIU). Unlike [Barnichon and Mesters \(2023\)](#), who rely on internal Federal Reserve forecasts, comparable official projections are unavailable for most emerging markets. The EIU forecasts offer three advantages: broad cross-country coverage since the mid-1990s, consistent methodology across countries, and institutional independence from domestic political pressures.

The results provide several new insights. I document strong evidence that many emerging markets have gradually placed increasing and more stable weight on inflation stabilization over time. This pattern is particularly pronounced in countries such as Brazil and Mexico, which experienced chronic high inflation in earlier decades. During crisis episodes, some central banks temporarily shift weight toward output stabilization, but many revert quickly to inflation-focused behavior, suggesting flexibility rather than abandonment of the inflation-target framework during distressed times.

The framework also highlights underperformance. For example, for South Korea, all benchmarks indicate that policy rates should have been lowered more aggressively during the 2008 Global Financial Crisis. For Turkey, all benchmarks consistently imply that policy rates should have been tighter, especially since the early 2020s, coinciding with rising political interference. In small open economies such as Israel, I find evidence of a shift toward output stabilization following the Israel–Gaza conflict, illustrating how geopolitical risks shape monetary trade-offs.

The approach has limitations. First, the EIU forecasts are not equivalent to central banks' internal projections, though they provide consistent cross-country coverage and may be less politically distorted than official data in some settings. Second, policy instruments are also measured imperfectly across countries, as many emerging markets target collateralized or longer-maturity rates rather than overnight interbank rates, and operational frameworks have evolved over time. These issues introduce noise but do not undermine the central contribution: providing a transparent, comparable, and theory-consistent measure of monetary policy

credibility where strong institutional benchmarks are absent.

2.1.1 Related Literature

This paper builds on the literature evaluating optimal monetary policy and credibility, building directly on the OPP framework of [Barnichon and Mesters \(2023\)](#). Its central contribution is extending this framework to emerging markets, where policy objectives are often opaque and institutional constraints are more severe.

A large literature emphasizes that emerging markets systematically deviate from textbook inflation-target behavior. [Calvo and Reinhart \(2002\)](#) document widespread “fear of floating”: low exchange-rate volatility combined with high volatility in interest rates and reserves. [Fraga, Goldfajn and Minella \(2003\)](#) argue that emerging markets face greater macroeconomic volatility, weaker credibility, and larger external exposure, necessitating more flexible inflation-target frameworks. Reviews by [Svensson \(2010\)](#) and [Gill Hammond \(2012\)](#) emphasize credibility, transparency, and flexibility as central pillars of successful inflation-target regimes.

On the theoretical side, [Svensson \(1997\)](#) shows that inflation-forecast targeting dominates both money growth and exchange rate targeting. [Ghosh, Ostry and Chamon \(2016\)](#) argue that sterilized foreign exchange intervention can be compatible with inflation-target and may even enhance credibility when capital mobility is imperfect.

The empirical evidence on inflation-target remains mixed. [Mishkin and Schmidt-Hebbel \(2007\)](#) find that inflation-target countries achieve lower long-run inflation and reduced sensitivity to external shocks. By contrast, [Zhang and Wang \(2022\)](#) and [Bhalla, Bhasin and Loungani \(2023\)](#) find limited short-run macroeconomic improvements following adoption and cannot reject mean reversion. Critics argue that inflation-target may come at the expense of output or financial stability, though [Walsh \(2009\)](#) finds no increase in output volatility and suggests that emerging markets may benefit more than advanced economies.

Rather than evaluating the average effectiveness of inflation-target, this paper focuses

on the *credibility*. Credibility matters because expectations play a central role in inflation persistence and monetary transmission, especially as forward guidance becomes more important (Bems et al., 2021; Baxa, Horváth and Vašíček, 2014). The closest paper is Uribe et al. (2025), who show that post-adoption behavior is path-dependent: countries with histories of high inflation respond more aggressively to deviations even under formally credible regimes.

The remainder of the paper is organized as follows. Section 2.2 presents the methodology. Section 2.3 describes the empirical framework. Section 2.4 introduces the data. Section 2.5 reports the results. Section 2.6 concludes.

2.2 METHODOLOGY

This section presents the empirical framework used to construct optimal policy perturbation (OPP) statistics. The OPP statistic requires two sufficient statistics: (i) $\mathcal{R}^0$, the impulse responses of macroeconomic variables to a monetary policy shock under the baseline policy rule ϕ^0 ; and (ii) $\mathcal{Y}_t^0$, the forecast paths of these variables under the baseline rule ϕ^0 in the absence of a policy perturbation (i.e., $\delta_t = 0$).

Section 2.2.1 motivates the OPP statistic in a simple open-economy New Keynesian model. Section 2.2.3 explains how I use EIU forecasts to proxy forecast path $\mathcal{Y}_t^0$. Sections 2.2.4 describe the Bayesian VAR, the recursive (Cholesky) identification scheme, and the construction of the impulse responses $\mathcal{R}^0$. Section 2.2.5 combines these objects to compute the OPP time series and credible intervals.

2.2.1 Motivation: A simple four-equation New Keynesian model

To motivate the OPP statistic, consider a discretionary New Keynesian (NK) model in a small open economy setting (Galí, 2015; Barnichon and Mesters, 2023). The model consists of an investment-saving (IS) curve, a New Keynesian Phillips curve (NKPC), an uncovered

interest parity (UIP) condition, and a baseline Taylor-style monetary policy (MP) rule.¹ The system is given as follows:

$$\begin{aligned}
(\text{IS}) \quad & x_t = \mathbb{E}_t[x_{t+1}] - \frac{1}{\sigma}(i_t - \mathbb{E}_t[\pi_{t+1}]) \\
(\text{NKPC}) \quad & \pi_t = \mathbb{E}_t[\pi_{t+1}] + \kappa x_t + \gamma \Delta \varepsilon_t + \xi_t \\
(\text{UIP}) \quad & \Delta \varepsilon_t = i_t + \eta_t \\
(\text{MP}^0) \quad & i_t^0 = \phi^0 \pi_t + \epsilon_t^{\text{MP}} \\
(\text{MP-OPP}) \quad & i_t = i_t^0 + \delta_t = \phi^0 \pi_t + \epsilon_t^{\text{MP}} + \delta_t
\end{aligned}$$

where x_t is the output gap, π_t is inflation, $\Delta \varepsilon_t$ is the currency depreciation rate, and i_t is the monetary policy rate. The shocks $(\xi_t, \eta_t, \epsilon_t^{\text{MP}})$ are a cost-push shock, a UIP wedge shock, and a monetary policy shock respectively. The parameter ϕ^0 governs the baseline monetary policy rule. Relative to the textbook three-equation NK model, the key difference is the UIP wedge η_t and its transmission to inflation through the NKPC.²

The OPP statistic δ_t represents an infinitesimal deviation from the baseline policy rate. It is introduced to examine whether a marginal adjustment in the policy instrument would reduce the central bank's loss. At a local optimum, the gradient of the central bank's loss function with respect to the policy instrument—evaluated at the observed policy rate—should equal zero (a necessary but not sufficient condition for optimality). Equivalently, if a small disturbance δ_t can move this gradient closer to zero, then the observed policy rate is not locally optimal. In particular, the sign of δ_t indicates the direction of the welfare-improving policy

¹The seminal paper by [Gali and Monacelli \(2005\)](#) shows that under certain conditions (e.g., log utility, unit elasticity of substitution between home and foreign good, and producer currency pricing), the optimal monetary policy of a small open economy is the domestic inflation targeting rule without an explicit exchange rate term. Moreover, [Taylor \(2001\)](#) argues that including the exchange rate directly in a Taylor rule need not improve macro stabilization because exchange rate movements already affect inflation dynamics (e.g., through the $\gamma \Delta \varepsilon_t$ term in the NKPC), and an explicit exchange rate term in the monetary policy rule may induce excessive interest rate volatility.

²A standard UIP condition is $\mathbb{E}_t[\Delta \varepsilon_{t+1}] = i_t - i_t^* + \zeta_t$, where ζ_t is a wedge that reconciles the empirical deviations from the UIP condition. For a small open economy, I treat the foreign interest rate i_t^* and the UIP wedge ζ_t as exogenous and absorb $(-i_t^* + \zeta_t)$ into a single exogenous shock η_t .

adjustment, and its magnitude summarizes how strongly the data call for that adjustment.

Under discretionary monetary policy, $\mathbb{E}_t[x_{t+1}]$ and $\mathbb{E}_t[\pi_{t+1}]$ are taken as given. In this simple environment, the effect of δ_t on contemporaneous macro outcomes admits a closed-form solution—the endogenous variables $(x_t, \pi_t, \Delta\varepsilon_t, i_t)$ can be solved as linear functions of the exogenous shocks $(\xi_t, \eta_t, \epsilon_t^{\text{MP}})$ and the policy perturbation δ_t :

$$\begin{aligned} x_t &= -\frac{\phi^0}{\sigma\Lambda^0}\xi_t - \frac{\gamma\phi^0}{\sigma\Lambda^0}\eta_t - \frac{1}{\sigma\Lambda^0}\epsilon_t^{\text{MP}} - \frac{1}{\sigma\Lambda^0}\delta_t \\ \pi_t &= \frac{1}{\Lambda^0}\xi_t + \frac{\gamma}{\Lambda^0}\eta_t + \frac{\gamma - \kappa/\sigma}{\Lambda^0}\epsilon_t^{\text{MP}} + \frac{\gamma - \kappa/\sigma}{\Lambda^0}\delta_t \\ \Delta\varepsilon_t &= \frac{\phi^0}{\Lambda^0}\xi_t + \frac{1 + (\kappa/\sigma)\phi^0}{\Lambda^0}\eta_t + \frac{1}{\Lambda^0}\epsilon_t^{\text{MP}} + \frac{1}{\Lambda^0}\delta_t \\ i_t &= \frac{\phi^0}{\Lambda^0}\xi_t + \frac{\gamma\phi^0}{\Lambda^0}\eta_t + \frac{1}{\Lambda^0}\epsilon_t^{\text{MP}} + \frac{1}{\Lambda^0}\delta_t \end{aligned}$$

where $\Lambda^0 \equiv 1 - (\gamma - \kappa/\sigma)\phi^0$ is a function that depends on the baseline rule ϕ^0 and other structural parameters.

Let $\mathbf{y}_t \equiv (x_t, \pi_t, \Delta\varepsilon_t)' \in \mathbb{R}^3$ collect the endogenous variables. Define the coefficient vectors $\Gamma^0 \equiv \frac{1}{\Lambda^0} \left(-\frac{\phi^0}{\sigma}, 1, \phi^0 \right)'$, $\Psi^0 \equiv \frac{1}{\Lambda^0} \left(-\frac{\gamma\phi^0}{\sigma}, \gamma, 1 + \frac{\kappa\phi^0}{\sigma} \right)'$, and $\mathcal{R}^0 \equiv \frac{1}{\Lambda^0} \left(-\frac{1}{\sigma}, \gamma - \frac{\kappa}{\sigma}, 1 \right)'$, where $\Gamma^0, \Psi^0, \mathcal{R}^0 \in \mathbb{R}^3$. Then the above equilibrium can be written compactly as

$$\mathbf{y}_t = \underbrace{\Gamma^0\xi_t + \Psi^0\eta_t + \mathcal{R}^0\epsilon_t^{\text{MP}}}_{\equiv \mathcal{Y}_t^0} + \mathcal{R}^0\delta_t \quad (2.1)$$

Equation (2.1) decomposes outcomes into two components: (i) $\mathcal{Y}_t^0$, the baseline forecast path of macroeconomic outcomes implied by the baseline rule ϕ^0 and the realized shocks $(\xi_t, \eta_t, \epsilon_t^{\text{MP}})$; and (ii) $\mathcal{R}^0\delta_t$, the deviation from that baseline forecast path induced by an arbitrary policy perturbation δ_t added to the baseline monetary rule.

Intuitively, in the absence of cost-push and exchange-rate shocks (i.e., $\xi_t = 0$ and $\eta_t = 0$), the term $\mathcal{R}^0\epsilon_t^{\text{MP}}$ in $\mathcal{Y}_t^0$ represents the contemporaneous impulse responses of macroeconomic outcomes $\mathbf{y}_t$ to a one-unit monetary policy shock ϵ_t^{MP} under the baseline rule ϕ^0 . By contrast, $\mathcal{R}^0\delta_t$ captures the marginal effect of an infinitesimal policy perturbation δ_t on

the macroeconomic outcomes.

Most importantly, both the monetary policy shock ϵ_t^{MP} and the policy perturbation δ_t affect macroeconomic outcomes $\mathbf{y}_t$ through the same coefficient vector $\mathcal{R}^0$. This equivalence underpins the sufficient statistics approach: empirical impulse responses to a monetary policy shock can be used to characterize how a marginal policy perturbation shifts macroeconomic outcomes, and therefore how it alters the central bank's loss function.

2.2.2 Central bank loss function and optimal policy perturbation

This subsection shows how the policy perturbation δ_t can be used to assess whether observed policy is locally optimal, or equivalently, whether there exists scope for improving policy implementation. I assume the central bank operates under a flexible inflation-target regime and evaluates outcomes using a standard quadratic loss function over the output gap, inflation, and exchange rate depreciation (Svensson, 2010):

$$\mathcal{L}_t \equiv \frac{1}{2} \mathbf{y}_t' \mathbf{W} \mathbf{y}_t = \frac{1}{2} (\omega^x x_t^2 + \omega^\pi \pi_t^2 + \omega^\varepsilon \Delta \varepsilon_t^2),$$

where $\mathbf{W} \equiv \text{diag}(\omega^x, \omega^\pi, \omega^\varepsilon) \in \mathbb{R}^{3 \times 3}$, $\omega^x + \omega^\pi + \omega^\varepsilon = 1$, and $\omega^x, \omega^\pi, \omega^\varepsilon \geq 0$.

Following the delegation logic of Rogoff (1985), I introduce a social planner who evaluates whether the observed policy rate is internally consistent with the central bank's stated objectives. The social planner asks whether a small perturbation δ_t to the policy rule could further reduce the central bank's loss. Using $\mathbf{y}_t = \mathcal{Y}_t^0 + \mathcal{R}^0 \delta_t$ from Equation 2.1, the social planner's problem is

$$\min_{\delta_t \in \mathbb{R}} \mathcal{L}_t = \frac{1}{2} \mathbf{y}_t' \mathbf{W} \mathbf{y}_t \quad \text{s.t.} \quad \mathbf{y}_t = \mathcal{Y}_t^0 + \mathcal{R}^0 \delta_t.$$

The first-order condition yields the (scalar) optimal policy perturbation

$$\delta_t^* = -(\mathcal{R}^{0'} \mathbf{W} \mathcal{R}^0)^{-1} (\mathcal{R}^{0'} \mathbf{W} \mathcal{Y}_t^0), \quad \delta_t^* \in \mathbb{R}. \quad (2.2)$$

Consistent with the discussion in Section 2.2.1, Equation (2.2) shows that computing the OPP statistic requires two inputs: (i) the baseline forecast path $\mathcal{Y}_t^0$ implied by the baseline policy rule under full information about all realized shocks; and (ii) the impulse response $\mathcal{R}^0$, which maps both a monetary policy shock and a marginal policy perturbation into macroeconomic outcomes.

The derivation above focuses on the contemporaneous horizon ($h = 0$). In practice, however, monetary policy affects macroeconomic outcomes over multiple future horizons, and central banks care about the entire forecast path rather than only contemporaneous outcomes.

Accordingly, I generalize (2.2) to a multi-horizon setting by stacking forecast paths and impulse responses of macroeconomic variables $n = 1, 2, \dots, N$ over horizons $h = 0, 1, \dots, H$, so that $\mathcal{Y}_t^0 \in \mathbb{R}^{N(H+1)}$ and $\mathcal{R}^0 \in \mathbb{R}^{N(H+1)}$. The simple New Keynesian example above corresponds to the special case $N = 3$ and $H = 0$, while the empirical implementation in Sections 2.2.3 and 2.2.4 allows both dimensions to be unrestricted.

2.2.3 EIU forecasts data as proxies for the counterfactual path $\mathcal{Y}_t^0$

Sections 2.2.1 and 2.2.2 both interpret the variable $\mathcal{Y}_t^0$ as the expected evolution of macroeconomic outcomes under the baseline policy rule ϕ^0 , given full information about all realized shocks $(\xi_t, \eta_t, \epsilon_t^{\text{MP}})$ and in the absence of any policy perturbation (i.e., $\delta_t = 0$).³ Under this interpretation, $\mathcal{Y}_t^0$ represents the counterfactual path that would prevail if the central bank implemented its announced policy rule and private agents formed expectations using all available information about relevant exogenous shocks.

I proxy this object $\mathcal{Y}_t^0$ using professional forecasts, interpreted as expectations formed using comprehensive contemporaneous information about macroeconomic conditions, external shocks, and the prevailing baseline policy rule (e.g., inflation-target). Under this interpretation, the reported EIU forecast paths summarize the expected trajectory of the

³Importantly, the baseline rule need not be optimal.

economy under the baseline policy rule, in the absence of policy perturbations.⁴

EIU reports provide annual (calendar-year; end-of-year) forecasts for output, inflation, depreciation, and other macroeconomic variables up to five years ahead. These forecasts are observed at discrete horizons $h \in \{0, 4, 8, 12, 16, 20\}$ quarters. I therefore set the maximum horizon to $H = 20$ in both the Bayesian VAR and the forecast data and linearly interpolate to obtain quarterly forecasts at all horizons $h = 0, 1, \dots, 20$.

To be specific, let $y_{t,h}^{0,n,\text{int}}$ denote the interpolated scalar forecast for variable n at horizon h , made at year-quarter t .⁵ I construct deviations from the long-run forecast $y_{t,H+1}^{0,n,\text{int}}$ and define $\tilde{y}_{t,h}^{0,n} \equiv y_{t,h}^{0,n,\text{int}} - y_{t,H+1}^{0,n}$. Stacking these deviations across all N variables and all $H + 1$ horizons yields the vector $\tilde{\mathcal{Y}}_t^0 \in \mathbb{R}^{N(H+1)}$ at year-quarter t :

$$\tilde{\mathcal{Y}}_t^0 \equiv \begin{pmatrix} \tilde{\mathbf{y}}_{t,0}^0 \\ \tilde{\mathbf{y}}_{t,1}^0 \\ \vdots \\ \tilde{\mathbf{y}}_{t,H}^0 \end{pmatrix} \in \mathbb{R}^{N(H+1)}, \quad \tilde{\mathbf{y}}_{t,h}^0 \equiv \begin{pmatrix} y_{t,h}^{0,1,\text{int}} - y_{t,H+1}^{0,1} \\ y_{t,h}^{0,2,\text{int}} - y_{t,H+1}^{0,2} \\ \vdots \\ y_{t,h}^{0,N,\text{int}} - y_{t,H+1}^{0,N} \end{pmatrix} \in \mathbb{R}^N.$$

To incorporate forecast uncertainty, I follow [Barnichon and Mesters \(2023\)](#) and model the baseline forecast path as a random vector centered on the observed forecast path. To be specific, I assume

$$\mathcal{Y}_t^{0,(m)} \sim \mathcal{N}\left(\tilde{\mathcal{Y}}_t^0, \Omega \otimes I_N\right), \quad m = 1, \dots, M$$

where $\mathcal{Y}_t^{0,(m)}$ represents the empirical counterpart to the theoretical object $\mathcal{Y}_t^0$ in each draw.

The matrix $\Omega \in \mathbb{R}^{(H+1) \times (H+1)}$ captures correlation across horizons and is specified as

$$\Omega = [\rho^{|k-j|}]_{k,j=0}^H, \quad \rho = 0.9.$$

⁴[Barnichon and Mesters \(2023\)](#) use the Federal Reserve's internal Greenbook forecasts, which they interpret as reflecting policymakers' information set under the baseline rule. Comparable internal forecasts are not publicly available for central banks in emerging markets.

⁵In particular, $y_{t,4}^{0,n,\text{int}}$, $y_{t,8}^{0,n,\text{int}}$, $y_{t,12}^{0,n,\text{int}}$, $y_{t,16}^{0,n,\text{int}}$, and $y_{t,20}^{0,n,\text{int}}$ correspond to the published EIU forecasts at horizons $h = 4, 8, 12, 16, 20$.

This AR(1)-type structure captures the idea that forecast errors at nearby horizons are more highly correlated than those at distant horizons. The simulated draws $\{\mathcal{Y}_t^{0,(m)}\}_{m=1}^M$ are used to propagate forecast uncertainty into the distribution of the OPP statistic.

2.2.4 Using Bayesian VAR to estimate impulse responses $\mathcal{R}^0$

This subsection describes the Bayesian VAR (BVAR) framework used to estimate the impulse responses $\mathcal{R}^0$ of macroeconomic variables to a one-unit monetary policy shock under the baseline policy rule ϕ^0 . The formulation closely follows [Koop \(2010\)](#) and [Geweke \(2005\)](#). Section 2.2.4.1 defines the reduced-form VAR and its stacked regression representation. Section 2.2.4.2 introduces the Normal-Inverse-Wishart prior and posterior sampling. Section 2.2.4.3 details the recursive identification scheme. Section 2.2.4.4 constructs impulse responses using the companion-form representation.

Macroeconomic variables often interact both contemporaneously and with lags. A VAR therefore provides a flexible reduced-form representation of these joint dynamics without requiring a fully specified structural model (e.g., explicit assumptions about expectation formation or deep parameters). Under standard regularity conditions, the reduced-form lag coefficients can be consistently estimated. To map reduced-form innovations into economically meaningful monetary policy shocks, I impose a Recursive (Cholesky) identification that orders the policy rate last, corresponding to the assumption that the policy rate affects other macroeconomic variables only with a lag within the period.

I estimate the VAR in a Bayesian framework to obtain a joint posterior distribution over impulse responses across horizons. This joint posterior preserves the variance-covariance structure of impulse response estimates across horizons, which is important for coherently propagating uncertainty into the OPP statistic.⁶

⁶I adopt a Bayesian VAR rather than local projections because local projections estimate each horizon separately and therefore do not preserve the cross-horizon variance-covariance structure by construction. Moreover, the Bayesian VAR posterior provides a natural way to propagate uncertainty into the OPP statistic via posterior simulation.

2.2.4.1 Notations and timing conventions

A structural interpretation of the VAR is that the variables of interest are jointly determined by both contemporaneous and lagged values of all variables in the system, including themselves. Consider an L -lag structural VAR. For variable n in period t , its value y_t^n can be written as

$$y_t^n = \sum_{i=1}^N \alpha^{ni} y_t^i + \sum_{\ell=1}^L \left(\sum_{i=1}^N \phi_\ell^{ni} y_{t-\ell}^i \right) + \epsilon_t^n,$$

where ϵ_t^n denotes the structural shock to variable n at time t .

Rewriting the system in matrix form, define $\mathbf{y}_t \equiv (y_t^1, y_t^2, \dots, y_t^N)' \in \mathbb{R}^N$ and $\boldsymbol{\epsilon}_t \equiv (\epsilon_t^1, \dots, \epsilon_t^N)' \in \mathbb{R}^N$, and let $\mathbf{A} = (\alpha^{ni}) \in \mathbb{R}^{N \times N}$ and $\Phi_\ell = (\phi_\ell^{ni}) \in \mathbb{R}^{N \times N}$. The structural VAR can then be written compactly as

$$\mathbf{A} \mathbf{y}_t = \sum_{\ell=1}^L \Phi_\ell \mathbf{y}_{t-\ell} + \boldsymbol{\epsilon}_t.$$

Assuming that $\mathbf{A}$ is invertible, the system admits the reduced-form representation

$$\mathbf{y}_t = \sum_{\ell=1}^L \mathbf{B}_\ell \mathbf{y}_{t-\ell} + \mathbf{e}_t, \quad \mathbf{B}_\ell \equiv \mathbf{A}^{-1} \Phi_\ell, \quad \mathbf{e}_t \equiv \mathbf{A}^{-1} \boldsymbol{\epsilon}_t,$$

where $\mathbf{e}_t \in \mathbb{R}^N$ and $\mathbf{B}_\ell \in \mathbb{R}^{N \times N}$. I assume $\mathbb{E}[\mathbf{e}_t] = 0$ and $\mathbb{E}[\mathbf{e}_t \mathbf{e}_t'] \equiv \boldsymbol{\Sigma} \in \mathbb{R}^{N \times N}$.

Under standard regularity conditions, the reduced-form coefficients $\{\mathbf{B}_\ell\}_{\ell=1}^L$ can be consistently estimated. However, identification of the contemporaneous impact matrix $\mathbf{A}^{-1}$ requires additional assumptions, which I impose below using a recursive (Cholesky) identification in Section 2.2.4.3.

Stacking the system over time yields the regression form

$$\mathbf{Y} = \mathbf{X} \mathbf{B} + \mathbf{e},$$

where the matrix $\mathbf{Y} \in \mathbb{R}^{(T-L) \times N}$ stacks macroeconomic outcomes; the regressor matrix

$\mathbf{X} \equiv (\mathbf{1} \mid \mathbf{Y}_1 \mid \cdots \mid \mathbf{Y}_L) \in \mathbb{R}^{(T-L) \times (1+NL)}$ stacks a row of ones $\mathbf{1}' \in \mathbb{R}^{1 \times N}$ and L lags of $\mathbf{Y}_\ell$; the coefficient matrix $\mathbf{B} \in \mathbb{R}^{(1+NL) \times N}$ stacks the intercept and lag matrices $\mathbf{B}_\ell \in \mathbb{R}^{N \times N}$; and the matrix $\mathbf{e} \in \mathbb{R}^{(T-L) \times N}$ stacks the reduced-form innovations. To be specific, their matrix forms are given by:

$$\mathbf{Y} \equiv \mathbf{Y}_{L+1}, \quad \mathbf{Y}_\ell \equiv \begin{pmatrix} \mathbf{y}'_\ell \\ \mathbf{y}'_{\ell+1} \\ \vdots \\ \mathbf{y}'_{\ell+(T-L-1)} \end{pmatrix}, \quad \mathbf{B} = \begin{pmatrix} \mathbf{C}_0 \\ \mathbf{B}_1 \\ \vdots \\ \mathbf{B}_L \end{pmatrix}, \quad \mathbf{e} \equiv \begin{pmatrix} \mathbf{e}'_{L+1} \\ \mathbf{e}'_{L+2} \\ \vdots \\ \mathbf{e}'_T \end{pmatrix}$$

where $\mathbf{y}'_t = (y_t^1, y_t^2, \dots, y_t^N) \in \mathbb{R}^{1 \times N}$, $\mathbf{e}'_t \equiv (e_t^1, e_t^2, \dots, e_t^N) \in \mathbb{R}^{1 \times N}$, and $\mathbf{X}$ is given by:

$$\mathbf{X} \equiv \begin{pmatrix} 1 & y_1^1 & \cdots & y_1^N & \cdots & y_L^1 & \cdots & y_L^N \\ \vdots & \vdots & \ddots & \vdots & & \vdots & \ddots & \vdots \\ 1 & y_{T-L}^1 & \cdots & y_{T-L}^N & \cdots & y_{T-1}^1 & \cdots & y_{T-1}^N \end{pmatrix} \in \mathbb{R}^{(T-L) \times (1+NL)}$$

At this stage, I only impose that the reduced-form innovations satisfy $\mathbb{E}[\mathbf{e}_t] = 0$ and $\mathbb{E}[\mathbf{e}_t \mathbf{e}_t'] = \mathbf{\Sigma}$, where $\mathbf{\Sigma} \in \mathbb{R}^{N \times N}$ is the symmetric positive-definite variance-covariance matrix

$$\mathbf{\Sigma} \equiv \begin{pmatrix} \mathbb{E}[(e_t^1)^2] & \cdots & \mathbb{E}[e_t^1 e_t^N] \\ \vdots & \ddots & \vdots \\ \mathbb{E}[e_t^N e_t^1] & \cdots & \mathbb{E}[(e_t^N)^2] \end{pmatrix} \in \mathbb{R}^{N \times N}$$

Note that $\mathbf{\Sigma}$ is related to the structural system via

$$\mathbf{\Sigma} = \mathbf{A}^{-1} (\mathbf{A}^{-1})'.$$

The key implication is that estimation of the reduced-form VAR delivers $\mathbf{B}$ and $\mathbf{\Sigma}$, but does not identify the contemporaneous impact matrix $\mathbf{A}^{-1}$. Recovering $\mathbf{A}^{-1}$ therefore requires additional identifying assumptions, which I impose using the recursive (Cholesky)

identification described in Section 2.2.4.3.

2.2.4.2 Bayesian VAR

Assume the reduced-form innovations $\{\mathbf{e}_t\}_{t=L+1}^T$ are independently and identically drawn from a normal distribution of mean $\mathbf{0} \in \mathbb{R}^N$ and variance-covariance matrix $\Sigma \in \mathbb{R}^{N \times N}$ across time $t = L + 1, \dots, T$, that is,

$$\mathbf{e}_t \stackrel{\text{i.i.d.}}{\sim} \mathcal{N}(\mathbf{0}, \Sigma) \quad t = L + 1, \dots, T$$

Stacking over time yields the matrix-normal representation

$$\mathbf{e} \sim \mathcal{MN}(\mathbf{0}, I_{T-L}, \Sigma)$$

where the identity matrix I_{T-L} reflects the assumption that reduced-form innovations $(\mathbf{e}_t)_{t=L+1}^T$ are independent and identically distributed over time.

I adopt a conjugate Normal-Inverse-Wishart (NIW) prior. To be specific, the innovation variance-covariance matrix $\Sigma \equiv \mathbb{E}[\mathbf{e}_t \mathbf{e}_t']$ follows

$$\Sigma \sim \mathcal{IW}(\underline{\mathbf{S}}, \underline{\alpha})$$

where $\mathcal{IW}$ denotes the inverse-Wishart distribution, $\underline{\mathbf{S}} \in \mathbb{R}^{N \times N}$ is a symmetric positive definite scale matrix, and $\underline{\alpha}$ is the degrees of freedom parameter.

Conditional on Σ , the coefficient matrix $\mathbf{B}$ follows a matrix-normal prior,

$$\mathbf{B} \mid \Sigma \sim \mathcal{MN}(\underline{\mathbf{B}}, \underline{\mathbf{V}}, \Sigma),$$

where $\underline{\mathbf{B}} \in \mathbb{R}^{(1+NL) \times N}$ is the prior mean matrix and $\underline{\mathbf{V}} \in \mathbb{R}^{(1+NL) \times (1+NL)}$ is a symmetric positive definite prior variance-covariance matrix. This Normal-Inverse-Wishart structure is conjugate and therefore delivers closed-form posterior updating.

Following an empirical Bayes approach, the hyperparameters governing the NIW prior—most importantly the overall tightness parameter $\lambda(h)$ —are selected by maximizing the marginal likelihood of the VAR, as in Giannone, Lenza and Primiceri (2015); Ferreira, Miranda-Agrippino and Ricco (2025); Miranda-Agrippino and Ricco (2021). The posterior remains NIW with closed-form updates. In particular, given prior parameters $(\underline{\mathbf{B}}, \underline{\mathbf{V}}, \underline{\mathbf{S}}, \underline{\alpha})$ and the observed data $(\mathbf{Y}, \mathbf{X})$, the joint posterior distribution of $(\boldsymbol{\Sigma}^{(d)}, \mathbf{B}^{(d)} \mid \boldsymbol{\Sigma}^{(d)})$ remains Normal-Inverse-Wishart. In particular, it can be shown that

$$\boldsymbol{\Sigma}^{(d)} \mid \mathbf{Y}, \mathbf{X} \sim \mathcal{IW}(\bar{\mathbf{S}}, \bar{\alpha})$$

and, conditional on $\boldsymbol{\Sigma}^{(d)} \in \mathbb{R}^{N \times N}$,

$$\mathbf{B}^{(d)} \mid \boldsymbol{\Sigma}^{(d)}, \mathbf{Y}, \mathbf{X} \sim \mathcal{MN}(\bar{\mathbf{B}}, \bar{\mathbf{V}}, \boldsymbol{\Sigma}^{(d)}).$$

where the posterior parameters are given by

$$\begin{aligned}\bar{\mathbf{V}} &= (\mathbf{X}'\mathbf{X} + \underline{\mathbf{V}}^{-1})^{-1} \\ \bar{\mathbf{B}} &= \bar{\mathbf{V}}(\mathbf{X}'\mathbf{Y} + \underline{\mathbf{V}}^{-1}\underline{\mathbf{B}}) \\ \bar{\mathbf{S}} &= \underline{\mathbf{S}} + \frac{1}{2}(\mathbf{Y}'\mathbf{Y} + \underline{\mathbf{B}}'\underline{\mathbf{V}}^{-1}\underline{\mathbf{B}} - \bar{\mathbf{B}}'\bar{\mathbf{V}}^{-1}\bar{\mathbf{B}}) \\ \bar{\alpha} &= \underline{\alpha} + (T - L).\end{aligned}$$

Posterior inference is conducted by drawing

$$\left\{ \left(\boldsymbol{\Sigma}^{(d)}, \mathbf{B}^{(d)} \mid \boldsymbol{\Sigma}^{(d)} \right) \right\}_{d=1}^D$$

from the posterior distribution.

2.2.4.3 Identification: Cholesky decomposition

This subsection describes how I use a recursive (Cholesky) identification scheme to map reduced-form posterior draws $\{(\boldsymbol{\Sigma}^{(d)}, \mathbf{B}^{(d)} \mid \boldsymbol{\Sigma}^{(d)})\}_{d=1}^D$ into structural impulse response functions $\{\mathcal{R}^{(d)}\}_{d=1}^D$, where $\mathcal{R}^{(d)} \equiv \left(\mathcal{R}_0^{(d)}, \dots, \mathcal{R}_H^{(d)}\right)' \in \mathbb{R}^{N(H+1)}$ with $\mathcal{R}_h^{(d)} \in \mathbb{R}^N$ for all impulse response draws $d = 1, \dots, D$ and horizons $h = 0, 1, \dots, H$.

Fix an arbitrary posterior draw $d \in \{1, \dots, D\}$. This draw provides a reduced-form variance-covariance matrix of innovations $\boldsymbol{\Sigma}^{(d)} \in \mathbb{R}^{N \times N}$ with $\mathbb{E}[\mathbf{e}_t \mathbf{e}_t'] = \boldsymbol{\Sigma}^{(d)}$, as well as a corresponding lag coefficient matrix $\mathbf{B}^{(d)} \mid \boldsymbol{\Sigma}^{(d)} \in \mathbb{R}^{(1+NL) \times N}$.

Let $\boldsymbol{\epsilon}_t \in \mathbb{R}^N$ denote the vector of structural shocks, normalized such that $\mathbb{E}[\boldsymbol{\epsilon}_t] = \mathbf{0}$ and $\mathbb{E}[\boldsymbol{\epsilon}_t \boldsymbol{\epsilon}_t' \mid \mathcal{F}_{t-1}] = I_N$. Assume that reduced-form innovations $\mathbf{e}_t$ are linear combinations of structural shocks $\boldsymbol{\epsilon}_t$ through the contemporaneous impact matrix $\mathbf{A}^{-1}$:

$$\mathbf{e}_t = \mathbf{A}^{-1} \boldsymbol{\epsilon}_t.$$

Then the reduced-form variance-covariance matrix satisfies

$$\boldsymbol{\Sigma}^{(d)} = \mathbb{E}[\mathbf{e}_t \mathbf{e}_t'] = \mathbf{A}^{-1} \mathbb{E}[\boldsymbol{\epsilon}_t \boldsymbol{\epsilon}_t' \mid \mathcal{F}_{t-1}] (\mathbf{A}^{-1})' = \mathbf{A}^{-1} (\mathbf{A}^{-1})'.$$

For each posterior draw d , I therefore recover the corresponding contemporaneous impact matrix $\mathbf{A}^{-1, (d)}$ by applying the Cholesky decomposition to $\boldsymbol{\Sigma}^{(d)}$. To be specific, let $\mathbf{P}^{(d)}$ denote the lower-triangular Cholesky factor such that

$$\boldsymbol{\Sigma}^{(d)} = \mathbf{P}^{(d)} \mathbf{P}^{(d)'}$$

I set $\mathbf{A}^{-1, (d)} \equiv \mathbf{P}^{(d)}$. This corresponds to a recursive identification scheme. The ordering of variables in $\mathbf{y}_t$ is chosen such that the policy rate appears last, implying that monetary policy can respond contemporaneously to all other variables in the system, while its effects on other variables occur with a lag within the period.

2.2.4.4 Constructing IRFs from the Bayesian VAR

To construct impulse responses for each posterior draw $d \in \{1, 2, \dots, D\}$, I begin with the posterior draw of the reduced-form lag coefficient matrix $\mathbf{B}^{(d)} \mid \boldsymbol{\Sigma}^{(d)}$, which stacks both the intercept and lag coefficients:

$$\mathbf{B}^{(d)} \mid \boldsymbol{\Sigma}^{(d)} = \begin{pmatrix} \mathbf{C}_0^{(d)} \\ \mathbf{B}_1^{(d)} \\ \vdots \\ \mathbf{B}_L^{(d)} \end{pmatrix} \in \mathbb{R}^{(1+NL) \times N}$$

where $\mathbf{C}_0^{(d)} \in \mathbb{R}^{1 \times N}$ is the intercept and each $\mathbf{B}_\ell^{(d)} \in \mathbb{R}^{N \times N}$ for $\ell = 1, \dots, L$ is the individual lag coefficient matrix. Since the intercept captures only deterministic time-trend components, I drop the intercept matrix $\mathbf{C}_0^{(d)}$ and work exclusively with the lag matrices $\left(\mathbf{B}_\ell^{(d)}\right)_{\ell=1}^L$ when constructing impulse responses.

To represent the dynamics in a state-space (companion) form, first define the stacked state vector of impulse responses for posterior draw d at horizon h as

$$\mathbf{R}_h^{(d)} \equiv \begin{pmatrix} \mathbf{y}_h^{\text{IRF},(d)} \\ \mathbf{y}_{h-1}^{\text{IRF},(d)} \\ \vdots \\ \mathbf{y}_{h-L+1}^{\text{IRF},(d)} \end{pmatrix} \in \mathbb{R}^{NL}, \quad \mathbf{y}_h^{\text{IRF},(d)} = \begin{pmatrix} y_h^{\text{IRF},1,(d)} \\ y_h^{\text{IRF},2,(d)} \\ \vdots \\ y_h^{\text{IRF},N,(d)} \end{pmatrix} \in \mathbb{R}^N.$$

where $y_h^{\text{IRF},n,(d)}$ denotes the impulse response of variable n at horizon h for posterior draw d , and $\mathbf{y}_h^{\text{IRF},(d)}$ stacks the impulse responses of all N variables at that horizon h for the posterior draw d .

Given the lag matrices $\left(\mathbf{B}_\ell^{(d)}\right)_{\ell=1}^L$, define the companion matrix for draw d as $\mathbf{F}^{(d)}$:

$$\mathbf{F}^{(d)} \equiv \begin{pmatrix} \mathbf{B}_1^{(d)} & \mathbf{B}_2^{(d)} & \cdots & \mathbf{B}_{L-1}^{(d)} & \mathbf{B}_L^{(d)} \\ I_N & \mathbf{0} & \cdots & \mathbf{0} & \mathbf{0} \\ \mathbf{0} & I_N & \cdots & \mathbf{0} & \mathbf{0} \\ \vdots & \vdots & \ddots & \vdots & \vdots \\ \mathbf{0} & \mathbf{0} & \cdots & I_N & \mathbf{0} \end{pmatrix} \in \mathbb{R}^{NL \times NL}$$

Let $\boldsymbol{\epsilon}_0^{MP} \in \mathbb{R}^N$ denote a unit monetary policy shock, defined as a vector with a one in the position corresponding to the policy rate (ordered last) and zeros elsewhere. The initial impact of a structural shock enters through the contemporaneous impact matrix $\mathbf{A}^{-1,(d)}$, so the state vector at impact is initialized as

$$\mathbf{R}_0^{(d)} \equiv \begin{pmatrix} \mathbf{A}^{-1,(d)} \boldsymbol{\epsilon}_0^{MP} \\ \mathbf{0} \\ \vdots \\ \mathbf{0} \end{pmatrix} \in \mathbb{R}^{NL}.$$

The impulse response dynamics then follow by iterating forward on the companion form:

$$\mathbf{R}_h^{(d)} = \mathbf{F}^{(d)} \mathbf{R}_{h-1}^{(d)}, \quad h \geq 1.$$

For each posterior draw d , I collect the impulse responses across horizons $h = 0, 1, \dots, H$ into the stacked vector

$$\mathcal{R}^{(d)} \equiv \begin{pmatrix} \mathcal{R}_0^{(d)} \\ \mathcal{R}_1^{(d)} \\ \vdots \\ \mathcal{R}_H^{(d)} \end{pmatrix} \in \mathbb{R}^{N(H+1)}, \quad \mathcal{R}_h^{(d)} \in \mathbb{R}^N,$$

where each $\mathcal{R}_h^{(d)}$ corresponds to the first N elements of the state vector $\mathbf{R}_h^{(d)}$ and therefore represents the impulse responses of all N variables at horizon h . Notice that the impulse response object $\mathcal{R}^{(d)}$ is time-invariant: it does not vary across year-quarter t , in contrast to the forecast-path object $\mathcal{Y}_t^{0,(m)}$, which varies over year-quarter t .

Throughout the analysis, I estimate the Bayesian VAR for each country using only observations from the post-inflation-target period. Accordingly, the resulting impulse response draws $\{\mathcal{R}^{0,(d)}\}_{d=1}^D$ correspond to the baseline inflation-target policy regime, where the superscript 0 denotes the inflation-target baseline rule.

2.2.5 Implementation

This subsection translates the theoretical framework into an empirical procedure. The key implication of Sections 2.2.3 and 2.2.4 is that, empirically, the OPP statistic δ_t depends on two sufficient statistics:

1. $\{\mathcal{R}^{0,(d)}\}_{d=1}^D$, where each $\mathcal{R}^{0,(d)} \in \mathbb{R}^{N(H+1)}$ is a posterior draw of impulse response coefficients to a one-unit monetary policy shock under the baseline policy rule ϕ^0 ;
2. $\{\mathcal{Y}_t^{0,(m)}\}_{m=1}^M$, where each $\mathcal{Y}_t^{0,(m)} \in \mathbb{R}^{N(H+1)}$ is a draw of the forecast path under the same baseline rule ϕ^0

Given the collections of draws $\{\mathcal{R}^{0,(d)}\}_{d=1}^D$ and $\{\mathcal{Y}_t^{0,(m)}\}_{m=1}^M$, I compute the (scalar) OPP statistic for each pair (m, d) as

$$\delta_t^{*,(m,d)} \equiv -(\mathcal{R}^{0,(d)'} \mathbf{W} \mathcal{R}^{0,(d)})^{-1} \left(\mathcal{R}^{0,(d)'} \mathbf{W} \mathcal{Y}_t^{0,(m)} \right) \in \mathbb{R}^1 \quad (2.3)$$

Uncertainty is summarized using posterior quantiles of the distribution of $\{\delta_t^{*,(m,d)}\}$ across all (m, d) at each year-quarter t . To be specific, I report the 68th, 90th, and 95th credible intervals (equal-tailed).

2.3 EMPIRICAL STRATEGY

This section explains how I use the dual interpretation of the OPP statistic to evaluate inflation-target monetary policy in emerging markets. The analysis proceeds in two steps.

First, Section 2.3.1 adopts a revealed-preference approach. I assume that, in each period, the observed policy rate is optimal given the central bank’s contemporaneous policy priorities and information set. Under this assumption, I recover the period-specific policy weight vector that minimizes deviations of the OPP statistic from zero. The resulting time series of weights summarizes the evolution of implicit trade-offs across output, inflation, and exchange rate stabilization. This framework allows me to study how policy priorities evolve around the adoption of inflation targeting and during crisis episodes. I also test whether the inferred inflation weight is statistically consistent with the benchmark value of 0.5 implied by a standard dual-mandate framework.

Second, Section 2.3.2 imposes time-invariant policy priorities and asks a different question: under fixed weights, is there systematic evidence of policy misalignment? I consider three benchmark weight schemes: an inflation-target benchmark $(\frac{1}{2}, \frac{1}{2}, 0)$, an equal-emphasis benchmark $(\frac{1}{3}, \frac{1}{3}, \frac{1}{3})$, and a globally optimal benchmark that minimizes the sum of squared deviations of the OPP statistic from zero across periods.

2.3.1 Revealed preference: period-specific optimal weights

In many emerging markets, formal policy mandates are incomplete or opaque, making it difficult to assess credibility by directly comparing observed actions to an explicit loss function. Instead, I infer the implicit policy priorities from observed policy rates and forecast data.

As an illustrative example, under a credible inflation-target framework, the recovered policy weights should resemble $(\frac{1}{2}, \frac{1}{2}, 0)$. If instead the estimated weight vector that rationalizes the observed policy rate is much closer to $(1, 0, 0)$, this would indicate that the central bank primarily prioritizes output stabilization, contradicting its stated inflation-target

commitment.

Formally, I recover the period-specific implicit trade-offs summarized by the weight vector

$$\boldsymbol{\omega}_t \equiv (\omega_t^x, \omega_t^\pi, \omega_t^\varepsilon), \quad \omega_t^x + \omega_t^\pi + \omega_t^\varepsilon = 1, \quad \omega_t^i \geq 0 \quad \forall i \in \{x, \pi, \varepsilon\}.$$

Intuitively, if the observed policy rate is optimal under the true contemporaneous objectives, the implied policy weight vector should make the OPP statistic close to zero: no small perturbation to the policy rate can further improve welfare. Although period-specific estimates are noisy, persistent patterns in the inferred weights over time provide informative evidence about underlying policy priorities.

To be specific, for each year-quarter t , I estimate $\boldsymbol{\omega}_t^*$ by minimizing the dispersion of the OPP statistic around zero, combining forecast draws $\{\mathcal{Y}_t^{0,(m)}\}_{m=1}^M$ with impulse-response draws $\{\mathcal{R}^{0,(d)}\}_{d=1}^D$.

Then, I solve

$$\boldsymbol{\omega}_t^* = \arg \min_{\boldsymbol{\omega}} \mathcal{L}_t^{\text{Huber}}(\boldsymbol{\omega}) = \frac{1}{MD} \sum_{m=1}^M \sum_{d=1}^D \ell \left(\delta_t^{(m,d)}(\boldsymbol{\omega}) \right) \quad (2.4)$$

where the draw-specific OPP statistic under candidate weight vector $\boldsymbol{\omega}$ is

$$\delta_t^{(m,d)}(\boldsymbol{\omega}) = - \left(\mathcal{R}^{0,(d)'} \mathbf{W} \mathcal{R}^{0,(d)} \right)^{-1} \left(\mathcal{R}^{0,(d)'} \mathbf{W} \mathcal{Y}_t^{0,(m)} \right), \quad \mathbf{W} \equiv \text{diag}(\boldsymbol{\omega}) \otimes I_{H+1} \quad (2.5)$$

I employ a Huber loss function to reduce sensitivity to extreme realizations:

$$\ell(\delta) = \begin{cases} \frac{1}{2} \left(\frac{\delta}{s_t} \right)^2, & \text{if } |\delta| \leq k s_t, \\ k \left| \frac{\delta}{s_t} \right| - \frac{1}{2} k^2, & \text{otherwise,} \end{cases}$$

where $s_t = \text{sd} \left(\{ \delta_t^{(m,d)}(\boldsymbol{\omega}) \}_{m,d} \right)$ and $k = 1.5$.

The optimization is performed over an unconstrained parameterization smoothly mapped

to the unit simplex and solved using the L-BFGS algorithm with multiple random initializations. The resulting sequence $\{\boldsymbol{\omega}_t^*\}_{t=1}^T$ summarizes the evolution of inferred policy priorities over time.

2.3.2 Time-Invariant policy weight benchmarks

2.3.2.1 Global-optimal benchmark

I next estimate a single *time-invariant* weight vector $\boldsymbol{\omega}^{\text{GO}}$ that best rationalizes the policy rate over the entire inflation-target period:

$$\boldsymbol{\omega}^{\text{GO}} = \arg \min_{\boldsymbol{\omega}} \frac{1}{T} \sum_{t=1}^T \left(\frac{1}{MD} \sum_{m=1}^M \sum_{d=1}^D \ell \left(\delta_t^{(m,d)}(\boldsymbol{\omega}) \right) \right),$$

where $\delta_t^{(m,d)}(\boldsymbol{\omega})$ is defined in Equation 2.5. In addition to the Huber loss, I include two regularization terms: (i) an ℓ_2 penalty on the unconstrained parameterization, and (ii) a mild entropy penalty on $\boldsymbol{\omega}$ to discourage corner solutions. The optimization is performed using the L-BFGS algorithm with multiple random initializations.

The resulting vector $\boldsymbol{\omega}^{\text{GO}}$ defines the global-optimal benchmark used in subsequent comparisons. Using this benchmark, I construct a time series of OPP statistics as follows. For each period t and each pair of draws (m, d) , I compute

$$\delta_t^{\text{GO},(m,d)} = - \left(\mathcal{R}^{0,(d)'} \mathbf{W}^{\text{GO}} \mathcal{R}^{0,(d)} \right)^{-1} \left(\mathcal{R}^{0,(d)'} \mathbf{W}^{\text{GO}} \mathcal{Y}_t^{0,(m)} \right), \quad \mathbf{W}^{\text{GO}} \equiv \text{diag}(\boldsymbol{\omega}^{\text{GO}}) \otimes I_{H+1}.$$

I then summarize the cross-draw distribution using the median,

$$\delta_t^{\text{GO},\text{med}} \equiv \text{median} \left\{ \delta_t^{\text{GO},(m,d)} \right\}_{m,d},$$

which yields the period-specific OPP series $(\delta_t^{\text{GO}})_{t=1}^T$ under the global-optimal benchmark.

2.3.2.2 Fixed-weight benchmark comparisons

In addition to the global-optimal weight vector discussed above, I evaluate policy rates under two standard fixed-weight benchmarks: (i) an equal-weight benchmark (EQ), $\boldsymbol{\omega}^{\text{EQ}} = (\frac{1}{3}, \frac{1}{3}, \frac{1}{3})$, which assigns equal emphasis to output, inflation, and exchange rate stabilization; and (ii) an inflation-target benchmark (IT), $\boldsymbol{\omega}^{\text{IT}} = (\frac{1}{2}, \frac{1}{2}, 0)$, which mimics a dual-mandate framework by assigning equal weight to output and inflation while placing zero weight on exchange rate stabilization.

For each year-quarter t , I combine forecast-path draws $\{\mathcal{Y}_t^{0,(m)}\}_{m=1}^M$ with impulse-response draws $\{\mathcal{R}^{0,(d)}\}_{d=1}^D$ to obtain a collection of OPP statistics under each benchmark. Specifically, under the equal-weight benchmark,

$$\delta_t^{\text{EQ},(m,d)} = -(\mathcal{R}^{0,(d)'} \mathbf{W}^{\text{EQ}} \mathcal{R}^{0,(d)})^{-1} \left(\mathcal{R}^{0,(d)'} \mathbf{W}^{\text{EQ}} \mathcal{Y}_t^{0,(m)} \right), \quad \mathbf{W}^{\text{EQ}} \equiv \text{diag}\left(\frac{1}{3}, \frac{1}{3}, \frac{1}{3}\right) \otimes I_{H+1},$$

and under the inflation-target benchmark,

$$\delta_t^{\text{IT},(m,d)} = -(\mathcal{R}^{0,(d)'} \mathbf{W}^{\text{IT}} \mathcal{R}^{0,(d)})^{-1} \left(\mathcal{R}^{0,(d)'} \mathbf{W}^{\text{IT}} \mathcal{Y}_t^{0,(m)} \right), \quad \mathbf{W}^{\text{IT}} \equiv \text{diag}\left(\frac{1}{2}, \frac{1}{2}, 0\right) \otimes I_{H+1}$$

For each year-quarter t , this yields three collections of OPP draws, each with MD elements: $\{\delta_t^{\text{IT},(m,d)}\}_{m,d}$, $\{\delta_t^{\text{EQ},(m,d)}\}_{m,d}$, and $\{\delta_t^{\text{GO},(m,d)}\}_{m,d}$.

For each benchmark $s \in \{\text{IT}, \text{EQ}, \text{GO}\}$, I summarize the cross-draw distribution using the median,

$$\delta_t^{s,\text{med}} \equiv \text{median} \left\{ \delta_t^{s,(m,d)} \right\}_{m,d}.$$

I plot the resulting median series $(\delta_t^{\text{IT},\text{med}})_{t=1}^T$, $(\delta_t^{\text{GO},\text{med}})_{t=1}^T$, and $(\delta_t^{\text{EQ},\text{med}})_{t=1}^T$ in panel (b) of the figures. In addition, because the primary focus of the paper is inflation targeting, I also report uncertainty bands for the IT benchmark by plotting the 68th and 90th percentiles of $\{\delta_t^{\text{IT},(m,d)}\}_{m,d}$ for each t .

2.3.3 Interpretation

How should we interpret the results? If all weighting schemes imply OPP prescriptions with the same direction and similar magnitude, the conclusion is robust: policy rates are systematically misaligned in that direction. In this case, under all benchmark preferences considered, no specification rationalizes the observed policy rate as optimal, implying that there remains scope for welfare-improving policy adjustment. Understanding the institutional or political motivations behind such episodes may require additional qualitative evidence, such as central bank communications or official policy statements (e.g., concerns about financial stability).

By contrast, if the inflation-target benchmark yields OPP values close to zero while alternative benchmarks imply significant deviations, this reveals the underlying trade-offs faced by policymakers. For example, if the equal-weight benchmark prescribes tightening while the inflation-target benchmark implies no deviation, this indicates that the central bank places relatively greater weight on inflation stabilization than on output or exchange-rate stabilization. Comparisons across benchmarks therefore help uncover the implicit policy priorities that rationalize observed decisions under one preference structure but not under others, especially during financial crises.

2.4 DATA

2.4.1 Macroeconomic variables

To estimate impulse responses, I use quarterly seasonally adjusted real GDP growth from the IMF’s International Financial Statistics (IFS), supplemented with the Macro-financial dataset of [Camous, Monnet and Puy \(2025\)](#). Quarterly consumer price indices are also drawn from the IFS. Since many CPI series are not seasonally adjusted, I deseasonalize them using

the X-13 method. Effective exchange rate indices are obtained from the IFS as well.⁷ All macroeconomic series are cross-checked against the comprehensive dataset of Obstfeld and Zhou (2022) to ensure consistency.

The most challenging aspect of data construction concerns monetary policy rates. Many central banks in emerging markets did not operate with a clearly defined policy rate until relatively recently, and operational instruments vary substantially across countries and over time. Unlike the Federal Funds rate—an unsecured interbank overnight rate—Table 2.2 shows that policy rates in emerging markets often correspond to collateralized short-term rates with maturities ranging from overnight to several weeks. In some cases, the reported “policy rate” reflects the central bank’s discount rate rather than the marginal rate targeted through open market operations.

Institutional changes further complicate measurement. Several central banks have altered their operating monetary instruments over time, and episodes of high inflation and currency redenomination can mechanically distort reported interest rates.⁸ These features introduce substantial noise into cross-country comparisons.

To mitigate these issues, I restrict the impulse response estimation to periods in which each country operated with a single, clearly defined policy rate, as documented in Table 2.2. These periods typically coincide with the formal adoption of inflation-targeting regimes. This restriction also avoids distortions arising from high-inflation episodes and currency redenomination.

2.4.2 Expected deviations from the forecast path

The OPP statistic requires expected forecast paths $\mathcal{Y}_t^0 \in \mathbb{R}^{N(H+1)}$ of macroeconomic variables. I construct these using professional forecasts from the Economist Intelligence Unit (EIU) for real GDP growth, inflation, and currency depreciation. As an independent organization, the

⁷Bilateral exchange rates against the U.S. dollar are often excessively volatile. I therefore rely on trade-weighted exchange rates, which provide smoother and more informative measures of external conditions.

⁸For example, Brazil’s currency redenominations in the early 1990s reduced published policy rates by several orders of magnitude in the BIS series.

EIU provides long-horizon forecasts for a broad set of countries—many extending back to 1994—using a consistent methodology that facilitates cross-country comparability.⁹

These forecast paths vary over time because the EIU updates forecasts in each publication, while the impulse response estimates remain fixed over the sample.

The publication frequency of EIU reports evolves over time. Earlier issues are typically quarterly, while reports become monthly in the early 2000s. For consistency, I use quarterly forecasts in earlier periods and select forecasts from the January, April, July, and October issues when monthly publications are available.

EIU forecasts are reported as point forecasts for end-of-calendar-year outcomes up to five years ahead. As a result, each issue provides a term structure of expectations at unevenly spaced horizons. For example, the January issue contains forecasts approximately 11, 23, 35, 47, and 59 months ahead, while the April issue corresponds to roughly 8, 20, 32, 44, and 56 months ahead. Later issues follow the same structure with progressively shorter effective horizons.

To align these forecasts with the quarterly horizons used in the impulse response analysis, I linearly interpolate forecast paths across horizons at the monthly frequency and then retain values closest to 12, 24, 36, 48, and 60 months ahead (corresponding to 4, 8, 12, 16, and 20 quarters). This procedure yields consistent forecast paths up to 20 quarters ahead in every period, despite differences in publication timing. I then follow the procedure detailed in Section 2.2.3 to construct posterior draws of forecast paths $\{\mathcal{Y}_t^{0,(m)}\}_{m=1}^M$.

2.5 MAIN RESULTS AND DISCUSSIONS

Figures 2.1–2.12 summarize the main results for each emerging market. Each country is evaluated using a common two-panel structure. Visual inspection reveals systematic pat-

⁹Central banks’ internal forecasts may contain private information unavailable to professional forecasters, and analyses based on real-time policymaker data can differ systematically from those using ex post revised data (Romer and Romer, 2000; Orphanides, 2001). However, consistent historical central-bank forecasts are largely unavailable for emerging markets. Moreover, even when official projections exist, they may be subject to political pressures that bias reported expectations.

terns across country groups sharing similar monetary policy traditions, so the discussion is organized by economically meaningful country groups rather than alphabetically.

Panels (a) report time-varying revealed policy priorities—i.e., the policy weight vectors that minimize the OPP deviation from zero in each year-quarter. These series are informative about whether the inflation weight is economically meaningful and whether it remains stable over time, providing an intuitive and transparent indicator of inflation-target credibility.

Panels (b) compare OPP statistics across alternative rule-based benchmarks, with the inflation-target benchmark—featuring only the inflation-output tradeoff—as the primary reference. Conditional on EIU forecasts approximating policymakers’ information sets, the OPP framework provides a disciplined way to evaluate whether observed monetary policy responses during crises were insufficient, excessive, or mistimed. Comparisons with the equal-weight and global-optimal benchmarks quantify both the direction and the magnitude of deviations from the inflation-target benchmark.

2.5.1 Asian emerging markets: stable inflation emphasis with pronounced crisis-driven deviations

I begin with four Asian emerging markets: South Korea, Thailand, Indonesia, and the Philippines. Prior to the 2000s, these economies exhibited limited monetary autonomy under de facto pegged exchange rate regimes supporting export-led growth. Following the 1997 Asian Financial Crisis, reserve accumulation and institutional reforms expanded monetary policy flexibility, allowing the inflation-target adoption without fully compromising exchange rate stability.

Panels (a) of Figures 2.1–2.2 show substantial cross-country heterogeneity in revealed policy priorities. South Korea and Indonesia place relatively little weight on exchange rate stabilization—global-optimal weights are approximately $(0.50, 0.44, 0.06)$ and $(0.51, 0.36, 0.13)$ respectively. Moreover, their time-varying revealed policy priorities become more stable over time, converging toward a balanced inflation-output trade-off, consistent with the maturation

of inflation-target regimes and comparatively stronger democratic institutions and central bank independence.

In contrast, Thailand and the Philippines retain substantially higher revealed emphasis on exchange rate stabilization. Their global-optimal weights, $(0.31, 0.34, 0.35)$ and $(0.33, 0.35, 0.32)$ respectively, lie close to the equal-weight benchmark. This pattern plausibly reflects institutional and structural differences. Thailand's monetary framework has been shaped by repeated political disruptions and higher export exposure. The Philippine economy, in turn, is heavily dependent on foreign remittances, implying that exchange rate movements carry salient distributional and political consequences.

Crisis episodes generate pronounced temporary deviations in policy trade-offs across all four economies. Panels (a) of Figures 2.1 and 2.2 show that during the 1997 Asian Financial Crisis, South Korea and Indonesia experienced sharp spikes in exchange rate weight followed by immediate declines after the pegs broke. In the post-crisis period, output weights temporarily increased before reverting back toward inflation. Similar crisis-driven reweighting appears, to varying degrees, in Thailand and the Philippines. Overall, crisis episodes and their subsequent economic downturns are consistently associated with elevated concern for real activity or exchange rate stability.

Panels (b) of Figures 2.1–2.4 translate constant-weight benchmarks into counterfactual policy prescriptions. For South Korea, Panel (b) of Figure 2.1 indicates that during the 2008 Global Financial Crisis all benchmarks imply sizable easing—roughly 150–250 basis points—suggesting that the policy rate could have responded more aggressively. In contrast, during the 1997 Asian Financial Crisis, the inflation-target and global-optimal benchmarks prescribe monetary easing of 100–150 basis points, whereas the equal-weight benchmark suggests tightening of around 150 basis points. This divergence highlights the underlying policy dilemma between anchoring inflation expectations versus stabilizing output growth and the exchange rate.

Indonesia, Thailand, and the Philippines all display analogous tensions. In Indonesia

(Panel (b) of Figure 2.2), the equal-weight benchmark prescribes a 150 basis point tightening during the 1997 Asian Financial Crisis, while inflation-target and global-optimal benchmarks suggest smaller adjustments; all benchmarks subsequently converge toward easing of about 50 basis points immediately after the crisis. Thailand (Panel (b) of Figure 2.3) occupies an intermediate position: equal-weight and global-optimal benchmarks both recommend monetary tightening of about 50 basis points during the 1997 Asian Financial Crisis but not during the 2008 Global Financial Crisis, whereas the inflation-target benchmark prescribes modest easing (about 25 basis points) in both episodes. For the Philippines (Panel (b) of Figure 2.4), benchmark disagreement during crisis periods—equal-weight and global-optimal benchmarks suggest monetary tightening of about 50–100 basis points—similarly reflects competing stabilization objectives.

Overall, heterogeneity in monetary policy prescriptions across countries and crisis episodes reflects differences in institutional credibility and external vulnerability—precisely the dimensions that the OPP framework is designed to capture. These findings are consistent with Fraga, Goldfajn and Minella (2003)’s argument that emerging markets’ constraints on credibility and external shocks (e.g., sudden stop) necessitate flexible inflation-target implementations that balance multiple stabilization objectives.

2.5.2 Historically high-inflation Latin American economies: convergence toward inflation dominance

The second country group consists of Brazil, Chile, Colombia, Peru, and Mexico—emerging economies with histories of chronic high inflation or hyperinflation. These cases provide a natural laboratory for examining how inflation-target adoption reshaped monetary policy priorities during the transition to low-inflation regimes.

Brazil exhibits the sharpest regime shift. Panel (a) of Figure 2.5 shows that prior to inflation-target adoption, policy priorities fluctuated substantially and placed limited emphasis on inflation stabilization, with greater attention to output and exchange rate. This

pattern is consistent with heightened concerns about regional contagion during Argentina's 2001 default episode. Following the consolidation of the inflation-target regime in the mid-2000s, the inflation weight rises sharply and becomes persistently dominant. The global-optimal benchmark (0.56, 0.33, 0.11) confirms this post-crisis inflation dominance and the stabilization of policy priorities.

Mexico displays a more gradual but similarly sharp and persistent transition. Panel (a) of Figure 2.6 indicates that before the formal adoption of the inflation-target framework in 2001, revealed policy priorities placed little emphasis on inflation stabilization, consistent with the post-Tequila Crisis environment. Afterward, the inflation weight rises sharply and remains elevated throughout the sample period. The global-optimal benchmark (0.32, 0.62, 0.06) suggests near-exclusive emphasis on inflation-output stabilization, with minimal concern for exchange-rate objectives.

Chile follows a comparable but more balanced pattern. Panel (a) of Figure 2.7 shows that after formal inflation targeting in 1999, inflation stabilization becomes the dominant objective, though exchange-rate and output considerations remain non-negligible. Temporary shifts toward output stabilization emerge during major downturns, most notably the 2008 Global Financial Crisis. The global-optimal benchmark (0.44, 0.31, 0.25) reflects this more balanced but still inflation-centered regime.

Colombia also exhibits sustained inflation dominance following the inflation-target adoption in 1999. Panel (a) of Figure 2.8 shows that post-adoption fluctuations in policy weights primarily reflect trade-offs between output and exchange rate, while the inflation weight remains consistently high and stable. The global-optimal benchmark (0.6, 0.36, 0.04) confirms limited concern for exchange rate stabilization. Peru, in contrast, stands out as an exception: Panel (a) of Figure 2.9 shows persistently elevated exchange rate weights even after inflation targeting, indicating continued emphasis on external stability.

Panels (b) of Figures 2.5–2.9 reveal greater volatility in monetary prescriptions than in the Asian group of Section 2.5.1. For Brazil, Panel (b) of Figure 2.5 shows that equal-weight

and global-optimal benchmarks often prescribe tighter monetary policy during major stress episodes, including the early-2000s turbulence, the 2008 Global Financial Crisis, and the 2015 political-fiscal crisis, while the inflation-target benchmark implies comparatively muted adjustments. This divergence underscores the extent to which Brazil’s realized policy appears aligned with inflation emphasis rather than broader stabilization objectives.

Chile and Peru exhibit smaller and smoother OPP deviations. Panel (b) of Figure 2.7 shows that in Chile, disagreement across benchmarks is modest and largely confined to crisis periods. For example, the equal-weight benchmark prescribes a mild 50 basis point tightening during the 2008 Global Financial Crisis, while the inflation-target benchmark suggests the opposite. Panel (b) of Figure 2.9, in contrast, shows consistent benchmark prescriptions of mild tightening of policy rates by about 50–100 basis points during the early inflation-target adoption period, the 2008 Global Financial Crisis, and the COVID-19 pandemic.

Mexico shows the most volatile OPP statistics in this country group. During the 1994–2000 currency crisis period, all benchmarks unanimously suggest substantial tightening, often exceeding 100–250 basis points. After the inflation-target adoption, the OPP deviations under the inflation-target benchmark become smaller and more stable, while equal-weight and global-optimal benchmarks remain considerably more volatile. This pattern suggests that the Bank of Mexico has consistently prioritized inflation stabilization relative to alternative objectives.

Overall, the Latin American evidence illustrates how inflation-target regimes facilitated rapid and durable convergence toward low and stable inflation environments. Even in economies with severe historical inflation episodes, the OPP framework shows a gradual stabilization of revealed policy priorities toward inflation dominance, consistent with successful re-anchoring of inflation expectations and institutional credibility.

2.5.3 Extreme performers and underperformers

This section highlights three notable examples for comparing inflation-target credibility: South Africa as a strong performer, Turkey as a clear underperformer, and Israel as an intermediate case.

South Africa stands out as a benchmark case of credible inflation targeting. Panel (a) of Figure 2.10 shows a persistently high and stable inflation weight, with limited emphasis on exchange rate stabilization. The global-optimal benchmark $(0.39, 0.54, 0.07)$ confirms this structure. Panel (b) further shows that the inflation-target OPP statistic remains tightly centered around zero, while the equal-weight benchmark occasionally prescribes modest tightening of 50–100 basis points. Together, these patterns indicate that South Africa consistently prioritizes inflation stabilization over competing objectives and exhibits close alignment with a credible inflation-target regime.

Turkey, by contrast, represents the clearest breakdown of inflation-target credibility in the sample. Panel (a) of Figure 2.12 shows persistent dominance of output stabilization with little sustained weight on inflation. Panel (b) displays large—but consistently aligned tightening prescriptions across all three benchmarks—OPP deviations throughout the sample. During 1994–2005, all benchmarks prescribe aggressive monetary tightening, often exceeding 250 basis points. Although OPP deviations appear more stable between 2005 and 2020, most benchmarks continue to imply that policy rates should have been higher by roughly 50 basis points. After 2020, deviations rise sharply again across all benchmarks, with prescriptions frequently exceeding 100–250 basis points. This pattern is consistent with a collapse of inflation-target credibility rather than simple mis-calibration of policy trade-offs, in line with well-documented political interference in monetary policy since the 2020s.

Israel occupies an intermediate position. Panel (a) of Figure 2.11 shows that inflation generally receives substantial weight over the sample, consistent with the global-optimal benchmark $(0.24, 0.48, 0.28)$. At the same time, policy priorities exhibit greater time variation

than in South Africa, with occasional shifts toward output stabilization, most noticeably since the 2020s. Panel (b) shows OPP deviations that are larger and more volatile than in South Africa but far smaller than in Turkey. During recent periods of heightened external and geopolitical stress, the inflation-target benchmark typically prescribes modest tightening of around 50 basis points, broadly consistent with the suggestions from other benchmarks. Overall, Israel appears to maintain inflation-target credibility but faces sharper short-run trade-offs due to geopolitical conflicts than highly stable benchmark cases.

Taken together, these cases underscore the central role of institutional stability for inflation-target credibility. Countries such as South Africa, where revealed policy priorities are stable and OPP deviations remain small, exhibit behavior consistent with the inflation-target framework. In contrast, Turkey illustrates how persistent political interference and shifting policy priorities can generate large and systematic deviations from inflation-target benchmarks. Intermediate cases such as Israel highlight the constraints faced by small open economies often exposed to external shocks and geopolitical risk, where maintaining inflation-target credibility may require temporary departures from strict rule-based policy.

2.5.4 Cross-country lessons

The above results yield three central lessons. First, time-varying revealed policy priorities provide a transparent and informative measure of evolving monetary policy priorities. Among successful inflation-target regimes—including Brazil, Chile, Colombia, Indonesia, Mexico, South Korea, and South Africa—the inflation-target adoption coincides with a pronounced and persistent increase in the inflation weight, alongside gradual convergence toward lower and more stable inflation observed over the past two decades. Countries such as Brazil, Colombia, and South Africa appear to maintain strict adherence to inflation dominance throughout, while others—such as Chile and Indonesia—nonetheless exhibit temporary reallocations toward output stabilization; however, those with historically stable inflation weights tend to revert more quickly to inflation primacy after shocks. Both patterns are consistent with

credible and internally coherent policy frameworks.

Second, the OPP statistic provides a quantitative measure of policy performance. Across most countries, alternative benchmarks—inflation-target, equal-weight, and global-optimal—deliver broadly similar directional policy prescriptions despite differences in magnitude. Extreme cases illustrate this contrast sharply. For Turkey, all benchmarks consistently suggest aggressive tightening, indicating severe credibility failures. For strong performers such as South Korea, the OPP statistics instead highlight missed opportunities for more timely and decisive easing during the 2008 Global Financial Crisis.

Third, cross-country heterogeneity reflects differences in institutional development and approaches to crisis management. Latin American economies with histories of high inflation exhibit the sharpest and most durable transitions toward inflation dominance following the inflation-target adoption. By contrast, countries with weaker institutional commitment—most notably Turkey—continue to display unstable policy weights on inflation and large OPP deviations despite formal regime adoption.

Overall, credible inflation-target regimes are characterized by convergence toward stable inflation dominance and declining OPP deviations over time. South Korea, Brazil, Chile, and Mexico exhibit this pattern. At the same time, small open economies such as Israel and Thailand face additional constraints from geopolitical risk and external shocks, leading to temporary departures from the strict inflation-target benchmark. Persistent OPP deviations across all benchmarks in countries such as Turkey, however, underscore the limits of formal regime adoption in the absence of institutional discipline.

2.6 CONCLUDING REMARKS

High inflation imposes substantial economic costs by distorting price signals, redistributing wealth arbitrarily, and weakening expectation formation—an essential channel through which both conventional and unconventional monetary policies operate. In emerging markets,

where expectations are often fragile and credibility hard-won, maintaining a consistent and transparent inflation-target framework is therefore central to macroeconomic stabilization.

This paper evaluates whether central banks in emerging markets credibly adhere to their stated inflation-target frameworks. While credibility is widely viewed as central to anchoring inflation expectations, many emerging markets have long histories of inflation volatility and repeated episodes of macroeconomic instability. This raises the question of whether formal inflation-target commitments meaningfully discipline monetary policy in practice.

To address this question, I develop an empirical framework based on optimal policy perturbation (OPP) statistics to assess monetary policy credibility. The analysis yields three main findings:

First, the revealed-preference estimates show that many emerging markets have placed increasing and more stable weight on inflation stabilization since the early 2000s. This pattern is especially pronounced in countries with histories of chronic or high inflation, such as Brazil, Chile, and Mexico. The evolution of implied policy weights suggests a structural reorientation of monetary policy toward price stability, consistent with a strengthening of inflation-target credibility over time.

Second, deviations from inflation-focused behavior are concentrated around crisis episodes. During severe downturns, central banks often shift weight temporarily toward output stabilization, reflecting the inherent trade-offs of crisis management in small open economies. Importantly, however, the optimal direction and magnitude of policy adjustment vary substantially across countries. For example, during the 2008 Global Financial Crisis, the OPP framework suggests that the Bank of Korea should have eased more aggressively, whereas policy in Peru should have been modestly tighter. These contrasts underscore the role of country-specific institutional features, external vulnerabilities, and inflation dynamics in shaping optimal monetary policy responses.

Third, the framework identifies clear cases of weak or deteriorating credibility. Most

notably, Turkey exhibits persistent and large deviations from the inflation-target-consistent policy path, even under identifying assumptions that are maximally favorable to the central bank. This evidence suggests that the formal adoption of inflation targeting is insufficient in the absence of institutional discipline and sustained political commitment.

Overall, the evidence is cautiously optimistic. Since the turn of the millennium, many central banks in emerging markets have increasingly prioritized inflation stabilization, helping to rationalize the sustained decline in inflation observed across these economies. While temporary departures occur during crises—particularly in small open economies exposed to financial or geopolitical shocks—most central banks subsequently revert to inflation-dominant policy behavior. This pattern points to a gradual strengthening of monetary credibility and supports a favorable outlook for the long-run effectiveness of inflation-target frameworks in emerging markets.

Table 2.1: Inflation-target framework and inflation outcomes in emerging markets

Country (1)	Target framework			Inflation outcomes (2024, YoY)		
	Start (2)	Type (3)	Current level(s) (4)	Actual (5)	Forecast (1yr) (6)	Forecast (avg) (7)
BRA	1999	Band	3% $\pm$ 1.5% pt	4.3%	3.8%	3.6%
CHL	1999	Point	3%	4.2%	4.1%	3.1%
COL	1999	Point	3%	6.4%	4.2%	3.2%
IDN	2005	Band	2.5% $\pm$ 1% pt	2.2%	2.5%	2.9%
ISR	1992	Band	1%–3%	3.0%	3.1%	1.9%
KOR	1997	Point	2%	2.3%	1.8%	1.7%
MEX	2001	Band	3% $\pm$ 1% pt	4.6%	3.9%	3.3%
PER	2002	Point	2%	2.0%	2.3%	2.4%
PHL	2002	Band	2%–4%	3.2%	2.6%	3.0%
THA†	2000	Band	1%–3%	0.5%	2.2%	2.1%
TUR†	2006	Band	5% $\pm$ 2% pt	49.4%	28.5%	18.0%
ZAF	2000	Point	4.5%	4.3%	4.2%	4.2%

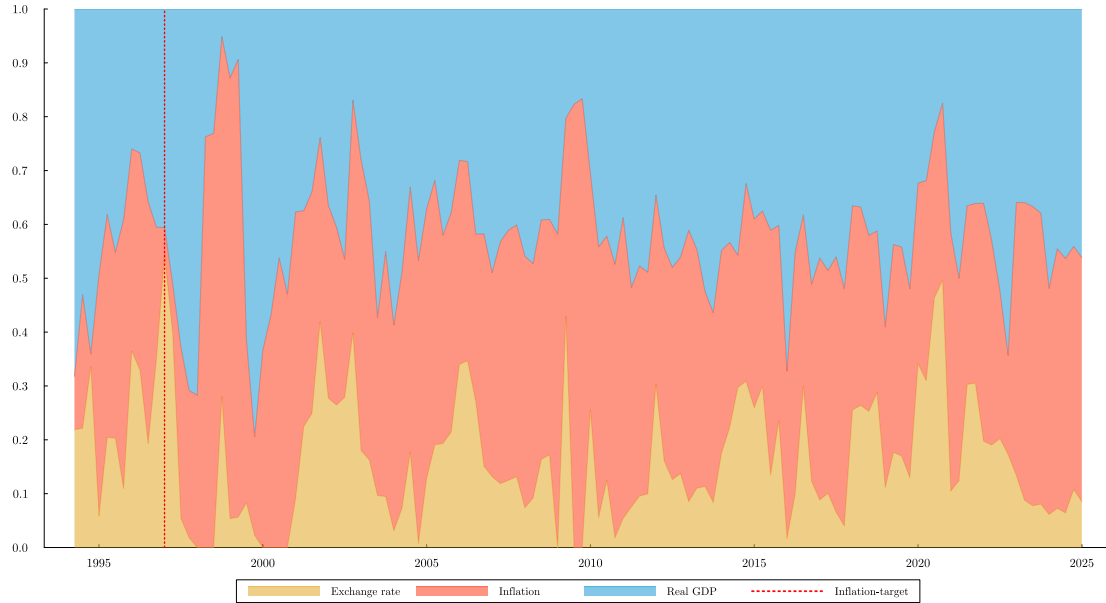
Notes: This table summarizes the recent implementation of inflation-target frameworks across 12 emerging markets. Information on the official adoption year, target type, and current target level(s) is obtained from central Banks’ official websites. *Current level(s)* refers to the most up-to-date announced inflation target, recognizing that many central Banks operate under flexible inflation-target regimes in which target bands or point targets are revised and announced several years in advance (for example, the Philippines typically announces revisions two years ahead). *Actual* inflation is measured as the 2024 average of quarterly year-on-year log changes in the consumer price index (CPI) using data from the IMF International Financial Statistics. Inflation expectations are taken from the Economist Intelligence Unit (EIU). *Forecast (1yr)* reports the EIU forecast for calendar-year CPI inflation in 2025, while *Forecast (avg)* reports the average of EIU forecasts for calendar-year CPI inflation over the 2025–2029 period, capturing medium-run inflation expectations. †: For Thailand and Turkey, actual inflation is computed using the latest available CPI data in 2024 (2023Q4–2024Q3). Except for Turkey, EIU forecasts are taken from the October 2024 EIU country forecast reports; for Turkey, forecasts are from the July 2024 report due to data availability.

Table 2.2: Monetary instruments under inflation-target in emerging markets

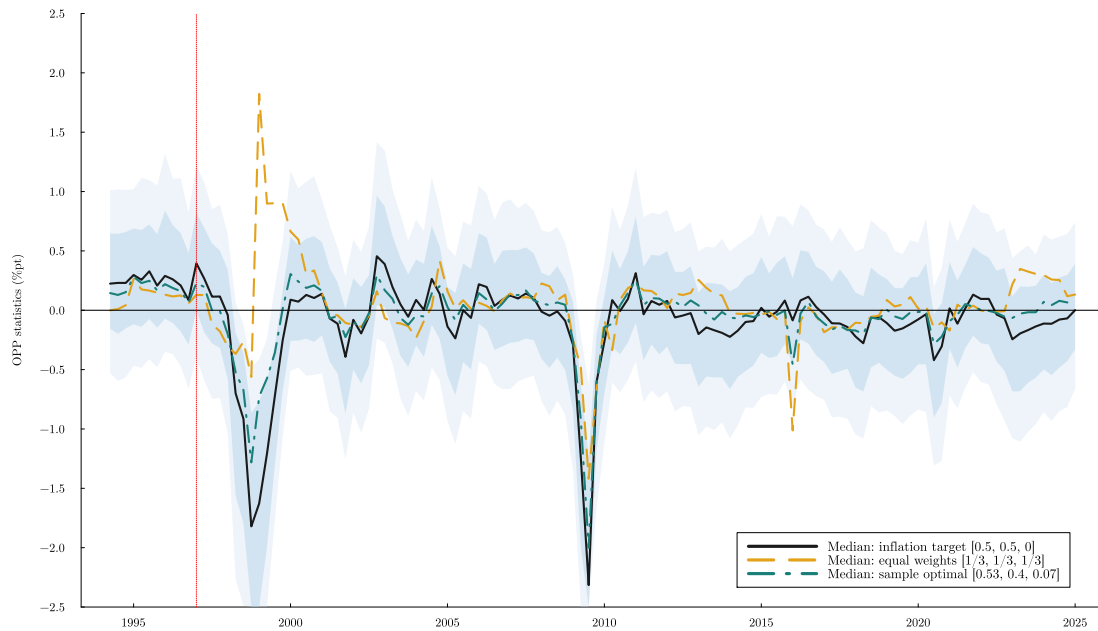
Country	Policy Instrument	Target Facility	Maturity	Secured	Policy Rate (%)	Stabilization Objectives		
						Growth	Price	EXR
(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
BRA	SELIC Rate	Interbank	O/N	✓	12.4 (5.0)		✓	
CHL	MP Rate	Interbank	O/N		4.2 (2.4)		✓	
COL	MP Rate	Interbank	O/N		6.6 (3.3)	✓	✓	
IDN	BI Rate	Central Bank	7-day	✓	6.6 (2.2)		✓	✓
ISR	SHIR Rate	Interbank	O/N		4.9 (4.8)		✓	
KOR	BOK Base Rate	Central Bank	7-day	✓	2.9 (1.4)		✓	
MEX	Target Rate	Interbank	O/N		6.1 (2.5)		✓	
PER	Reference Rate	Interbank	O/N		3.7 (1.7)		✓	
PHL	Reverse Repo Rate	Central Bank	Pre-specified	✓	4.7 (2.0)		✓	
THA	1-day Repo Rate	Central Bank	O/N	✓	2.1 (1.1)	✓		
TUR	7-day Repo Rate	Central Bank	7-day	✓	13.4 (11.3)		✓	
ZAF	Repo Rate	“Tiered-floor System”	O/N	✓	7.2 (2.3)	✓	✓	

Notes: This table summarizes heterogeneity in the operational implementation of monetary policy across emerging markets, focusing on the target facility, maturity, and collateralization of policy instruments, relative to the Federal Funds Rate, which is defined as the target overnight unsecured interbank borrowing rate in the United States. *Policy Instrument* reports the official name of each central bank’s monetary policy target rate. *Target Facility* indicates whether monetary policy implementation operates through the interbank market (i.e., transactions among commercial banks via open market operations) or directly through borrowing from or lending to the central bank (e.g., repurchase or reverse repurchase facilities). *Maturity* refers to the tenor of the instrument targeted by the central bank, with “O/N” denoting overnight maturity. *Secured* indicates whether transactions are collateralized by government securities. *Policy Rate (%)* reports the mean policy rate over the inflation-target period, with the standard deviation shown in parentheses; sample periods vary across countries depending on the timing of inflation-target adoption and data availability. Columns (7)–(9) summarize central banks stated macroeconomic policy objectives, based on official mission statements published on their official websites, indicating whether output stabilization, price stability, or exchange rate considerations are explicitly mentioned. All information is obtained from the most recent official central bank websites.

Figure 2.1: South Korea—Revealed policy priorities and OPP-based policy evaluation
(a) Time-varying policy weights that rationalize observed policy rate

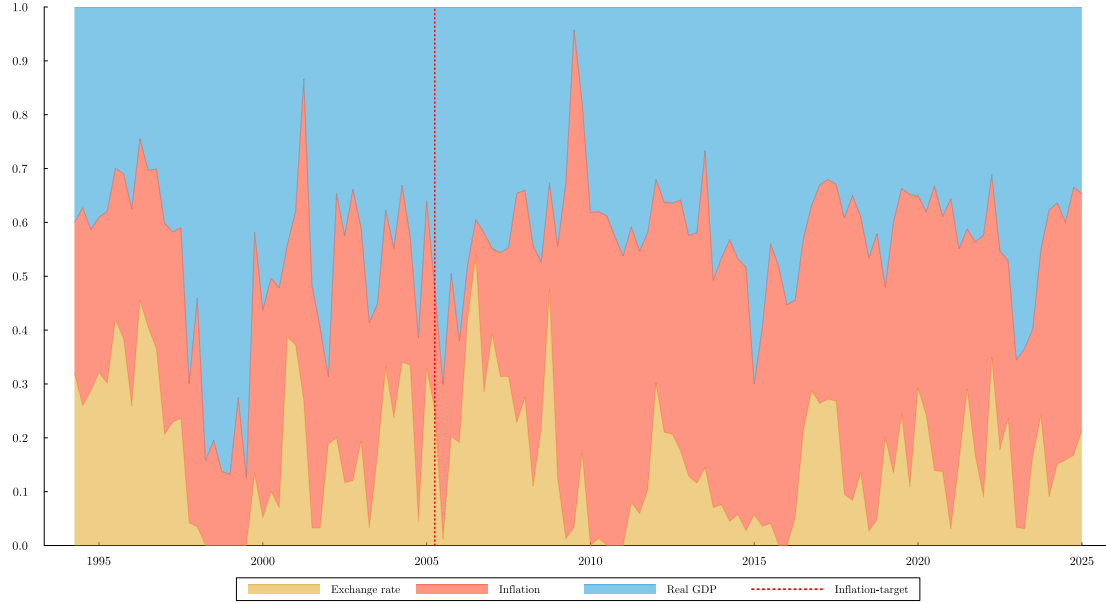


(b) OPP statistics across different constant-weight benchmarks

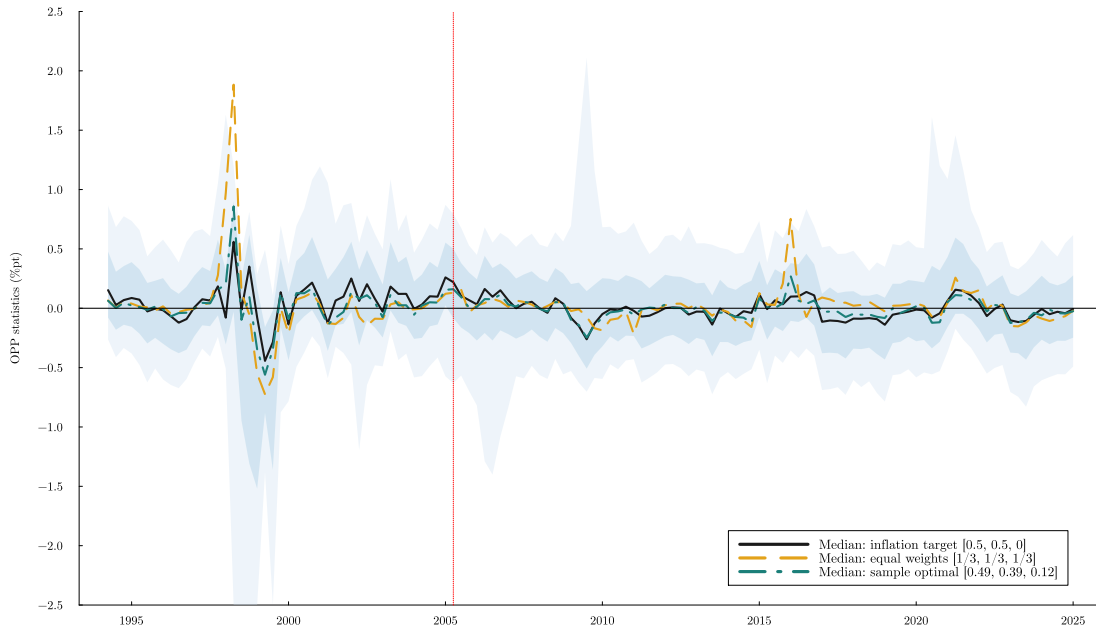


Notes: This figure evaluates the credibility of inflation-target monetary policy in South Korea. Panel (a) reports the period-specific optimal policy weights—interpreted as revealed preferences over policy trade-offs—that minimize the OPP deviation from zero in each year-quarter. Panel (b) plots the median OPP statistic under the inflation-target benchmark as well as its 68% and 90% percentile bands. For comparison, the median OPP series under the equal-weight benchmark and the global-optimal benchmark are also shown. The vertical red dotted line marks the official adoption of inflation-target in South Korea.

Figure 2.2: Indonesia—Revealed policy priorities and OPP-based policy evaluation
(a) Time-varying policy weights that rationalize observed policy rate

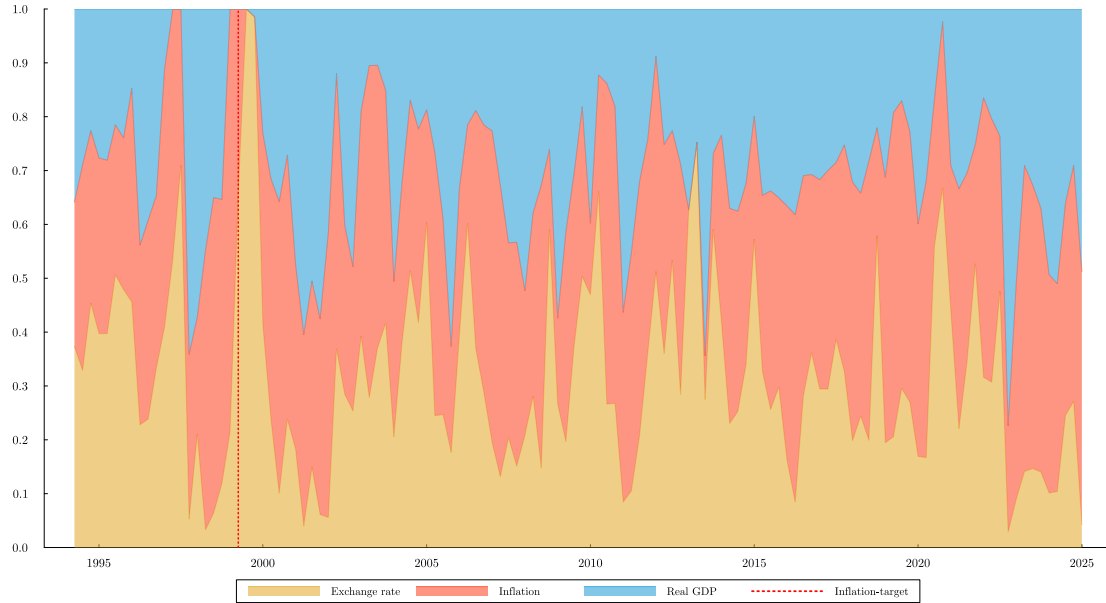


(b) OPP statistics across different constant-weight benchmarks

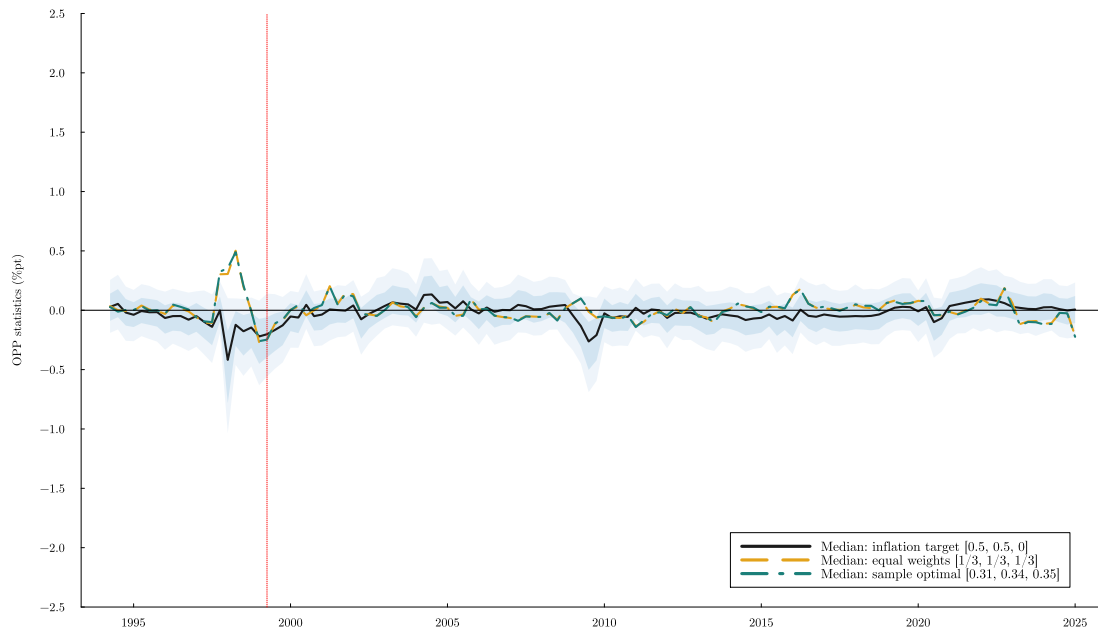


Notes: This figure evaluates the credibility of inflation-target monetary policy in Indonesia. Panel (a) reports the period-specific optimal policy weights—interpreted as revealed preferences over policy trade-offs—that minimize the OPP deviation from zero in each year-quarter. Panel (b) plots the median OPP statistic under the inflation-target benchmark as well as its 68% and 90% percentile bands. For comparison, the median OPP series under the equal-weight benchmark and the global-optimal benchmark are also shown. The vertical red dotted line marks the official adoption of inflation-target in Indonesia.

Figure 2.3: Thailand—Revealed policy priorities and OPP-based policy evaluation
(a) Time-varying policy weights that rationalize observed policy rate

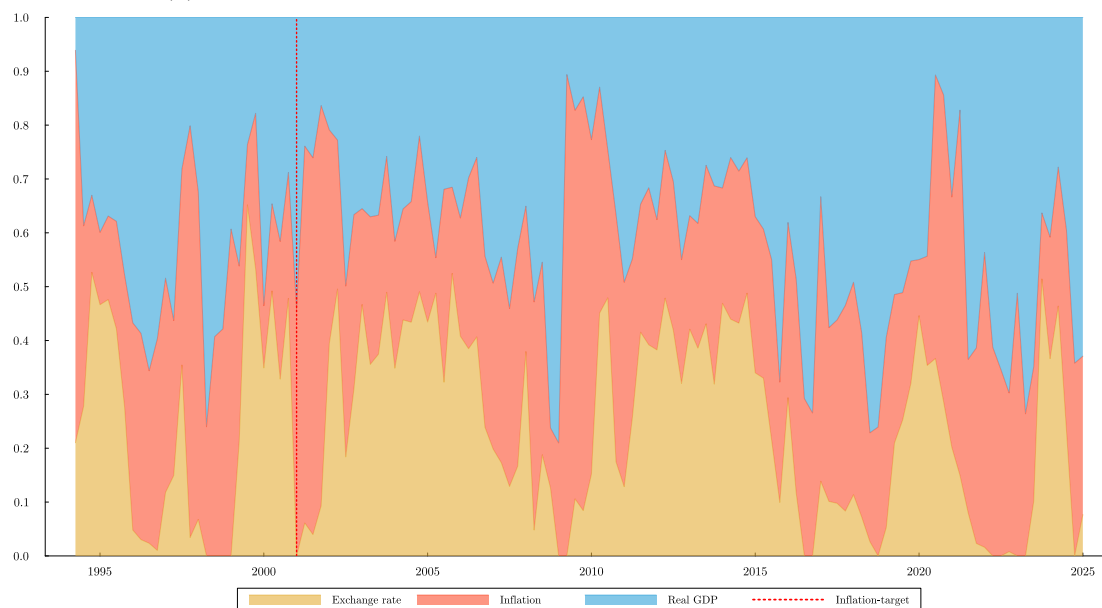


(b) OPP statistics across different constant-weight benchmarks

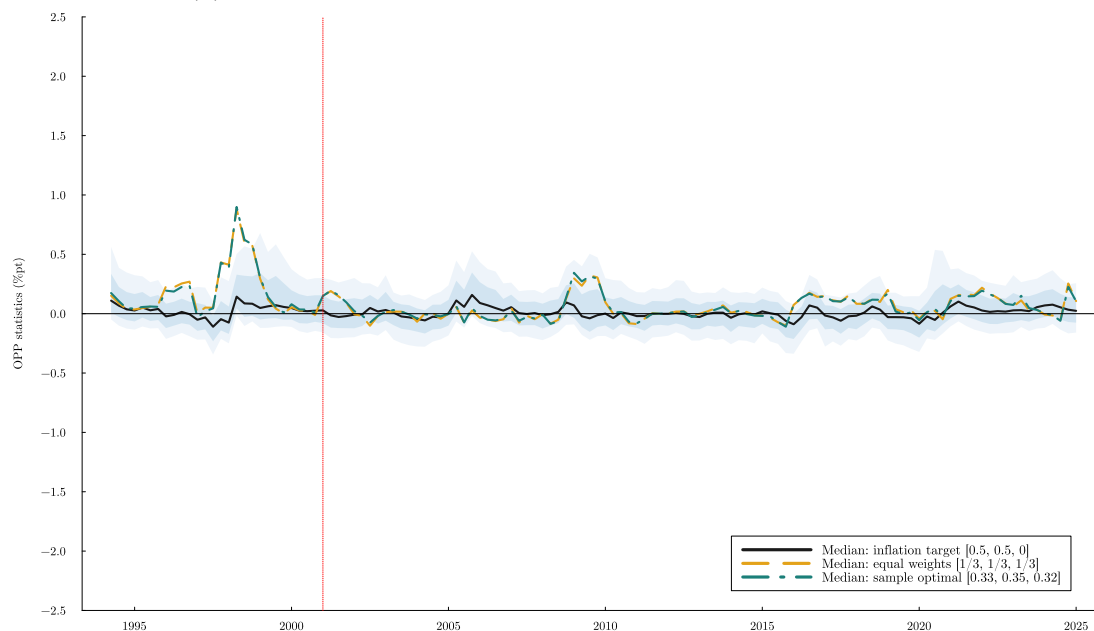


Notes: This figure evaluates the credibility of inflation-target monetary policy in Thailand. Panel (a) reports the period-specific optimal policy weights—interpreted as revealed preferences over policy trade-offs—that minimize the OPP deviation from zero in each year-quarter. Panel (b) plots the median OPP statistic under the inflation-target benchmark as well as its 68% and 90% percentile bands. For comparison, the median OPP series under the equal-weight benchmark and the global-optimal benchmark are also shown. The vertical red dotted line marks the official adoption of inflation-target in Thailand.

Figure 2.4: The Philippines—Revealed policy priorities and OPP-based policy evaluation
(a) Time-varying policy weights that rationalize observed policy rate

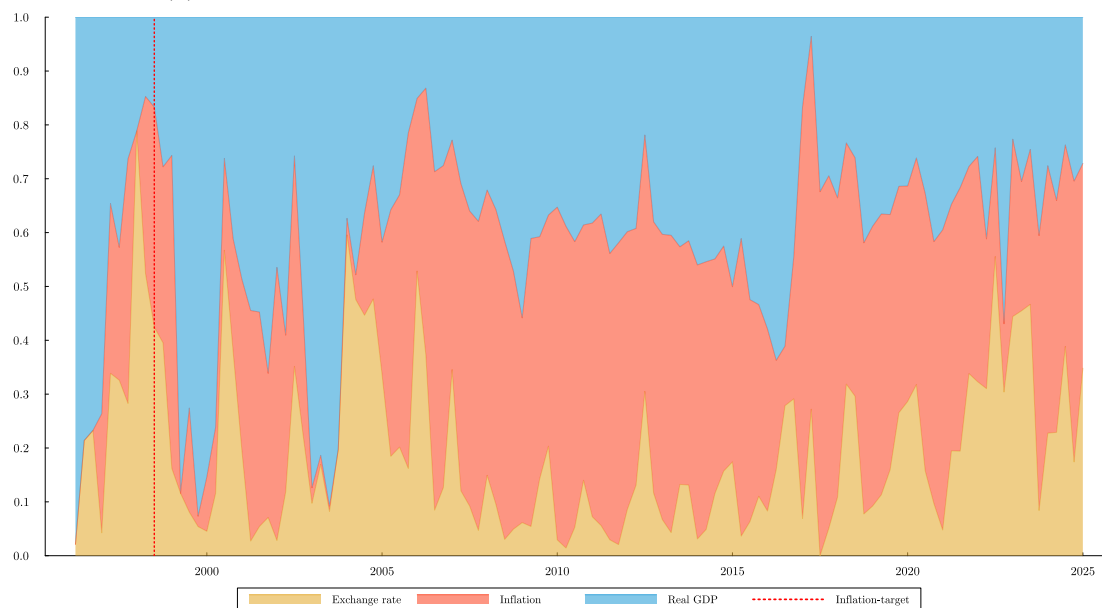


(b) OPP statistics across different constant-weight benchmarks

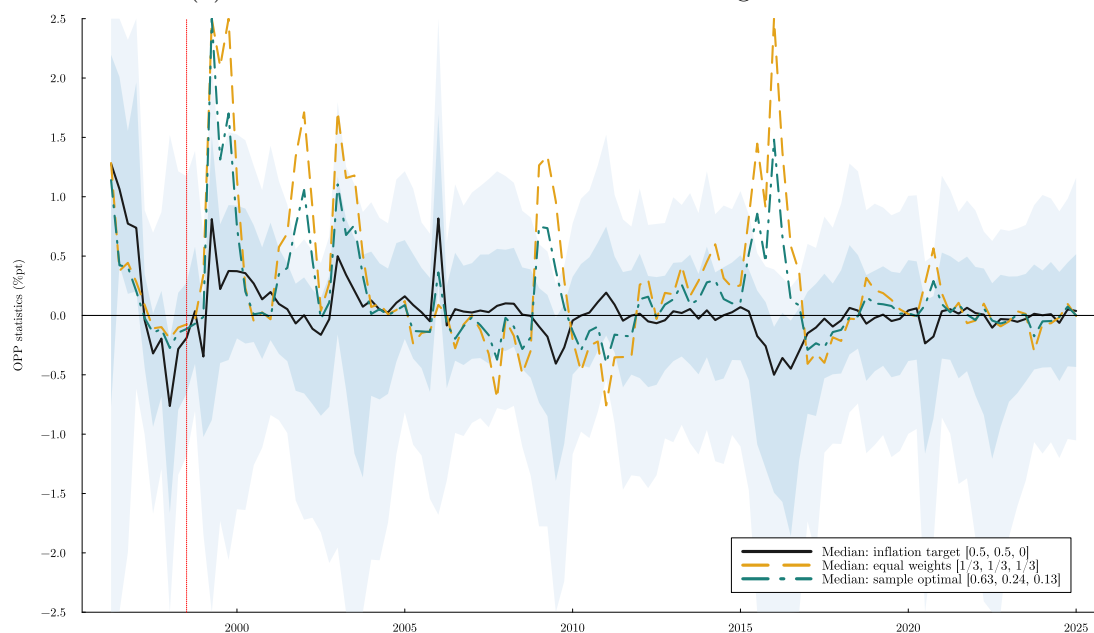


Notes: This figure evaluates the credibility of inflation-target monetary policy in The Philippines. Panel (a) reports the period-specific optimal policy weights—interpreted as revealed preferences over policy trade-offs—that minimize the OPP deviation from zero in each year-quarter. Panel (b) plots the median OPP statistic under the inflation-target benchmark as well as its 68% and 90% percentile bands. For comparison, the median OPP series under the equal-weight benchmark and the global-optimal benchmark are also shown. The vertical red dotted line marks the official adoption of inflation-target in The Philippines.

Figure 2.5: Brazil—Revealed policy priorities and OPP-based policy evaluation
(a) Time-varying policy weights that rationalize observed policy rate

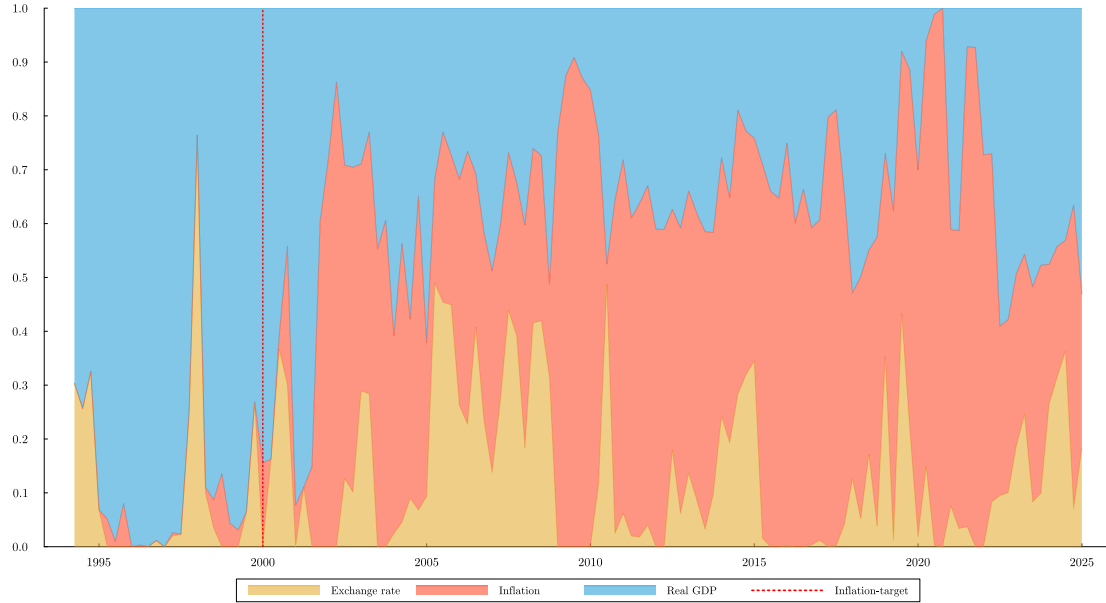


(b) OPP statistics across different constant-weight benchmarks

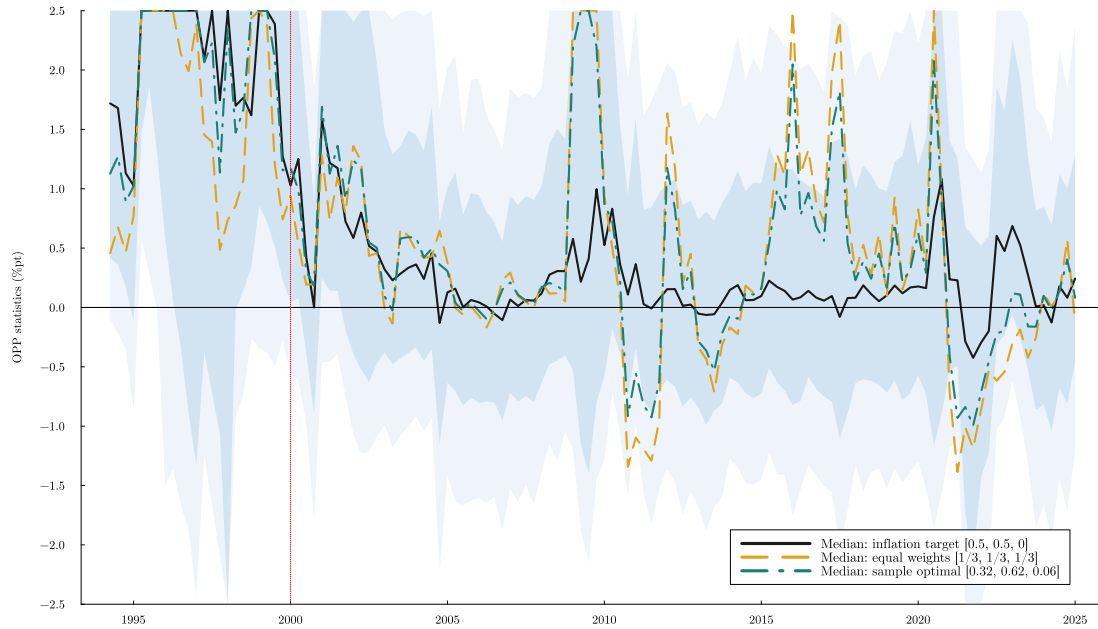


Notes: This figure evaluates the credibility of inflation-target monetary policy in Brazil. Panel (a) reports the period-specific optimal policy weights—interpreted as revealed preferences over policy trade-offs—that minimize the OPP deviation from zero in each year-quarter. Panel (b) plots the median OPP statistic under the inflation-target benchmark as well as its 68% and 90% percentile bands. For comparison, the median OPP series under the equal-weight benchmark and the global-optimal benchmark are also shown. The vertical red dotted line marks the official adoption of inflation-target in Brazil.

Figure 2.6: Mexico—Revealed policy priorities and OPP-based policy evaluation
(a) Time-varying policy weights that rationalize observed policy rate

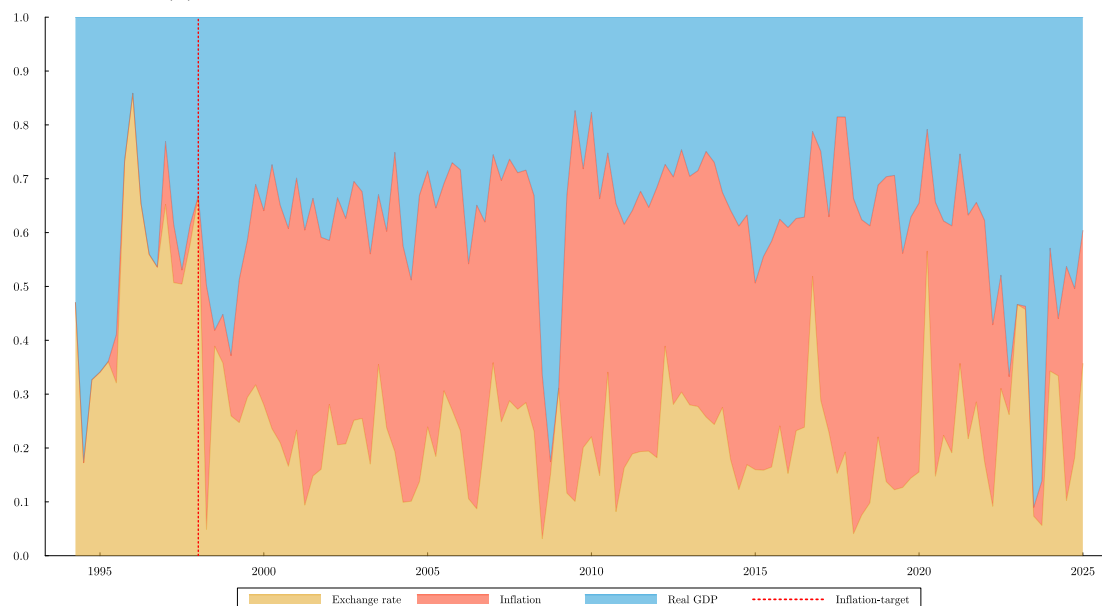


(b) OPP statistics across different constant-weight benchmarks

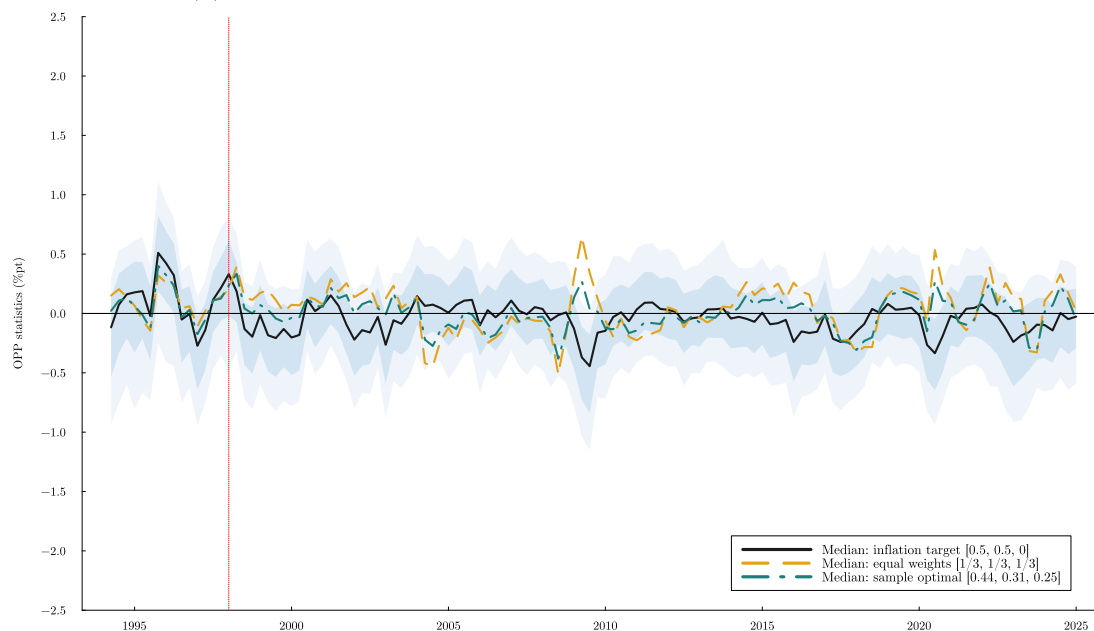


Notes: This figure evaluates the credibility of inflation-target monetary policy in Mexico. Panel (a) reports the period-specific optimal policy weights—interpreted as revealed preferences over policy trade-offs—that minimize the OPP deviation from zero in each year-quarter. Panel (b) plots the median OPP statistic under the inflation-target benchmark as well as its 68% and 90% percentile bands. For comparison, the median OPP series under the equal-weight benchmark and the global-optimal benchmark are also shown. The vertical red dotted line marks the official adoption of inflation-target in Mexico.

Figure 2.7: Chile—Revealed policy priorities and OPP-based policy evaluation
(a) Time-varying policy weights that rationalize observed policy rate

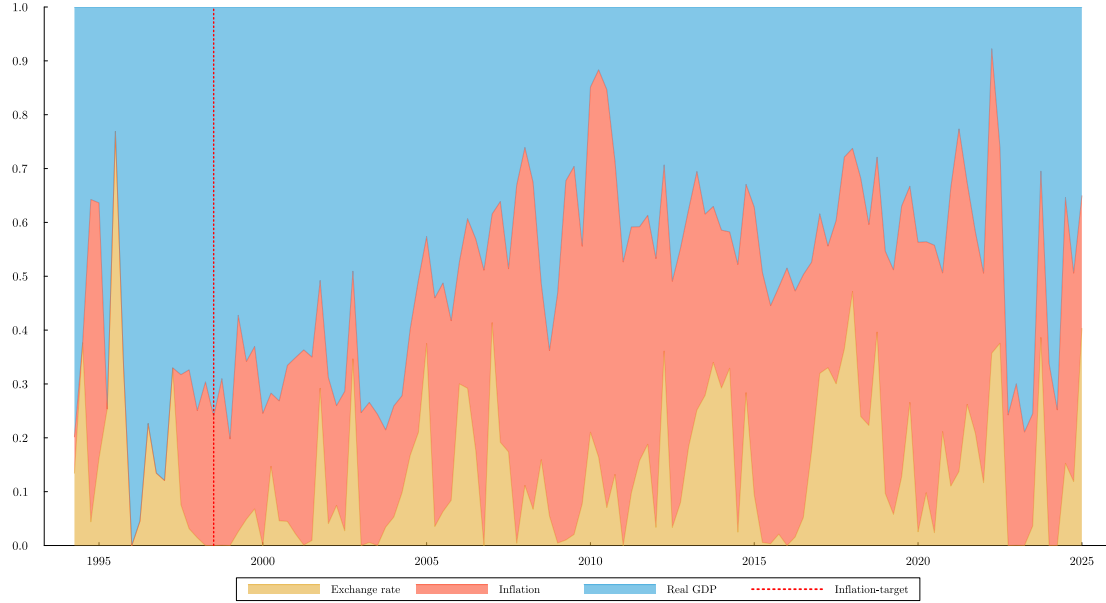


(b) OPP statistics across different constant-weight benchmarks

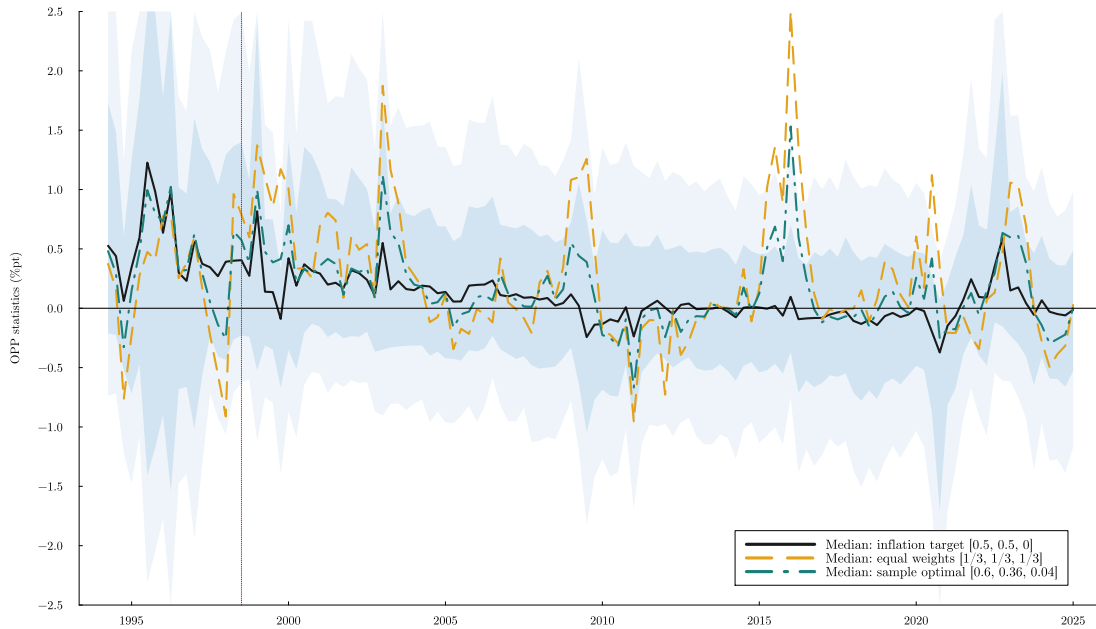


Notes: This figure evaluates the credibility of inflation-target monetary policy in Chile. Panel (a) reports the period-specific optimal policy weights—interpreted as revealed preferences over policy trade-offs—that minimize the OPP deviation from zero in each year-quarter. Panel (b) plots the median OPP statistic under the inflation-target benchmark as well as its 68% and 90% percentile bands. For comparison, the median OPP series under the equal-weight benchmark and the global-optimal benchmark are also shown. The vertical red dotted line marks the official adoption of inflation-target in Chile.

Figure 2.8: Colombia—Revealed policy priorities and OPP-based policy evaluation
(a) Time-varying policy weights that rationalize observed policy rate

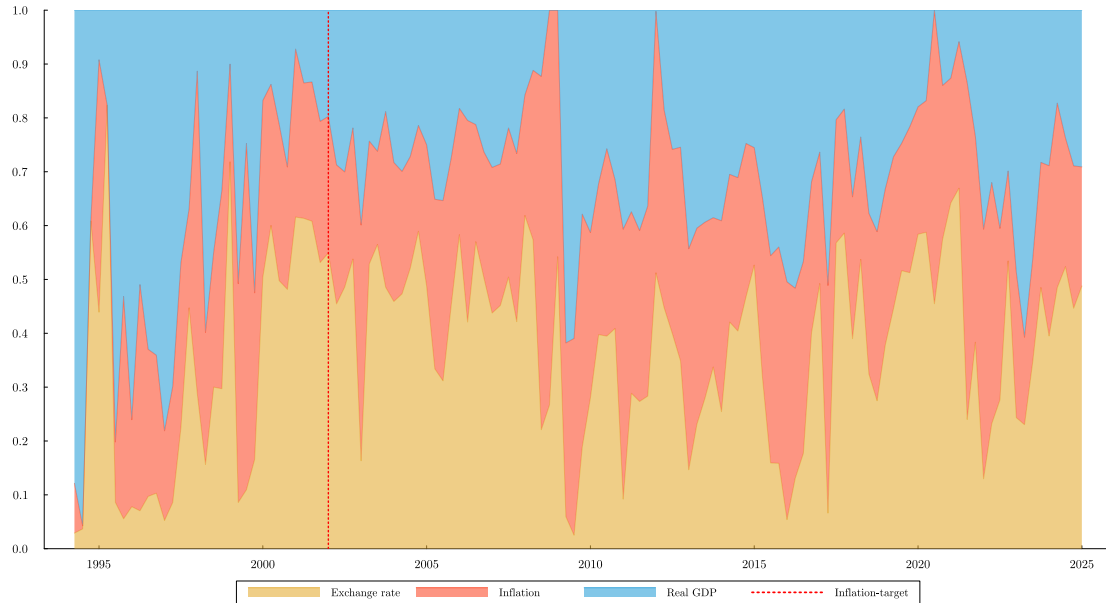


(b) OPP statistics across different constant-weight benchmarks

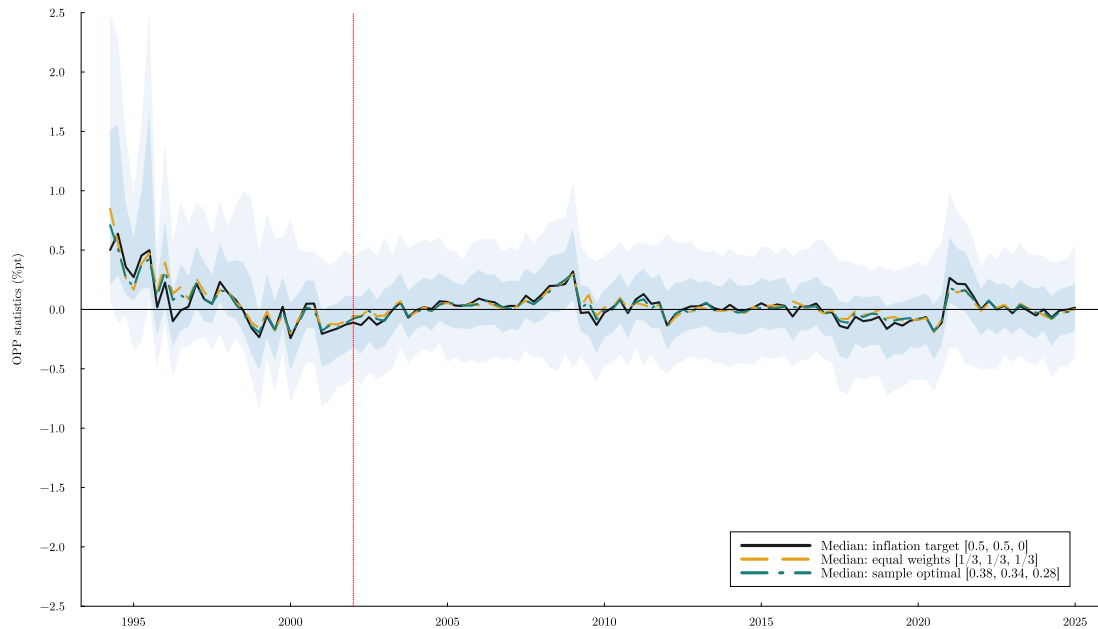


Notes: This figure evaluates the credibility of inflation-target monetary policy in Colombia. Panel (a) reports the period-specific optimal policy weights—interpreted as revealed preferences over policy trade-offs—that minimize the OPP deviation from zero in each year-quarter. Panel (b) plots the median OPP statistic under the inflation-target benchmark as well as its 68% and 90% percentile bands. For comparison, the median OPP series under the equal-weight benchmark and the global-optimal benchmark are also shown. The vertical red dotted line marks the official adoption of inflation-target in Colombia.

Figure 2.9: Peru—Revealed policy priorities and OPP-based policy evaluation
(a) Time-varying policy weights that rationalize observed policy rate

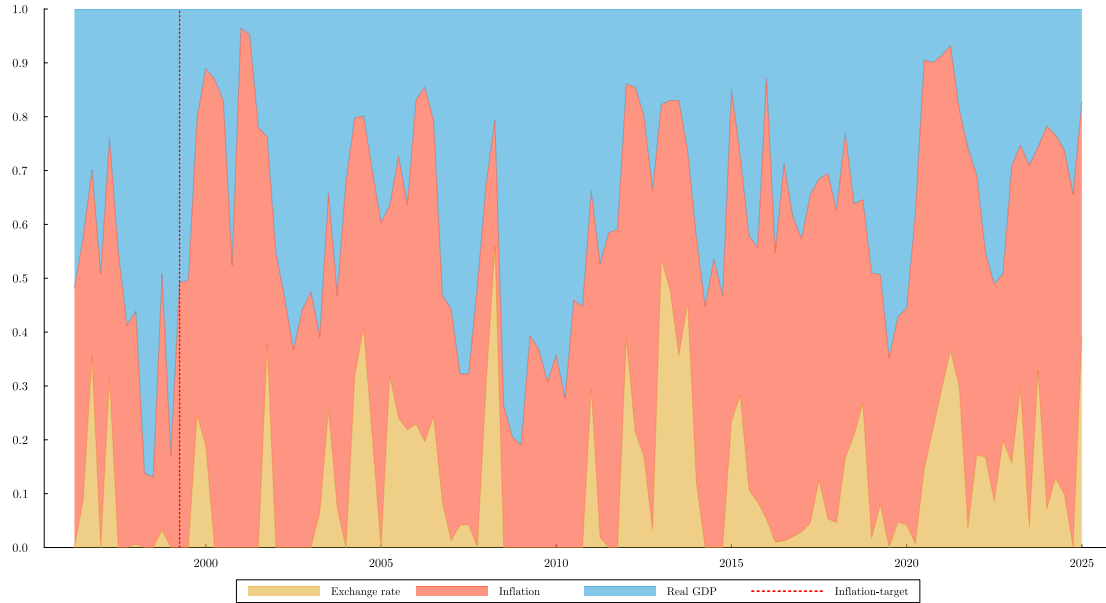


(b) OPP statistics across different constant-weight benchmarks

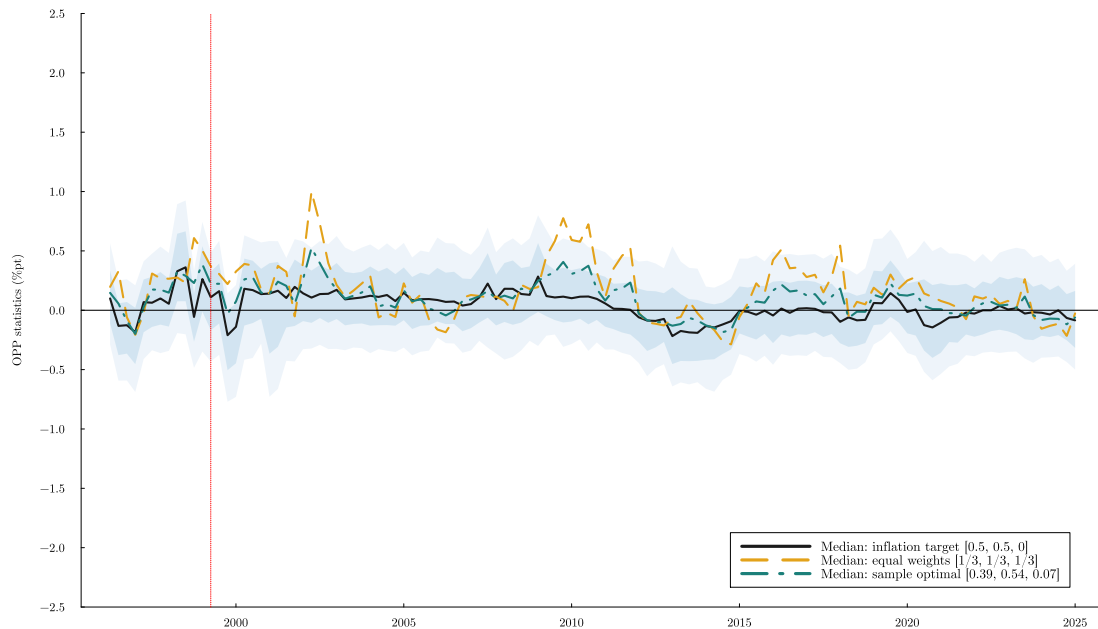


Notes: This figure evaluates the credibility of inflation-target monetary policy in Peru. Panel (a) reports the period-specific optimal policy weights—interpreted as revealed preferences over policy trade-offs—that minimize the OPP deviation from zero in each year-quarter. Panel (b) plots the median OPP statistic under the inflation-target benchmark as well as its 68% and 90% percentile bands. For comparison, the median OPP series under the equal-weight benchmark and the global-optimal benchmark are also shown. The vertical red dotted line marks the official adoption of inflation-target in Peru.

Figure 2.10: South Africa—Revealed policy priorities and OPP-based policy evaluation
(a) Time-varying policy weights that rationalize observed policy rate

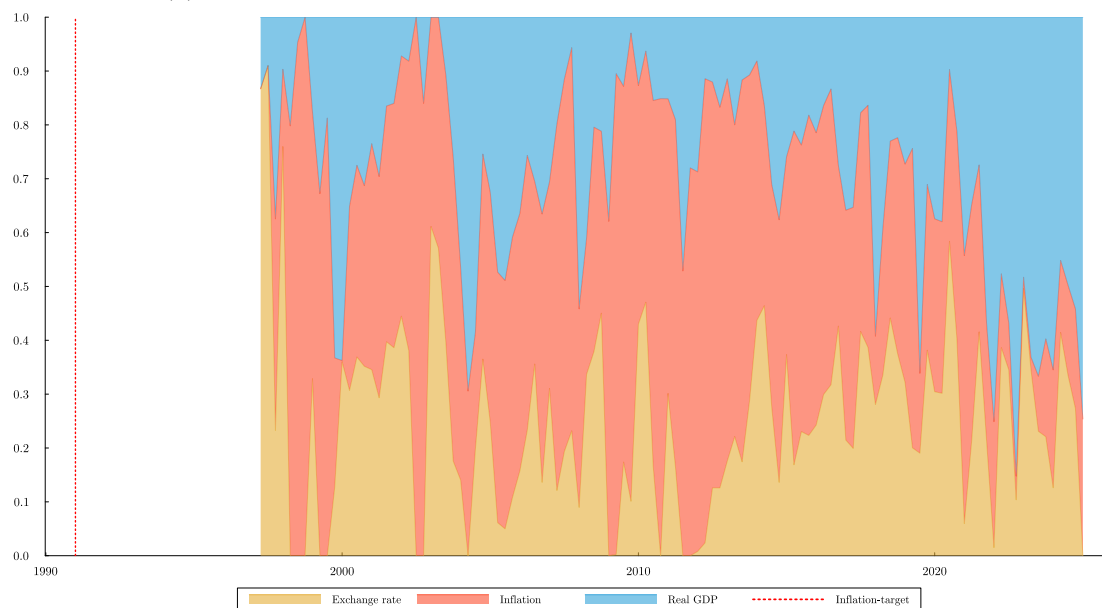


(b) OPP statistics across different constant-weight benchmarks

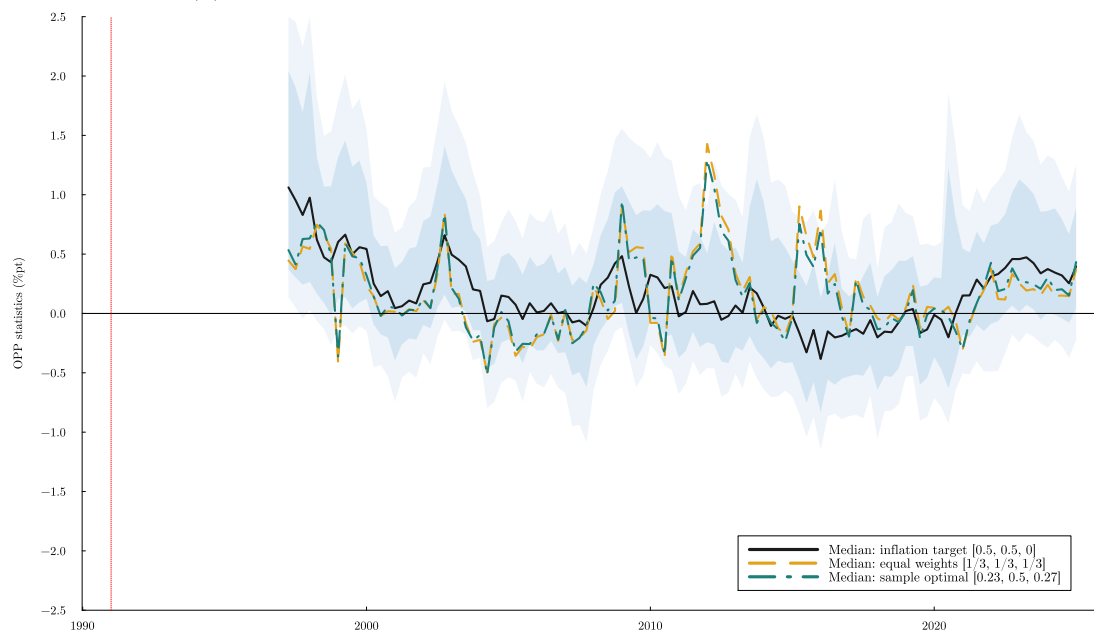


Notes: This figure evaluates the credibility of inflation-target monetary policy in South Africa. Panel (a) reports the period-specific optimal policy weights—interpreted as revealed preferences over policy trade-offs—that minimize the OPP deviation from zero in each year-quarter. Panel (b) plots the median OPP statistic under the inflation-target benchmark as well as its 68% and 90% percentile bands. For comparison, the median OPP series under the equal-weight benchmark and the global-optimal benchmark are also shown. The vertical red dotted line marks the official adoption of inflation-target in South Africa.

Figure 2.11: Israel—Revealed policy priorities and OPP-based policy evaluation
(a) Time-varying policy weights that rationalize observed policy rate

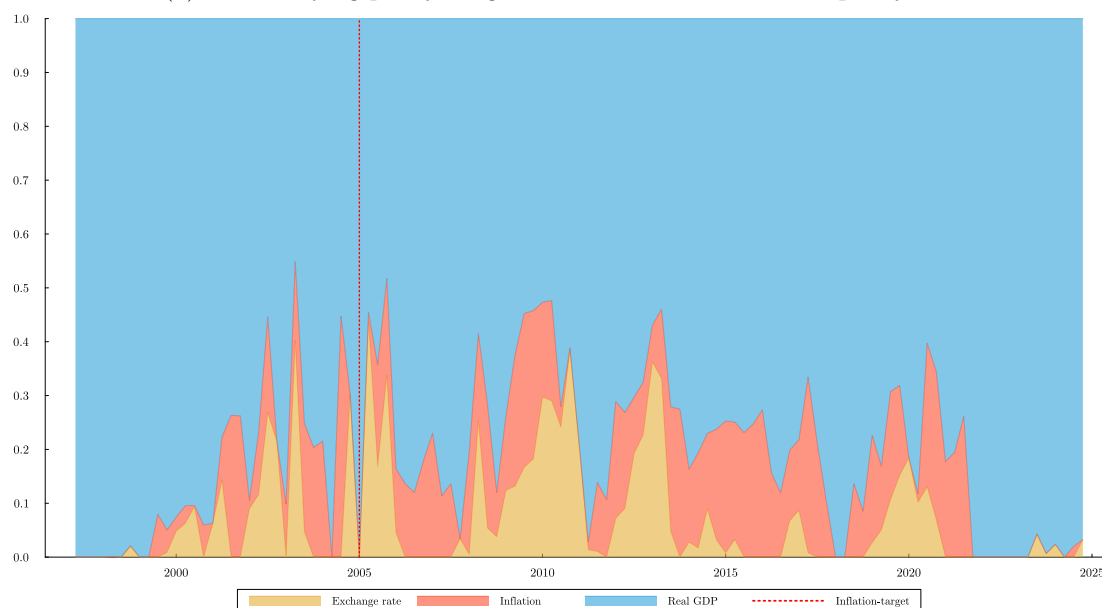


(b) OPP statistics across different constant-weight benchmarks

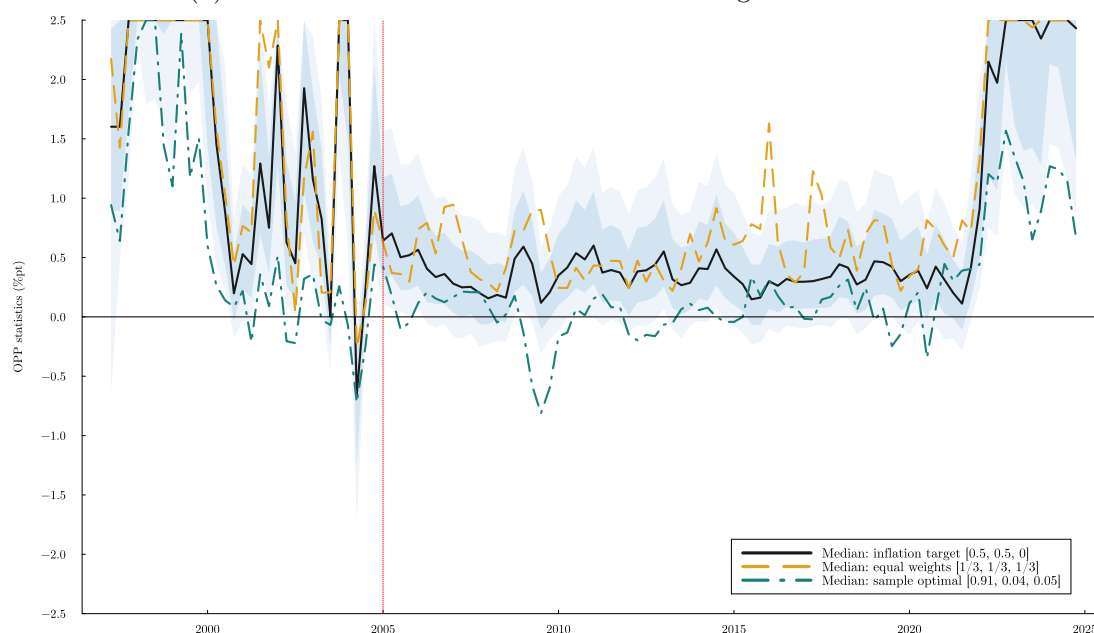


Notes: This figure evaluates the credibility of inflation-target monetary policy in Israel. Panel (a) reports the period-specific optimal policy weights—interpreted as revealed preferences over policy trade-offs—that minimize the OPP deviation from zero in each year-quarter. Panel (b) plots the median OPP statistic under the inflation-target benchmark as well as its 68% and 90% percentile bands. For comparison, the median OPP series under the equal-weight benchmark and the global-optimal benchmark are also shown. The vertical red dotted line marks the official adoption of inflation-target in Israel.

Figure 2.12: Turkey—Revealed policy priorities and OPP-based policy evaluation
(a) Time-varying policy weights that rationalize observed policy rate



(b) OPP statistics across different constant-weight benchmarks



Notes: This figure evaluates the credibility of inflation-target monetary policy in Turkey. Panel (a) reports the period-specific optimal policy weights—interpreted as revealed preferences over policy trade-offs—that minimize the OPP deviation from zero in each year-quarter. Panel (b) plots the median OPP statistic under the inflation-target benchmark as well as its 68% and 90% percentile bands. For comparison, the median OPP series under the equal-weight benchmark and the global-optimal benchmark are also shown. The vertical red dotted line marks the official adoption of inflation-target in Turkey.

2.7 APPENDIX

2.7.1 Detailed Bayesian VAR Derivation

This section provides a detailed derivation of the Bayesian VAR model used in this paper.

First, we have $\mathbf{Y} \sim \mathcal{MN}(\mathbf{XB}, \Sigma, I_{T-p})$ implies the PDF is given as:

$$f(\mathbf{Y} \mid \mathbf{XB}, \Sigma, I_{T-p}) = \left(\frac{1}{2\pi}\right)^{(T-p)k/2} \left(\frac{1}{|\Sigma|}\right)^{(T-p)/2} \exp\left(-\frac{1}{2} \text{tr}[\Sigma^{-1}(\mathbf{Y} - \mathbf{XB})'(\mathbf{Y} - \mathbf{XB})]\right)$$

At the same time, we have $\mathbf{B} \sim \mathcal{MN}(\underline{\mathbf{B}}, \underline{\mathbf{V}}, \Sigma)$ implies the PDF is given as:

$$f(\mathbf{B} \mid \underline{\mathbf{B}}, \underline{\mathbf{V}}, \Sigma) = \left(\frac{1}{2\pi}\right)^{(T-p)k/2} \left(\frac{1}{|\Sigma|}\right)^{(T-p)/2} \left(\frac{1}{|\underline{\mathbf{V}}|}\right)^{\frac{k}{2}} \exp\left(-\frac{1}{2} \text{tr}[\Sigma^{-1}(\mathbf{B} - \underline{\mathbf{B}})' \underline{\mathbf{V}}^{-1}(\mathbf{B} - \underline{\mathbf{B}})]\right)$$

Last but not least, we have $\Sigma \sim \mathcal{IW}(\underline{\mathbf{S}}, \underline{\alpha})$ implies the PDF is given as:

$$f(\Sigma) = \left(\frac{|\underline{\mathbf{S}}|^{\underline{\alpha}}}{\Gamma(\underline{\alpha})}\right) (|\Sigma|)^{-\left(\underline{\alpha} + \frac{T+1}{2}\right)} \exp(-\text{tr}[\Sigma^{-1}\underline{\mathbf{S}}])$$

Therefore, by the Baye's theorem, we have

$$\begin{aligned}
f(\mathbf{B}, \Sigma \mid Y) &= \frac{f(\mathbf{B}, \Sigma, Y)}{f(\mathbf{Y})} \\
&\propto f(\mathbf{Y} \mid \mathbf{B}, \Sigma) f(\mathbf{B} \mid \Sigma) f(\Sigma) \\
&\propto \left(\frac{1}{2\pi}\right)^{(T-p)k/2} \left(\frac{1}{|\Sigma|}\right)^{(T-p)/2} \exp\left(-\frac{1}{2} \text{tr}[\Sigma^{-1}(\mathbf{Y} - \mathbf{XB})'(\mathbf{Y} - \mathbf{XB})]\right) \\
&\quad \times \left(\frac{1}{2\pi}\right)^{(T-p)k/2} \left(\frac{1}{|\Sigma|}\right)^{(T-p)/2} \left(\frac{1}{|\mathbf{V}|}\right)^{\frac{k}{2}} \\
&\quad \times \exp\left(-\frac{1}{2} \text{tr}[\Sigma^{-1}(\mathbf{B} - \underline{\mathbf{B}})' \underline{\mathbf{V}}^{-1}(\mathbf{B} - \underline{\mathbf{B}})]\right) \\
&\quad \times \left(\frac{|\underline{\mathbf{S}}|^{\alpha}}{\Gamma(\underline{\alpha})}\right) (|\Sigma|)^{-\left(\alpha + \frac{(T-p)+1}{2}\right)} \\
&\quad \times \exp(-\text{tr}[\Sigma^{-1} \underline{\mathbf{S}}]) \\
&\propto \left(|\Sigma|^{-\left(\alpha + \frac{(T-p)+1}{2} + \frac{(T-p)}{2} + \frac{(T-p)}{2}\right)}\right) \\
&\quad \exp\left(-\frac{1}{2} \text{tr}\left(\Sigma^{-1} ((\mathbf{Y} - \mathbf{XB})'(\mathbf{Y} - \mathbf{XB}) + (\mathbf{B} - \underline{\mathbf{B}})' \underline{\mathbf{V}}^{-1}(\mathbf{B} - \underline{\mathbf{B}}) + 2\underline{\mathbf{S}})\right)\right)
\end{aligned}$$

Notice that

$$\begin{aligned}
&(\mathbf{Y} - \mathbf{XB})'(\mathbf{Y} - \mathbf{XB}) + (\mathbf{B} - \underline{\mathbf{B}})' \underline{\mathbf{V}}^{-1}(\mathbf{B} - \underline{\mathbf{B}}) \\
&= \mathbf{Y}'\mathbf{Y} - 2\mathbf{B}'\mathbf{X}'\mathbf{Y} + \mathbf{B}'\mathbf{X}'\mathbf{XB} + \mathbf{B}'\underline{\mathbf{V}}^{-1}\mathbf{B} - 2\underline{\mathbf{B}}'\underline{\mathbf{V}}^{-1}\mathbf{B} + \underline{\mathbf{B}}'\underline{\mathbf{V}}^{-1}\underline{\mathbf{B}} \\
&= -2\mathbf{B}'(\mathbf{X}'\mathbf{Y} + \underline{\mathbf{V}}^{-1}\underline{\mathbf{B}}) + \mathbf{B}'(\mathbf{X}'\mathbf{X} + \underline{\mathbf{V}}^{-1})\mathbf{B} + \mathbf{Y}'\mathbf{Y} + \underline{\mathbf{B}}'\underline{\mathbf{V}}^{-1}\underline{\mathbf{B}}
\end{aligned}$$

Using guess and verify, assume $\bar{\mathbf{V}} = (\mathbf{X}'\mathbf{X} + \underline{\mathbf{V}}^{-1})^{-1}$ and $\bar{\mathbf{B}} = \bar{\mathbf{V}}(\mathbf{X}'\mathbf{Y} + \underline{\mathbf{V}}^{-1}\underline{\mathbf{B}})$, we can show that

$$\begin{aligned}
(\mathbf{B} - \bar{\mathbf{B}})' \bar{\mathbf{V}}^{-1} (\mathbf{B} - \bar{\mathbf{B}}) &= \mathbf{B}' \bar{\mathbf{V}}^{-1} \mathbf{B} - 2\mathbf{B}' \bar{\mathbf{V}}^{-1} \bar{\mathbf{B}} + \bar{\mathbf{B}}' \bar{\mathbf{V}}^{-1} \bar{\mathbf{B}} \\
&= -2\mathbf{B}' \bar{\mathbf{V}}^{-1} \bar{\mathbf{V}} (\mathbf{X}' \mathbf{Y} + \underline{\mathbf{V}}^{-1} \underline{\mathbf{B}}) + \mathbf{B}' (\mathbf{X}' \mathbf{X} + \underline{\mathbf{V}}^{-1}) \mathbf{B} + \bar{\mathbf{B}}' \bar{\mathbf{V}}^{-1} \bar{\mathbf{B}} \\
&= -2\mathbf{B}' (\mathbf{X}' \mathbf{Y} + \underline{\mathbf{V}}^{-1} \bar{\mathbf{B}}) + \mathbf{B}' (\mathbf{X}' \mathbf{X} + \underline{\mathbf{V}}^{-1}) \mathbf{B} + \bar{\mathbf{B}}' \bar{\mathbf{V}}^{-1} \bar{\mathbf{B}}
\end{aligned}$$

Therefore, we can rewrite

$$\begin{aligned}
&(\mathbf{Y} - \mathbf{XB})' (\mathbf{Y} - \mathbf{XB}) + (\mathbf{B} - \underline{\mathbf{B}})' \underline{\mathbf{V}}^{-1} (\mathbf{B} - \underline{\mathbf{B}}) \\
&= (\mathbf{B} - \bar{\mathbf{B}})' \bar{\mathbf{V}}^{-1} (\mathbf{B} - \bar{\mathbf{B}}) - \left(-\bar{\mathbf{B}}' \bar{\mathbf{V}}^{-1} \bar{\mathbf{B}} + \mathbf{Y}' \mathbf{Y} + \underline{\mathbf{B}}' \underline{\mathbf{V}}^{-1} \underline{\mathbf{B}} \right)
\end{aligned}$$

For the last step, notice that

$$\begin{aligned}
f(\mathbf{B}, \Sigma \mid Y) &\propto (|\Sigma|)^{-\frac{(T-p)}{2}} \exp \left(-\frac{1}{2} \text{tr} \left[\Sigma^{-1} (\mathbf{B} - \bar{\mathbf{B}})' \bar{\mathbf{V}}^{-1} (\mathbf{B} - \bar{\mathbf{B}}) \right] \right) \\
&\times (|\Sigma|)^{-\left(\bar{\alpha} + \frac{T+1}{2} \right)} \exp \left(-\text{tr} \left[\Sigma^{-1} \left(\underline{\mathbf{S}} + \frac{1}{2} \left(\mathbf{Y}' \mathbf{Y} + \underline{\mathbf{B}}' \underline{\mathbf{V}}^{-1} \underline{\mathbf{B}} - \bar{\mathbf{B}}' \bar{\mathbf{V}}^{-1} \bar{\mathbf{B}} \right) \right) \right] \right) \\
&\left(|\Sigma|^{-\left(\underline{\alpha} + \frac{(T-p)+1}{2} + \frac{(T-p)}{2} + \frac{(T-p)}{2} \right)} \right) \\
&\frac{\quad}{(|\Sigma|)^{-\left(\bar{\alpha} + \frac{(T-p)+1}{2} + \frac{(T-p)}{2} \right)}}
\end{aligned}$$

Therefore, for the posterior parameters that follow the Normal-Inverse-Wishart distribution, I further assume:

$$\bar{\mathbf{S}} \equiv \underline{\mathbf{S}} + \frac{1}{2} \left(\mathbf{Y}' \mathbf{Y} + \underline{\mathbf{B}}' \underline{\mathbf{V}}^{-1} \underline{\mathbf{B}} - \bar{\mathbf{B}}' \bar{\mathbf{V}}^{-1} \bar{\mathbf{B}} \right)$$

$$\text{As well as for } \frac{\left(|\Sigma|^{-\left(\underline{\alpha} + \frac{(T-p)+1}{2} + \frac{(T-p)}{2} + \frac{(T-p)}{2} \right)} \right)}{(|\Sigma|)^{-\left(\bar{\alpha} + \frac{(T-p)+1}{2} + \frac{(T-p)}{2} \right)}} \text{ to be equal to 1, this implies}$$

$$\bar{\alpha} = \underline{\alpha} + \frac{T - p}{2}$$

Therefore, the posterior parameters $(\bar{\mathbf{B}}, \bar{\mathbf{V}}, \bar{\mathbf{S}}, \bar{\alpha})$ given as follows:

$$\bar{\mathbf{V}} = (\mathbf{X}'\mathbf{X} + \underline{\mathbf{V}}^{-1})^{-1}$$

$$\bar{\mathbf{B}} = \bar{\mathbf{V}}(\mathbf{X}'\mathbf{Y} + \underline{\mathbf{V}}^{-1}\underline{\mathbf{B}})$$

$$\bar{\mathbf{S}} = \underline{\mathbf{S}} + \frac{1}{2}(\mathbf{Y}'\mathbf{Y} + \underline{\mathbf{B}}'\underline{\mathbf{V}}^{-1}\underline{\mathbf{B}} - \bar{\mathbf{B}}'\bar{\mathbf{V}}^{-1}\bar{\mathbf{B}})$$

$$\bar{\alpha} = \underline{\alpha} + T$$

2.7.2 Optimal Policy Perturbations Derivation

This section provides a detailed derivation of the optimal policy perturbations used in this paper. Again, we start with the baseline impulse response function of endogenous variables $(x_t, \pi_t, \Delta\varepsilon_t, i_t)$ as function of parameters and structural shocks $(\xi_t, \eta_t, \epsilon_t^0)$. Since we are in the monetary discretion framework, we have $\mathbb{E}_t[\pi_{t+1}] = 0$ and $\mathbb{E}_t[x_{t+1}] = 0$. Therefore, we can rewrite the baseline model as follows:

$$(IS) \quad x_t = -\frac{1}{\sigma}i_t$$

$$(PC) \quad \pi_t = \kappa x_t + \gamma \Delta\varepsilon_t + \xi_t$$

$$(UIP) \quad \Delta\varepsilon_t = i_t + \eta_t$$

$$(MP^0) \quad i_t = \phi\pi_t + \epsilon_t$$

We first start from plugging in the IS curve and UIP curve into the Phillips curve, and

then plug in the monetary policy rule into the resulting equation. This gives us:

$$\begin{aligned}
\pi_t &= -\frac{\kappa}{\sigma}i_t + \gamma(i_t + \eta_t) + \xi_t \\
&= \left(\gamma - \frac{\kappa}{\sigma}\right)\phi\pi_t + \left(\gamma - \frac{\kappa}{\sigma}\right)\epsilon_t + \gamma\eta_t + \xi_t + \epsilon_t \\
&= \frac{1}{1 - \phi\left(\gamma - \frac{\kappa}{\sigma}\right)}\xi_t + \frac{\gamma}{1 - \phi\left(\gamma - \frac{\kappa}{\sigma}\right)}\eta_t + \frac{\gamma - \frac{\kappa}{\sigma}}{1 - \phi\left(\gamma - \frac{\kappa}{\sigma}\right)}\epsilon_t
\end{aligned}$$

Similarly, we can derive the expressions for x_t , $\Delta\varepsilon_t$, and i_t using the same substitution method, as well as adding the optimal policy perturbation term ϵ_t^* to the monetary policy rule.

For the optimal policy perturbation statistics δ_t^* , to solve

$$\min_{\delta_t} L_t = \min_{\delta_t} \frac{1}{2} \mathbf{Y}_t' \mathbf{W} \mathbf{Y}_t \quad \text{s.t.} \quad \mathbf{Y}_t = \mathcal{Y}_t^0 + \mathcal{R}^0 \delta_t$$

We can rewrite the loss function as follows:

$$\begin{aligned}
\min_{\delta_t} L_t &= \min_{\delta_t} \frac{1}{2} (\mathcal{Y}_t^0 + \mathcal{R}^0 \delta_t)' \mathbf{W} (\mathcal{Y}_t^0 + \mathcal{R}^0 \delta_t) \\
&= \min_{\delta_t} \frac{1}{2} (\mathbf{Y}_t^{0'} \mathbf{W} \mathcal{Y}_t^0 + 2\delta_t' \mathcal{R}^{0'} \mathbf{W} \mathcal{Y}_t^0 + \delta_t^2 \mathcal{R}^{0'} \mathbf{W} \mathcal{R}^0)
\end{aligned}$$

Take the first order condition with respect to δ_t and set it to zero, we have:

$$\frac{1}{2} (\mathcal{R}^{0'} \mathbf{W} \mathcal{Y}_t^0) + \delta_t^* (\mathcal{R}^{0'} \mathbf{W} \mathcal{R}^0) = 0 \implies \delta_t^* = -(\mathcal{R}^{0'} \mathbf{W} \mathcal{R}^0)^{-1} (\mathcal{R}^{0'} \mathbf{W} \mathcal{Y}_t^0)$$

2.7.3 Construction of forecast path of macroeconomic variables

Therefore, for quarterly interpolation and horizon mapping, denote m as the month of the publication, where for latter years, and $q(m)$ be the function of the mapping from month to quarter, in particular, $q(1) = 1, q(4) = 2, q(7) = 3, q(10) = 4$. And then, since forecast horizons are pointed to the end of the year, while publication values are

referring to the forecast horizon in different publication episodes, so I first place discrete forecast points on quarterly grid. Denote $j_f \in \{0, 1, 2, 3, 4\}$ be the forward forecast, while $j_b \in \{1, 2, 3\}$ to be the back filled placeholders. For example, $(j_f, j_b) = (4, 0)$ for $q(t) = 1$, and for $q(t) = 2$, $(j_f, j_b) = (3, 1)$, for $q(t) = 3$, $(j_f, j_b) = (2, 2)$, and last but not least, for $q(t) = 4$, $(j_f, j_b) = (1, 3)$. The issue is that, the EIU forecast data indeed provides forecast data up to 5 years ahead, for example, if the book is the January 2000, then it provides forecast data on the end of the year for 2000, 2001, 2002, 2003, 2004, 2005, so, this rich structure in the term structure provide me with many flexible tools to utilize this properties. In particular, for $q(t) = 1$, then $j_f \in \{4, 8, 12, 16, 20\}$ and $j_b = 1$, similar, for $q(t) = 2$, $j_f \in \{3, 7, 11, 15, 19\}$, $j_b = 1$, with the forecast horizon term structure, then I use the interpolation method to fill the quarterly horizon grids with the available forecast points and linearly interpolate the missing intermediate horizons. In other words, the focal points in each period are different based on the periodic period, that is, the forecast horizons that have real values (i.e., not interpolated) are different for each year-quarter.

Figure 2.13: Impulse responses to a 1%pt monetary policy shock (Bayesian VAR; Cholesky)

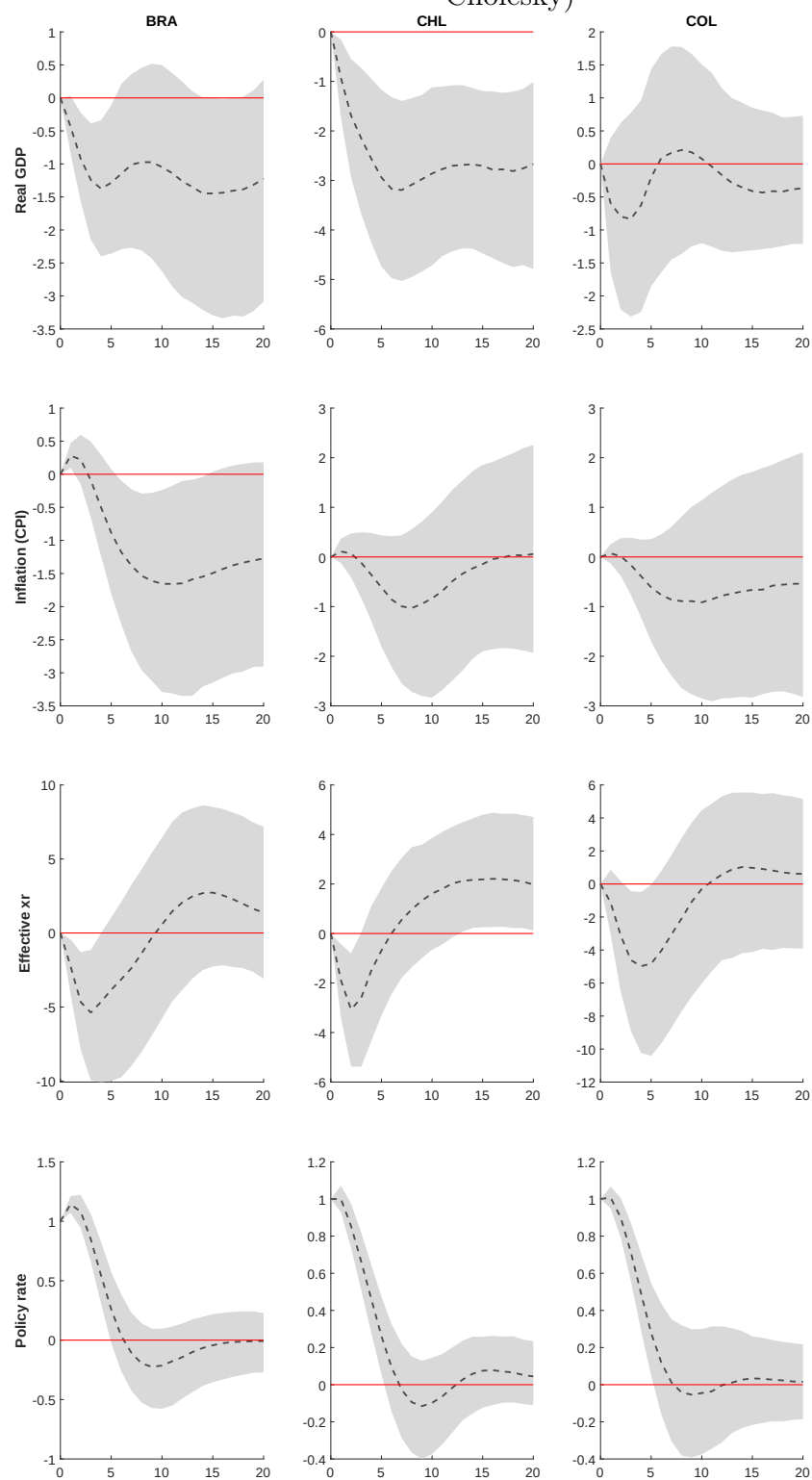


Figure 2.13: (continued)

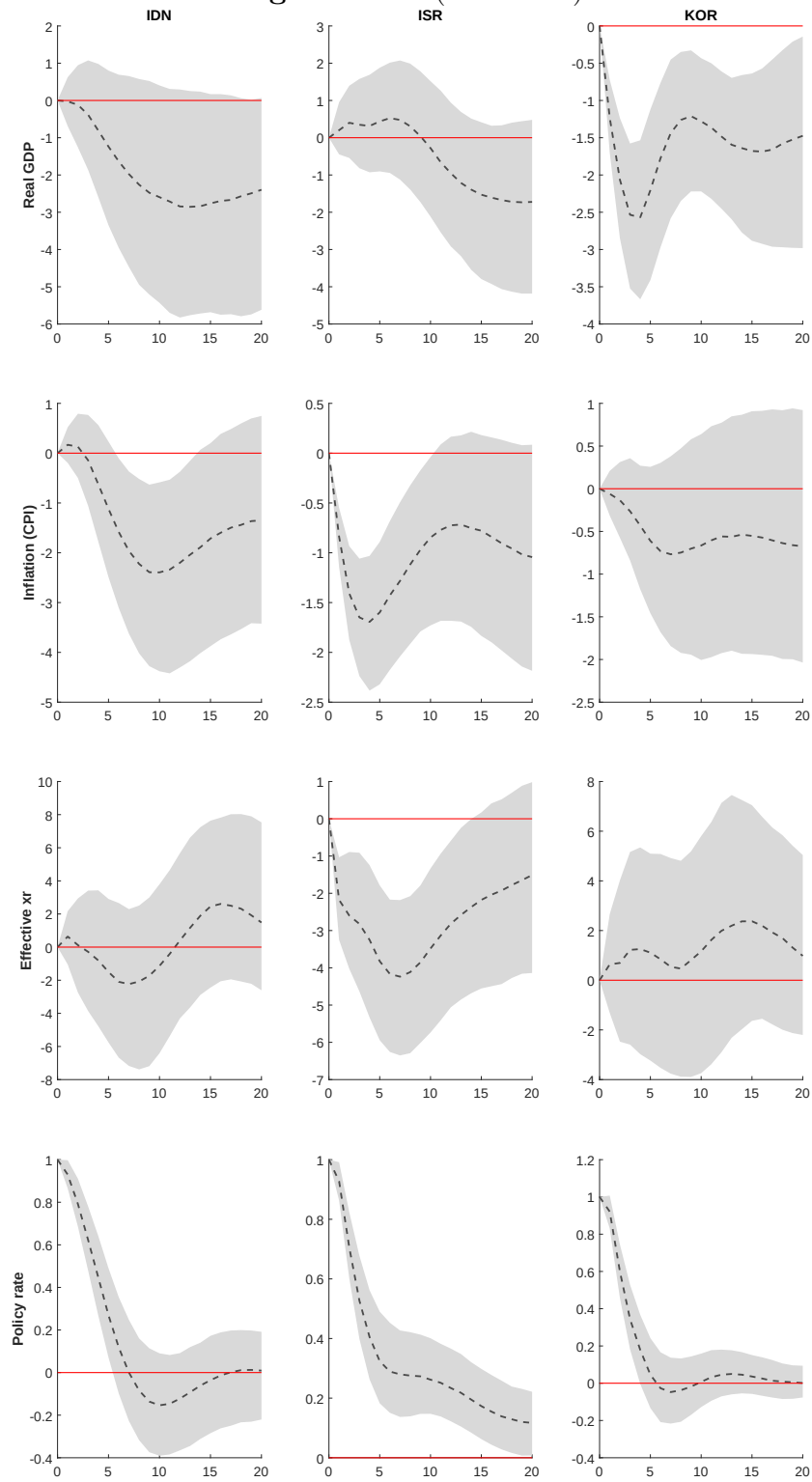


Figure 2.13: (continued)

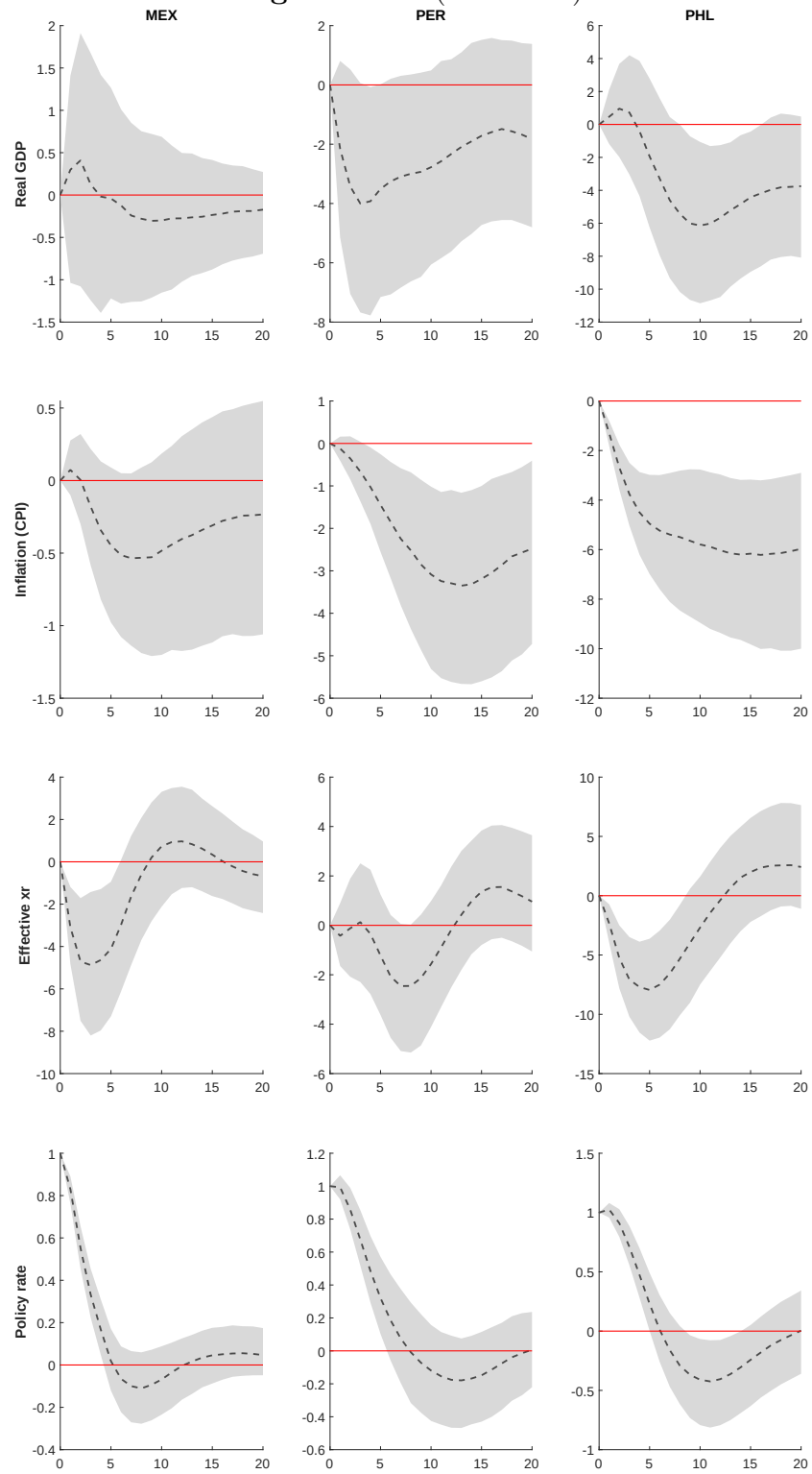
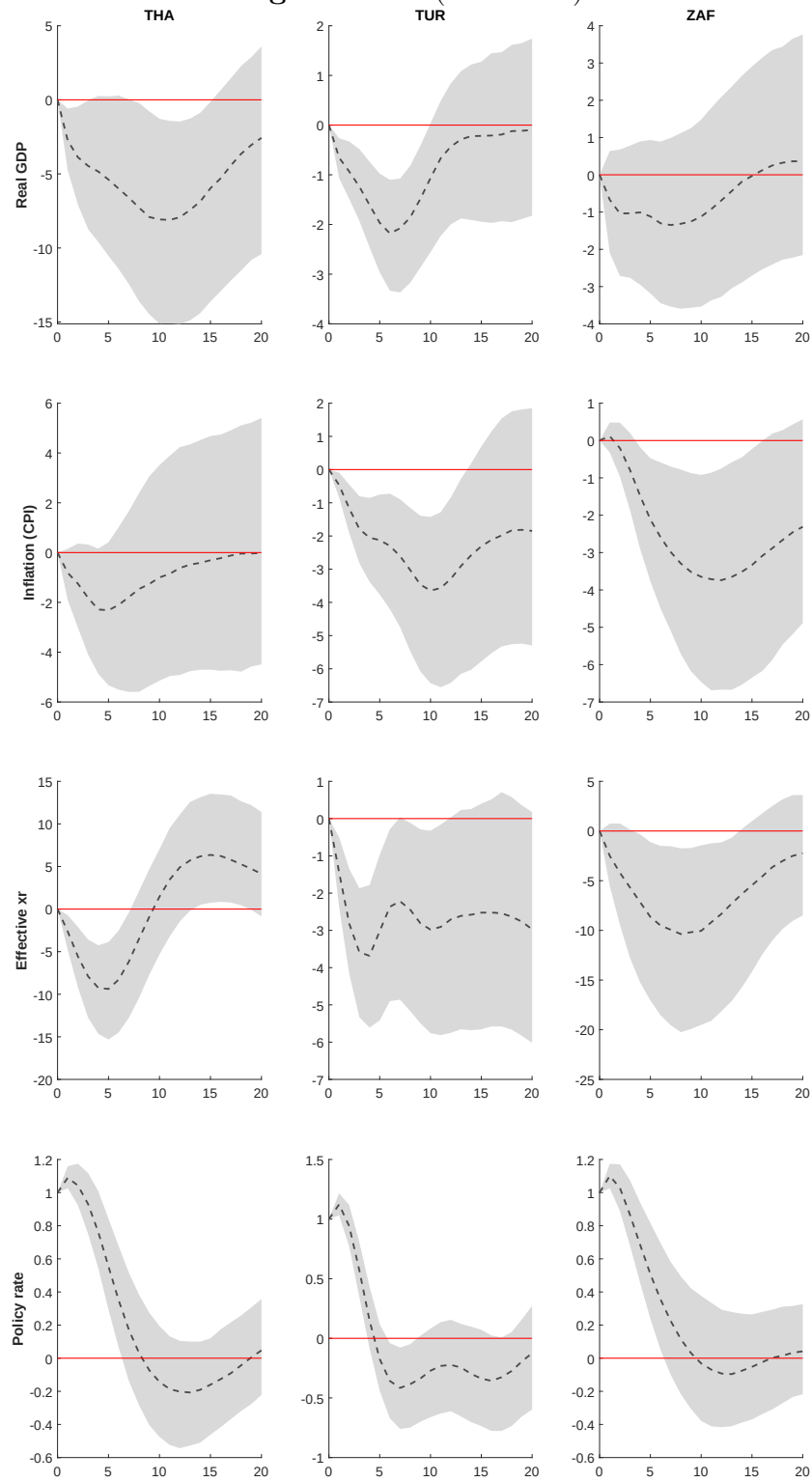


Figure 2.13: (continued)



Chapter 3

Term Structure, Interest Volatility, and the Failure of the UIP Condition

3.1 INTRODUCTION

The uncovered interest parity (UIP) condition states that the capital loss from expected depreciation offsets the yield advantage of holding a high interest currency. Empirically, when regressing the log-difference of the exchange rate Δs_{t+1} against interest rate differentials $i_t - i_t^*$,

$$\Delta s_{t+1} = \alpha + \beta(i_t - i_t^*) + e_t$$

the UIP condition implies $\alpha = 0$ and $\beta = +1$. Equivalently, if we define the excess return from holding foreign currency as $\text{rx}_{t+1} \equiv (i_t^* + \Delta s_{t+1}) - i_t$, and the foreign-to-domestic interest rate differential as $i_t^* - i_t$, regressing excess returns on interest rate differentials

$$\text{rx}_{t+1} = \alpha + \gamma(i_t^* - i_t) + u_t \tag{3.1}$$

should yield $\gamma = 0$ (i.e., $\gamma = 1 - \beta$). This no-arbitrage condition is a cornerstone of rational-expectations models and frictionless DSGE macroeconomic frameworks.

However, a large literature has repeatedly rejected this prediction—empirical studies

often find that high-interest-rate currencies tend to appreciate rather than depreciate on average (i.e., $\beta < 0$, or equivalently, $\gamma > 0$), and the regression typically exhibits very low explanatory power as measured by the R-square values. In such cases, investors holding the high interest currency earn not only the higher interest return but also an additional return from currency appreciation. This empirical pattern is commonly referred to as the *UIP puzzle*.

The literature offers two broad explanations for this puzzle. The macroeconomic approach emphasizes delayed responses of exchange rates to monetary shocks due to sticky expectations or information frictions. In this view, UIP deviations arise because exchange rates adjust sluggishly to monetary policy changes. The finance-based interpretation, by contrast, emphasizes that higher-interest-rate currencies may offer additional compensation because they are systematically exposed to greater risk. Models with time-varying risk premia—including those with Epstein-Zin preferences in long-run risk frameworks as well as rare disaster models—can generate UIP deviations that are consistent with the observed empirical patterns. Understanding the driving forces behind UIP deviations therefore helps bridge macroeconomic models with asset-pricing approaches to currency markets.

This paper introduces a new empirical perspective to this debate by studying the term structure of UIP regressions. Most standard tests of the UIP condition focus either on short-maturity assets (e.g., 1-month interest rates) or on carry-trade portfolios constructed from long-horizon assets (e.g., 10-year government bonds).¹ By contrast, this paper uses Eurocurrency/LIBOR deposit rates at the 1-, 3-, 6-, and 12-month maturities, which provide horizon-specific expected returns quoted at the same point in time. This distinction matters because traditional carry-trade portfolios examine arbitrage opportunities across consecutive periods, whereas the term structure of deposit rates captures the expected compensation from holding foreign-currency deposits over different horizons rather than the price evolution

¹By carry trade returns, the literature typically defines returns from holding zero-coupon bonds, such as $R_{t+1}^{(k)} \equiv \frac{P_{t+1}^{(k-1)}}{P_t^{(k)}}$, which capture the excess returns from sequentially holding bonds of the same maturity across periods.

of the same asset over time.

Studying the term structure of UIP regressions provides particularly useful insights during the zero lower bound period and in the recent debate surrounding the “New Fama Puzzle”—the apparent sign reversal of UIP coefficients following the 2008 Global Financial Crisis. The zero lower bound is especially relevant because conventional monetary policy operates primarily through very short-maturity assets (e.g., the federal funds rate, which is the overnight interbank borrowing rate). When short-term interest rates are pushed close to zero, however, the variance of short-term interest rate differentials collapses, mechanically reducing both the precision and magnitude of UIP coefficient estimates. As a result, the apparent sign reversal of the UIP coefficient after the 2008 Global Financial Crisis may partly reflect this econometric compression rather than a genuine structural shift in investors’ risk appetite or regulatory changes affecting currency markets.

The zero lower bound environment also provides a useful thought experiment. During normal times—when central banks can influence short-term policy rates—delayed overshooting implies that exchange rates may respond more strongly at shorter horizons than at longer horizons. As a result, short-maturity currency trades tend to command larger excess return premia. By contrast, long-horizon exchange rates tend to evolve more smoothly, consistent with the empirical observation that the UIP condition tends to hold better over longer horizons.

When short-term policy rates are compressed toward zero, however, conventional monetary policy becomes constrained, while longer-term interest rates remain positive and continue to reflect expectations about future monetary policy paths as well as compensation for macroeconomic uncertainty. The persistence of positive long-term rates, in contrast to near-zero short-term rates, suggests that risk premia may shift toward longer maturities during distressed periods in which short-term rates are constrained by the zero lower bound. Such shifts may also be reinforced by unconventional monetary policies, such as quantitative easing, which explicitly target longer-term yields and the slope of the yield curve. The

term structure of UIP coefficients therefore provides a natural way to distinguish whether UIP deviations arise primarily from delayed overshooting, time-varying risk premia, or a combination of the two.

Using maturity-specific Eurocurrency/LIBOR deposit rates for the U.S. dollar and the British pound, I find empirical evidence that UIP coefficient estimates vary systematically across both maturities and monetary regimes. The United Kingdom provides a particularly useful testing ground because its exchange rate arrangements have historically moved across pegged, banded, and floating regimes while also experiencing episodes of currency stress and stabilization. In addition, the sterling market is characterized by deep financial markets, minimal capital controls, and relatively limited direct foreign exchange intervention. These features make the bilateral dollar–sterling setting well suited for studying how monetary regimes and financial conditions shape currency excess returns.

Following the monetary identification literature, I further interpret financial regimes through the signs of the yield curve slopes in the U.S. and the U.K., examining how excess return premia vary across maturities and regimes. When the two yield curves have the same sign—either both upward sloping or both inverted—I interpret these episodes as being dominated by common global factors, such as the global financial cycle. When the yield curve signs differ across the two countries, I interpret the episode as one in which country-specific or idiosyncratic shocks play a larger role. The empirical results suggest that the largest UIP deviations occur precisely during episodes in which the two yield curves have opposite signs.

Given the hegemonic role of the U.S. dollar and the safe-asset properties of U.S. financial markets, the slope of the U.S. yield curve—often interpreted as a proxy for the state of the U.S. economy—appears to wield a particularly strong influence on UIP deviations. I find that when both the U.S. and the U.K. yield curves are upward sloping, the UIP condition tends to hold more closely. However, this interpretation should be treated with caution, as such observations are more common in the sample. When the U.S. yield curve is inverted, by contrast, excess return premia tend to be higher regardless of the sign of the U.K. yield

curve. When the U.S. yield curve slope is positive, excess return premia remain present but are smaller in magnitude than during periods with an inverted U.S. yield curve. This pattern highlights the dominant role of U.S. financial conditions and U.S. monetary policy in shaping global currency risk premia.

The central message of this paper is straightforward: the UIP puzzle cannot be fully understood through a single unconditional regression. During normal periods, UIP deviations appear to be dominated by delayed overshooting mechanisms, as reflected in larger excess returns associated with short-maturity deposits. During distressed periods when short-term policy rates are compressed toward zero, risk premia become embedded in the term structure of deposit rates. The yield curves—in particular their signs—therefore provide a concise indicator of when these relationships shift. The contribution of this paper is both conceptual and empirical: it reframes the UIP puzzle as a maturity-dependent and regime-specific phenomenon rather than a single unconditional anomaly.

3.1.1 Related Literature

Starting from the seminal [Fama \(1984\)](#) paper, the literature has proposed a range of explanations for the UIP puzzle. Broadly speaking, two major approaches have emerged. The first is the finance-based literature, including [Lustig, Roussanov and Verdelhan \(2011\)](#), [Bansal and Shaliastovich \(2013\)](#), and [Colacito and Croce \(2013\)](#), which emphasizes time-varying risk premia, Epstein-Zin preferences that separate risk aversion from the elasticity of intertemporal substitution, and the role of stochastic discount factor shocks in explaining currency returns. For example, [Lustig and Verdelhan \(2007\)](#) argue that UIP deviations arise because low interest currency hedge domestic investors against aggregate consumption shocks. [Lustig, Roussanov and Verdelhan \(2011\)](#) document that currency excess returns are predictable from the U.S. business cycle conditions. [Farhi and Gabaix \(2016\)](#) introduce a rare-disaster framework in which high interest currencies command risk premia during normal times because they are expected to depreciate in states associated with heightened uncertainty.

The second approach comes from the macroeconomic literature on delayed overshooting and information frictions, in which deviations from UIP arise because exchange rates do not adjust immediately to monetary shocks. [Engel \(2016\)](#) provides a comprehensive review of this literature and documents that UIP tends to hold more closely at longer horizons. He also notes that neither the macroeconomic nor the finance-based explanations fully account for the empirical pattern in which high interest currencies initially depreciate but subsequently appreciate at longer horizons. To reconcile this pattern, he proposes a liquidity-based explanation for exchange rate dynamics.

More recently, the apparent sign reversal of UIP coefficients following the 2008 Global Financial Crisis has generated renewed discussion, often referred to as the “New Fama Puzzle” ([Bussière et al., 2022](#)). This debate has been further examined by [Engel et al. \(2022\)](#) and [Yoon \(2025\)](#). These studies acknowledge that the compression of interest rate differentials during the zero lower bound period complicates the interpretation of UIP regressions. In particular, [Engel et al. \(2022\)](#) argue that UIP coefficients appear inherently unstable and that inflation differentials have become stronger predictors of exchange rate movements in the post-crisis period, although they cannot rule out risk-premium explanations. [Yang, Zhang and Zhang \(2025\)](#) use a Bayesian VAR approach to decompose the contribution of different shocks to UIP deviations and highlight substantial heterogeneity in the transmission mechanisms through which macroeconomic shocks affect currency excess returns.

This paper builds on these insights but advances the argument further by showing that UIP coefficients are not merely unstable—they are systematically regime-dependent. By incorporating the term structure of UIP regressions and relating it to the sign of the yield curve slope, I show that maturity variation contains valuable information about the underlying sources of UIP deviations. While the literature has explored the role of maturity and yield curves in explaining the UIP puzzle, existing work typically focuses either on very short maturities or on very long horizons. The closest related paper is [Lustig, Stathopoulos and Verdelhan \(2019\)](#), which studies the infinite-horizon holding returns of government

zero-coupon bonds. Their results suggest that reconciling UIP in the long run requires stronger assumptions about underlying asset-pricing mechanisms. By contrast, [Marin and Lloyd \(2024\)](#) examine how the magnitude of the yield curve slope affects UIP regressions across horizons and show that the UIP coefficient varies systematically with business cycle conditions.

The novelty of my approach is to focus on the *sign* of the yield curve slope—to be specific, whether the yield curve is upward sloping or inverted—because these categories naturally capture regime shifts in monetary and financial conditions. Rather than emphasizing the steepness of the yield curve, I focus on the directional signal it provides. Given the limited set of maturities available in the data (1-, 3-, 6-, and 12-month), this approach provides a transparent classification of macro-financial regimes. The consistent patterns observed across maturities reinforce the idea that UIP coefficients are closely linked to yield curve characteristics that reflect broader macroeconomic conditions.

While much of the existing literature focuses on advanced economies, there is increasing attention to UIP deviations in emerging markets. For example, [Cristi et al. \(2024\)](#), [Kalemli-Ozcan and Varela \(2024\)](#), and [Kalemli-Ozcan \(2019\)](#) show that contractionary U.S. monetary policy can generate sizable UIP premia in emerging markets, with mechanisms that appear weaker in advanced economies. The contribution of this paper is to show that even among advanced economies, macro-financial conditions play an important role in determining the sign and magnitude of UIP deviations.

The remainder of the paper proceeds as follows: Section [3.2](#) describes the empirical strategy, Section [3.3](#) discusses the main results, and Section [3.4](#) concludes.

3.2 EMPIRICAL FRAMEWORK

This section describes the empirical strategy used to test maturity-specific UIP regressions. Section [3.2.1](#) introduces the notation and data sources. Section [3.2.2](#) presents the regression

design.

3.2.1 Notations and Data Sources

The main novelty of the paper is that I extend maturity-specific U.S. dollar–British pound deposit rate series back to 1973 using monthly observations at the 1-, 3-, 6-, and 12-month maturities. For the earlier period prior to 1998, I hand-collect Eurocurrency deposit rate quotations from the historical *Financial Times* archive on the last business day of each month.² Figure A1 shows the Eurocurrency rates as reported on July 31, 1973. This figure provides a representative snapshot of bid and offer quotations (i.e., a range) gathered from market participants in the Eurocurrency market, offering a rare long historical series of short- and medium-maturity offshore interest rates.³ Since 1998, I use LIBOR rates downloaded from the Bloomberg terminal, following the standard approach in the literature.

The *Financial Times* reports Eurocurrency rates in annualized terms, so I convert them into maturity-consistent returns before constructing interest differentials. Let $i_t^{(k),\text{p.a.}}$ and $i_t^{*,(k),\text{p.a.}}$ denote the annualized Eurocurrency rate at maturity $k \in \{1, 3, 6, 12\}$ at year-month t for the U.S. and the U.K. respectively, and $i_t^{(k)}$, and $i_t^{*,(k)}$ denotes the corresponding return over the relevant holding period. Using the standard log approximation for compound interest, one can show:

$$i_t^{(k),\text{p.a.}} \approx \ln(1 + i_t^{(k),\text{p.a.}}) = \ln\left((1 + i_t^{(k)})^{\frac{12}{k}}\right) = \frac{12}{k} \ln(1 + i_t^{(k)}) \approx \frac{12}{k} i_t^{(k)} \implies i_t^{(k)} = \frac{k}{12} i_t^{(k),\text{p.a.}}$$

In other words, the maturity-consistent return is proportional to the annualized quote, scaled by $\frac{k}{12}$. This transformation allows interest rates at different maturities to be compared on a common basis.

The dollar–pound exchange rate is taken directly from the Federal Reserve’s daily series

²When a month-end observation is too blurry to read because of scanning quality, I use the preceding business day instead.

³The *Financial Times* reports quote Eurocurrency rates in fractional form: in earlier periods, quotations appear in increments of $\frac{1}{8}$, later evolving to $\frac{1}{16}$ and eventually $\frac{1}{32}$.

at the last business day of each month. I define the h period ahead depreciation rate as the log change in the exchange rate between month t and $t + h$,

$$\Delta s_{t \rightarrow t+h} \equiv \ln \left(\frac{S_{t+h}}{S_t} \right) = s_{t+h} - s_t \quad s_t \equiv \ln(S_t) \quad (3.2)$$

where S_t denotes the level of the exchange rate measured as U.S. dollars per British pound. An increase in S_t therefore represents a depreciation of the U.S. dollar against the British pound.

3.2.2 Empirical Designs

From the perspective of a U.S. investor, the excess return on a k -month sterling deposit equals the sterling deposit return plus the capital gain (loss) from the appreciation (depreciation) of the British pound (i.e., depreciation of the U.S. dollar), minus the return on a U.S. deposit of the same maturity. Under the concept of fractional discounting, this measure maps directly into the excess return interpretation of UIP regressions,

$$\text{rx}_t^{(k)} \equiv (i_t^{*,(k)} + \Delta s_{t \rightarrow t+k} - i_t^{(k)}) = \alpha_t + \beta_t^{(k)}(i_t^{*,(k)} - i_t^{(k)}) + e_t \quad (3.3)$$

Using the Eurocurrency/LIBOR rates together with the exchange rate series, I estimate rolling-window UIP regressions over 10-year windows (120 months) for each maturity, following [Engel et al. \(2022\)](#). The coefficient of interest measures the response of excess returns to the interest-rate differential at the same maturity. I am particularly interested in whether these coefficients differ across maturities and whether they vary systematically over time and across regimes. The empirical design therefore focuses on the evolution of the coefficients rather than solely on their full-sample averages.

The interpretation of the regression is straightforward: a positive coefficient $\beta_t^{(k)}$ implies that higher-interest currencies earn additional compensation beyond the interest-rate spread—coming from capital gains from appreciate of pounds. In particular, I test two hy-

potheses of interest: first, whether UIP holds across maturities, $\beta_t^{(1)} = \beta_t^{(3)} = \beta_t^{(6)} = \beta_t^{(12)} = 0$, and second, whether the UIP coefficient is stable over time, $\beta_{t_1}^{(k)} = \beta_{t_2}^{(k)} = \dots$ for all year-months t .

The coefficients $\beta_t^{(k)}$ are estimated using rolling windows of 10 years (120 months) in order to illustrate the time-varying properties of UIP deviations. The primary objective is to study the point estimates rather than the statistical significance of individual coefficients. When statistical inference is required, I compute Newey-West standard errors to account for overlapping horizons.

The empirical analysis proceeds along three dimensions. First, I relate rolling-window UIP coefficient estimates to the volatility of 1-month interest-rate differentials. Second, I compare the coefficient estimates across maturities to examine whether shorter- and longer-horizon currency trades carry different risk loadings across different regimes. Third, I split the sample according to exchange rate regimes and according to the yield curve signs. Throughout the analysis, Newey-West standard errors are used when conducting inference so that overlapping horizons do not mechanically understate uncertainty. Estimating these coefficients using rolling windows allows the data to reveal whether the UIP relationship is stable, gradually evolving, or driven by specific monetary regimes.

3.3 MAIN RESULTS AND DISCUSSIONS

The empirical results are organized around three questions. First, how do the coefficient estimates on excess return premia evolve with the volatility of interest-rate differentials over time? Section 3.3.1 shows that there is a systematic relationship along the time dimension. Second, how do these coefficients differ across maturities? Section 3.3.2 shows that excess premia are concentrated in short-maturity deposits during normal time but then shift toward long-maturity deposits near the zero lower bound. Third, do exchange rate regimes and yield curve signs (upward-sloping versus inverted) help explain the observed heterogeneity? I find

that the yield curve signs provide useful information about the time-varying properties of excess return premia. Taken together, these results suggest that the UIP puzzle is not a single static empirical fact but rather a set of evolving patterns that vary with both maturity horizon and monetary regimes.

3.3.1 Volatility and Time Evolution

The first result is a strong positive relationship between rolling-window coefficient estimates on excess return premia and the volatility of interest-rate differentials. When the volatility of interest-rate differentials is high, the coefficient estimates tend to be more positive, indicating a textbook violation of the UIP condition. When the volatility of interest-rate differentials approaches zero—especially during the zero lower bound periods when central banks worldwide synchronically lower interest rates—the coefficient estimates on excess return premia turns negative, which is consistent with UIP condition saying high-interest currency is expected to depreciation. However, they are both more extreme in magnitude and also less precisely estimated, as the volatility of the regressor collapses, so periods of low volatility must be interpreted cautiously.

This observation is particularly relevant for the recent studies that interpret the sign reversal of the UIP coefficient after the 2008 Global Financial Crisis as evidence that UIP began to hold more closely—the “New Fama Puzzle.” Figure 3.1, however, suggests a more cautious interpretation: the coefficient estimates evolve gradually alongside the decline in the volatility of interest-rate differentials, which monotonically decline over time rather than exhibiting a sharp structural break around 2008. The relevant empirical pattern is therefore a slow-moving evolution across decades rather than an abrupt post-crisis regime shift.

From an econometric perspective, this relationship also warrants careful interpretation. The OLS estimates depends on the variance of the regressor, $\hat{\beta}_t^{(k)} = \frac{\text{Cov}(\text{rx}_t^{(k)}, i_t^{*,(k)} - i_t^{(k)})}{\mathbb{V}(i_t^{*,(k)} - i_t^{(k)})}$. A collapse in volatility of the interest-rate differential (i.e., a decline in $\mathbb{V}(i_t^{*,(k)} - i_t^{(k)})$) therefore mechanically affects the estimated coefficient.

This pattern also aligns closely with the exchange rate regimes experienced by the U.K.. The path of excess return coefficient estimates moves through the currency-board period, the pre-ERM band, the ERM episode, the post-ERM period, and the later floating regime in a way that suggests regime dependence rather than simple instability. In this sense, the UIP coefficient is not merely “unstable” but appears tied to broader monetary and exchange rate conditions. Without accounting for the long-run gradual decline in interest-rate differential volatility, it is difficult to attribute the post-2008 sign reversals in UIP coefficients purely to changes in market regulations or shifts in monetary policy frameworks.

In short, the central finding of Figure 3.1 is not only the econometric identity itself but also the gradual decline in interest rate volatility over time and its alignment with monetary regimes and policy environments. This gradual evolution is the key empirical fact that the subsequent subsections seek to explain.

3.3.2 Maturity Dependence

Section 3.3.1 establishes that discussions of post-2008 UIP sign reversals must take into account the gradual decline in short-term interest-rate differential volatility over time. This observation raises the question of the mechanism behind the pattern. From a financial perspective, the volatility of interest-rate differentials provides information about the broad risk environment—except in periods when central banks compress short-term rates through monetary policy interventions. This is precisely why focusing exclusively on short-term returns (e.g., the 1-month rate) can be misleading. At the same time, the delayed overshooting literature and much of the empirical evidence suggest that the UIP condition tends to fail at short horizons but hold more closely at longer horizons. Ignoring the speculative nature of short-run currency trading—where exchange rates exhibit substantial larger short-run volatility and hence carry a risk premium—would therefore overlook an important feature of currency markets as financial assets.

Traditional monetary policy operates primarily through short-term rates. For exam-

ple, the federal funds rate is the overnight interbank borrowing rate in the U.S., while the Bank Rate is the interest rate paid on overnight deposits at the Bank of England by commercial banks. As a result, the macro interpretation—sluggish response to monetary shocks—naturally focus on short-horizon exchange rate responses. At the same time, while traditional monetary theory predicts that long-maturity assets should yield higher returns because they compensate investors for postponing consumption, the behavior of exchange rates introduces an additional dimension. Short-run exchange rates are typically more volatile than long-run exchange rates, implying that short-maturity currency trades may require larger compensation for exchange-rate risk. Long-maturity deposits, by contrast, embed smoother expectations about the path of future short rates as well as more predictable long-horizon exchange rate movements, and therefore load less heavily on short-run currency trade. Examining the term structure of UIP regressions is therefore essential for distinguishing between short-run monetary explanations and broader risk-premium interpretations for the UIP puzzle.

The main novelty of this paper is therefore to examine excess return coefficients across maturity horizons (1m, 3m, 6m, and 12m).⁴ Figure 3.2 supports this perspective. During periods of high interest-rate differential volatility, the 1-month excess-return coefficient often lies above the 12-month coefficient, indicating that short-maturity currency trades command larger premia. Around and after the Global Financial Crisis, however, this ranking weakens and in some cases reverses as short-term rates are compressed toward zero. In such environments, longer maturities become relatively more informative because they continue to reflect uncertainty about future monetary policy and macroeconomic conditions.

This interpretation is consistent with the experience of quantitative easing. Figure 3.2 also shows that, since the 2008 Global Financial Crisis, the maturity-specific UIP regressions produce less negative coefficients for long-maturity (12-month) compared with short-maturity

⁴While the literature has explored the term structure of carry trades (e.g., [Lustig, Stathopoulos and Verdelhan \(2019\)](#)), most studies focus on infinite-horizon zero-coupon bonds rather than the gradual maturity transitions documented here.

(1-month). While short-term monetary policy rates approach the zero lower bound, Long-term rates, however, typically do not fall to zero as easily as short-term rates because they embed expectations about future policy paths and compensation for macroeconomic uncertainty, as supported by the unconventional monetary policies such as quantitative easing to influence the yield curve slopes. In this sense, longer maturities become the natural location where currency risk premia remain visible when short-term interest rate differentials are artificially compressed with the concerted efforts from central banks worldwide.

Figure 3.3 further sharpens this interpretation by showing that the responsiveness of excess-return coefficients to interest-rate differential volatility declines with maturity. In other words, excess return premia become less sensitive to overall economic risk at longer maturities. Shorter maturities are more responsive to exchange rate volatility than longer maturities, implying that 1-month currency trades are more exposed to risk conditions than 12-month currency trades (with coefficient estimates of 28.5 versus 16.4). Interpreted quantitatively, a one basis point (0.01%pt) increase in the standard deviation of interest-rate differentials is associated with roughly a 0.285%pt increase in currency excess returns. This pattern is both economically intuitive and empirically important, suggesting that the term structure of UIP coefficients contains meaningful information about how investors price currency risk across horizons.

3.3.3 Yield-Curve Regimes

The previous subsection shows that excess return premia vary across maturities and riskiness states. I now extend the analysis by classifying observations according to the signs of yield-curve slopes the U.S. and the U.K.. This approach is inspired by the idea of distinguishing information and real effects using the responses of the stock market and the bond market to monetary shocks as in [Jarociński and Karadi \(2020\)](#). The yield curve slope is defined as the difference between the 12-month and 1-month Eurocurrency/LIBOR rates, measured in annualized terms, that is, $\text{sign}(i_t^{(12),\text{p.a.}} - i_t^{(1),\text{p.a.}})$ and $\text{sign}(i_t^{*,(12),\text{p.a.}} - i_t^{*,(1),\text{p.a.}})$ for each

year-month t .⁵

This classification yields two natural cases. First, when both countries have yield curves with the same sign—either both upward sloping or both inverted—I interpret the episode as being dominated by common global factors, such as global risk-on or risk-off investment behavior. When the signs differ across countries, I interpret the episode as one in which country-specific or idiosyncratic shocks play a larger role. This distinction is particularly meaningful for the U.K., which has experienced both global financial shocks and country-specific currency crises, such as the 1992 sterling crisis.

Figure 3.5 reports pooled-sample regressions that separate observations into same-sign and opposite-sign yield-curve episodes. The results are striking. When the two yield curves share the same sign, excess return coefficients are small and often not statistically different from zero, broadly consistent with UIP holding more closely. When the yield-curve signs differ, however, excess return coefficients become positive and statistically significant across maturities, and they tend to decline with maturity. Together with the positive and significant overall regressions, these findings suggest that the overall UIP deviations tend to be driven primarily during episodes of financial stress when the yield curves are in opposite directions. At the same time, the magnitude of excess returns declines with maturity, indicating that longer-horizon exchange rates predictability may partially smooth the currency risk.

These results should nevertheless be interpreted with caution. Table 3.2 shows that there are more observations in periods when both yield curves share the same sign (302 observations) than when they differ (152 observations). However, this imbalance also suggests that UIP deviations are disproportionately concentrated—and thus driven—by periods when yield-curve signs diverge.⁶

⁵I primarily focus on the signs of yield curve (i.e., upward sloping versus inverted) rather than its magnitude, because their directions already provide useful information about whether the economy is in a normal or stressed state. See [Marin and Lloyd \(2024\)](#) for interpretations based on the magnitude of the yield curve slope.

⁶Observations in which either country exhibits a flat yield curve—where the difference between 12-month and 1-month Eurocurrency/LIBOR rates is zero—are excluded. There are only 15 such observations in the entire sample.

I further refine the yield curve classification into four categories: positive-positive, positive-negative, negative-positive, and negative-negative. This decomposition allows me to examine whether U.S. financial conditions exert a disproportionate influence on global currency risk premia. This question is particularly relevant in light of the literature on the Global Financial Cycle driven by U.S. monetary policy and the role of the U.S. dollar as the dominant international reserve currency (Miranda-Agrippino and Rey, 2022, 2020; Rey, 2013)

Figure 3.5 shows that the most pronounced excess-return premia arise when the U.S. yield curve is inverted, and the U.K. yield curve signs seem to play a smaller role. By contrast, there is little evidence of UIP deviations when both yield curves are upward sloping. These findings suggest that UIP tends to hold more closely during normal periods and that U.S. financial conditions play a disproportionate role in shaping global currency risk premia.

More broadly, the results indicate that the yield curve contains economically meaningful information about the nature of UIP deviations. Episodes in which yield curves share the same sign—especially when both are upward sloping—are more closely aligned with the UIP condition. In contrast, episodes in which yield curves move in opposite directions are associated with larger excess returns and more pronounced UIP deviations. An inverted U.S. yield curve, in particular, appears to generate stronger UIP deviations regardless of U.K. conditions. These findings reinforce the regime-dependent view of the UIP puzzle: the UIP deviations are not constant over time and cannot be explained by a single static mechanism across all states of the world.

3.4 CONCLUDING REMARKS

This paper argues that UIP coefficients are not merely unstable; they are both regime-specific and maturity-dependent. Using Eurocurrency/LIBOR rates at the 1-, 3-, 6-, and 12-month maturities, I show that UIP coefficients evolve with the volatility of interest-rate differentials,

vary systematically across maturities, and respond to yield-curve regimes. The apparent sign reversal of UIP coefficient estimates after the 2008 Global Financial Crisis must therefore be interpreted with caution—the compression of short-term interest-rate differentials near the zero lower bound introduces econometric complications that can distort the interpretation of the “New Fama Puzzle”.

To address this issue, I introduce the term structure of currency deposit rates to disentangle macroeconomic delayed-overshooting mechanisms from financial risk-premium effects. The results suggest that these two explanations are not mutually exclusive. Delayed overshooting helps explain why short-horizon currency trades tend to violate the UIP condition during normal periods: short-run exchange-rate fluctuations are more volatile than long-run exchange rate movements. By contrast, when short-term interest rates are compressed toward zero by monetary authorities, the term structure considerations become more important in explaining why longer maturities become more informative and why yield curve shapes play a central role. A complete explanation of the UIP puzzle must therefore incorporate both maturity and regime dependence.

Last but not least, this paper shows that UIP deviations are particularly pronounced when U.S. and U.K. yield curves move in opposite directions, suggesting that country-specific shocks become more important when domestic economic conditions diverge. At the same time, the evidence points to a dominant role for U.S. financial conditions in global currency pricing. Episodes with an inverted U.S. yield curve are associated with especially large excess return premia, consistent with the central role of the U.S. dollar in international finance and the convenience-yield interpretation of U.S. safe assets. In short, the UIP condition cannot be understood through a single static regression. Instead, it reflects regime-dependent dynamics in which macroeconomic delayed overshooting and financial risk premia—each operating with different strengths across risk environments—jointly shape the observed UIP deviations.

Table 3.1: Summary statistics: exchange rate changes and interest rate differentials.

		1-month		12-month	
	#obs	$\Delta s_{t \rightarrow t+1}$	$i_t^{*,(1)} - i_t^{(1)}$	$\Delta s_{t \rightarrow t+12}$	$i_t^{*,(12)} - i_t^{(12)}$
	(1)	(2)	(3)	(4)	(5)
<u>(a) All observations</u>					
	469	-0.10 (2.95)	0.19 (0.23)	-1.19 (11.29)	2.01 (2.29)
<u>(b) Exchange rate regimes</u>					
Currency board	18	0.07 (1.86)	0.03 (0.17)	-0.16 (6.04)	0.58 (0.71)
Pre-ERM moving bands	219	-0.11 (3.22)	0.25 (0.28)	-1.69 (13.02)	2.53 (2.81)
ERM crawling bands	23	-0.39 (4.11)	0.51 (0.05)	-8.36 (13.22)	5.25 (0.83)
Post-ERM moving bands	100	-0.20 (2.54)	0.09 (0.10)	-0.74 (5.36)	1.00 (0.99)
Managed moving bands	96	-0.01 (2.49)	0.14 (0.10)	1.08 (11.74)	1.61 (1.14)
Freely floating	13	0.43 (3.41)	0.04 (0.02)	-1.59 (6.55)	0.13 (0.20)
<u>(c) Yield curve signs: same vs. opposite</u>					
Same	302	-0.19 (2.78)	0.15 (0.19)	-1.59 (11.10)	1.78 (2.08)
Opposite	152	0.08 (3.29)	0.25 (0.27)	-0.50 (12.02)	2.38 (2.61)

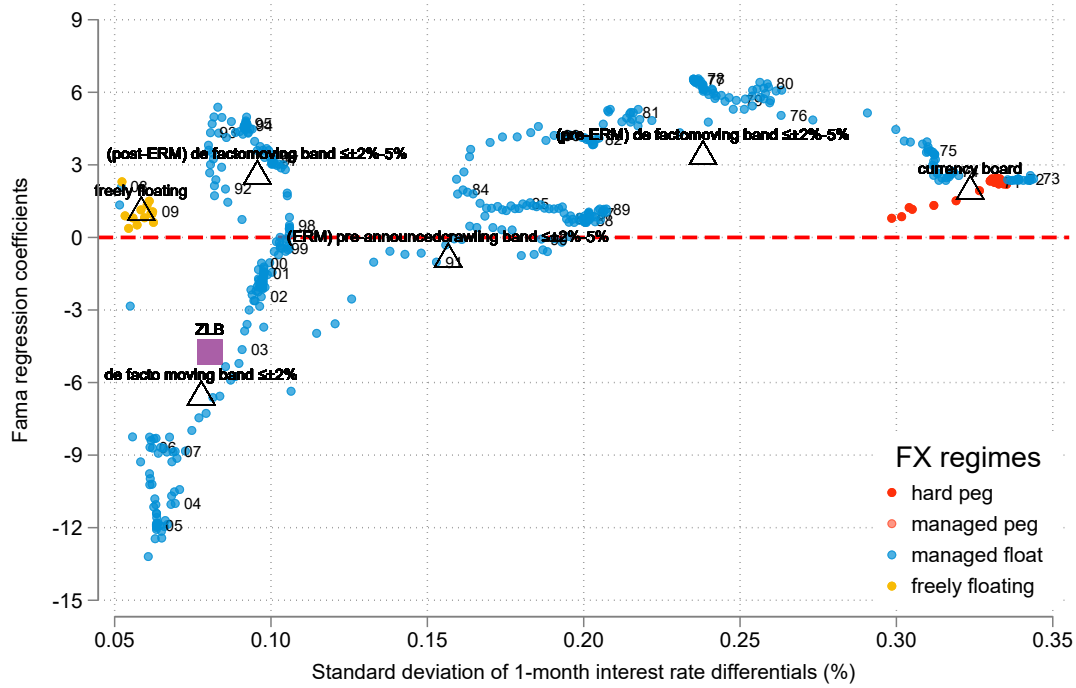
Continued on next page

Table 3.2: Summary statistics: exchange rate changes and interest rate differentials—continued

		1-month		12-month	
	#obs	$\Delta s_{t \rightarrow t+1}$	$i_t^{*,(1)} - i_t^{(1)}$	$\Delta s_{t \rightarrow t+12}$	$i_t^{*,(12)} - i_t^{(12)}$
	(1)	(2)	(3)	(4)	(5)
<i>(d) Yield curve signs: pp, pn, np, nn</i>					
pp	247	-0.22 (2.80)	0.13 (0.17)	-1.34 (10.56)	1.59 (1.98)
pn	116	0.19 (3.43)	0.32 (0.24)	0.37 (11.92)	2.71 (2.60)
np	36	-0.28 (2.81)	0.03 (0.24)	-3.30 (12.06)	1.32 (2.40)
nn	55	-0.05 (2.69)	0.23 (0.24)	-2.70 (13.30)	2.60 (2.29)

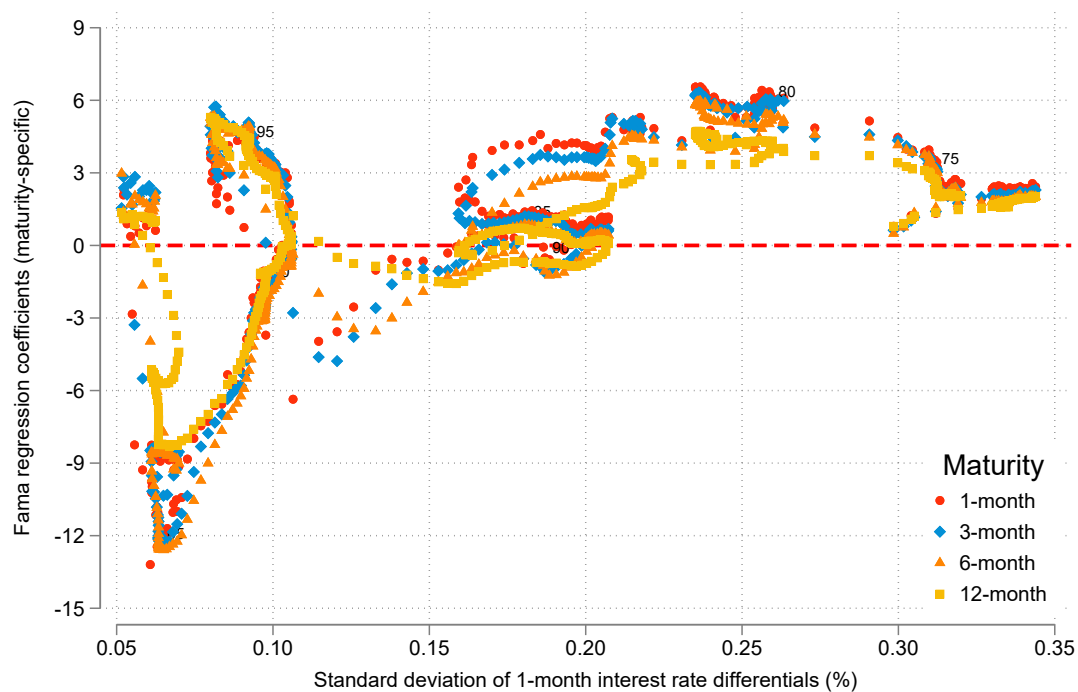
Notes: Notes: This table reports regime-specific summary statistics for exchange rate changes and interest rate differentials at the 1-month and 12-month horizons. In panel (b), exchange rate regime classifications follow [Ilzetzk, Reinhart and Rogoff \(2019\)](#). Yield-curve direction is defined by the sign of the difference between the 12-month and 1-month Eurocurrency or LIBOR rates. In panel (c), “same” indicates that the U.S. and U.K. yield curves share the same sign, while “opposite” indicates that the signs differ. In panel (d), “pp” denotes positive-positive, “pn” positive-negative, “np” negative-positive, and “nn” negative-negative yield-curve configurations. Observations with a flat yield curve are excluded; there are 15 such year-month observations out of 469 in the full sample.

Figure 3.1: Negative relationship between excess return coefficients and the volatility of interest rate differentials.



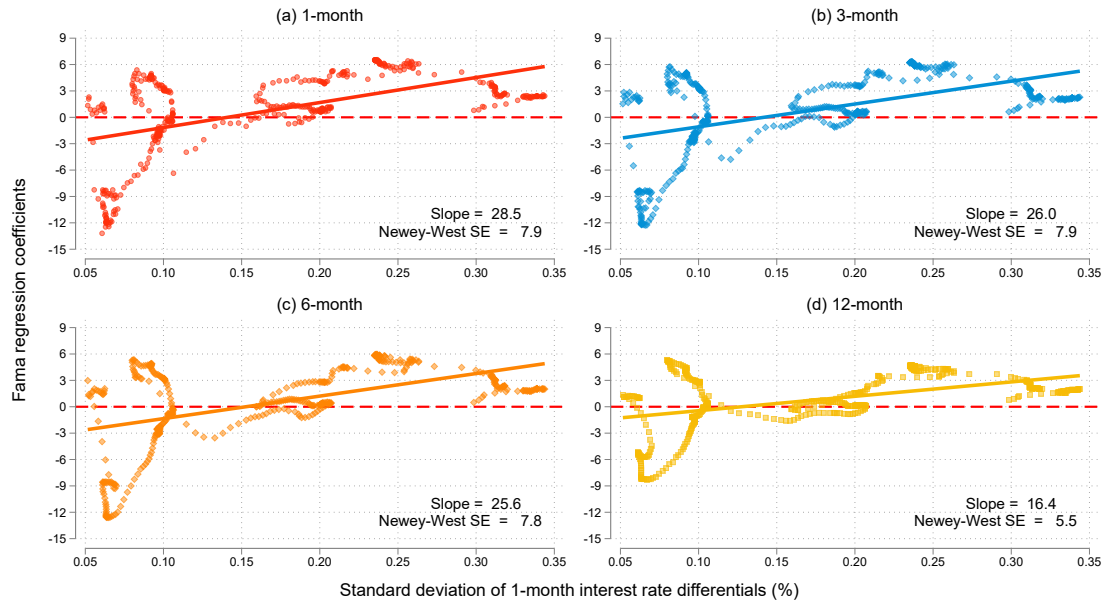
Notes: This figure shows that there is a negative relationship between the excess return coefficient estimates against the standard deviation of 1-month interest rate differentials. Each dot represents one rolling-window estimate (10 years; 120 year-month observations). Dot colors indicate exchange rate regime classifications following [Ilzetzi, Reinhart and Rogoff \(2019\)](#). Point estimates are computed using the same small-sample correction methodology as in [Engel et al. \(2022\)](#). Confidence intervals are omitted for presentation purpose.

Figure 3.2: Term structure of excess return premium before and after the 2008 Global Financial Crisis.



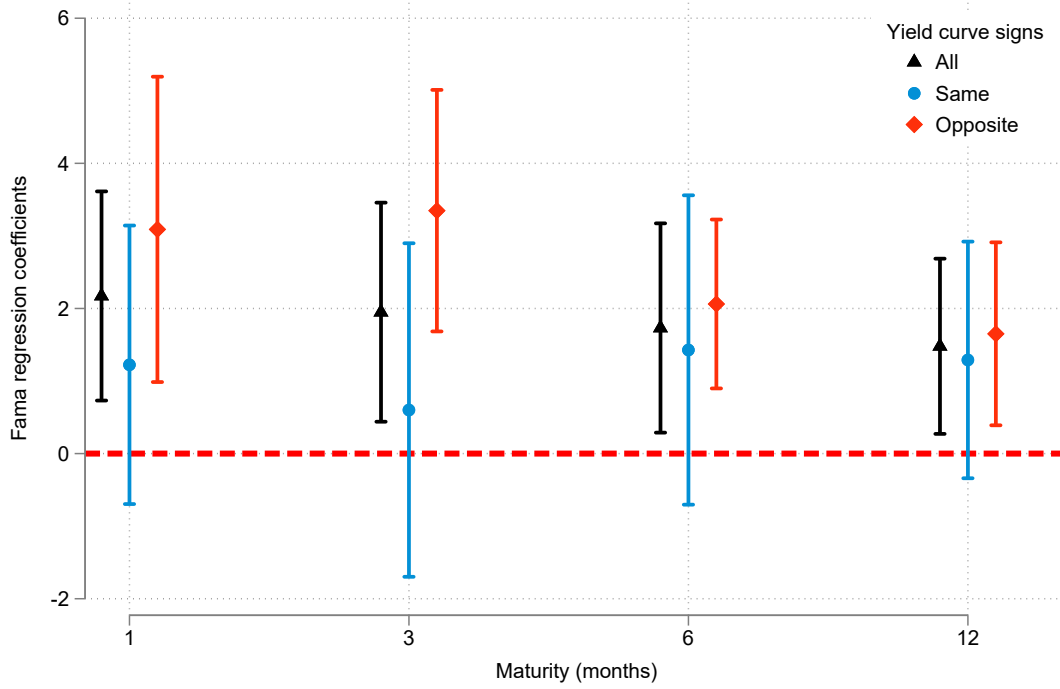
Notes: This figure plots the term structure of excess return coefficients at the 1-, 3-, 6-, and 12-month maturities against the standard deviation of 1-month interest rate differentials. Pre-1998 maturity-specific deposit rates are based on *Financial Times*. Eurocurrency rate quotations, while post-1998 observations use Bloomberg data. The regression specification follows that used in Figure 3.1.

Figure 3.3: Responsiveness of excess return coefficients to interest rate differential volatility across maturities.



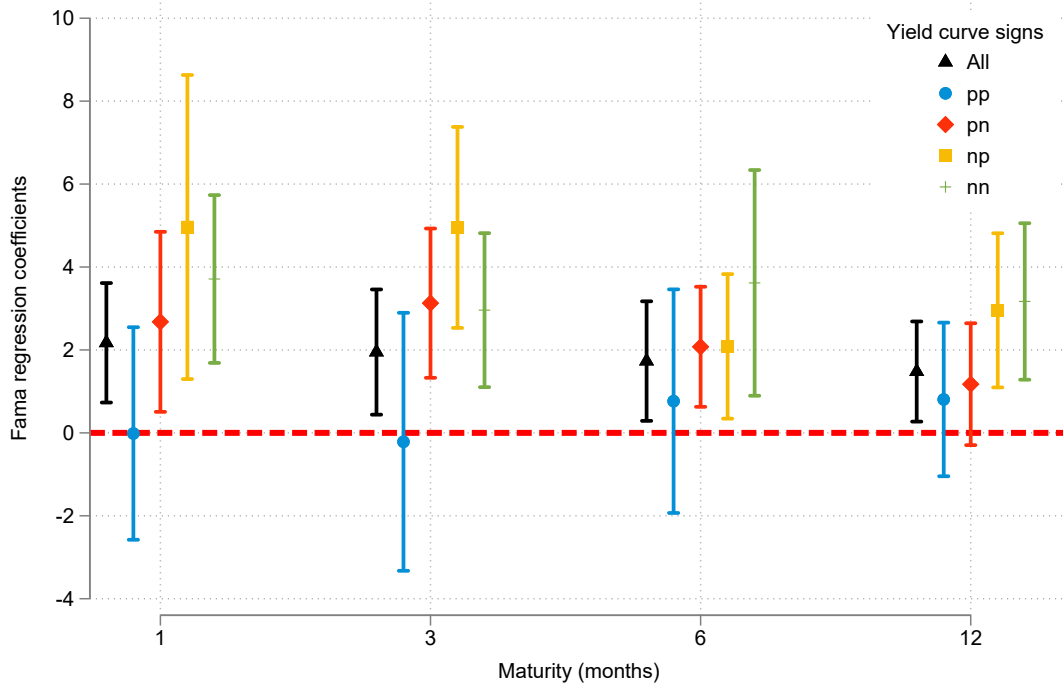
Notes: This figure illustrates the maturity-specific relationship between excess return coefficients and the standard deviation of 1-month interest-rate differentials. The estimated slopes indicate that short-maturity excess returns are more responsive to interest rate differential volatility, which is a proxy for macro-financial risk, than longer maturities. Newey-West standard errors are reported.

Figure 3.4: Excess return coefficients under same versus opposite yield curve signs.



Notes: This figure shows that the UIP deviations are primarily driven by observations in which the U.S. and U.K. yield curves have opposite signs. It reports maturity-specific excess return coefficient estimates after splitting observations according to whether the U.S. and U.K. yield curves share the same or opposite signs. The yield curve slope is defined as the difference between the 12-month and 1-month Eurocurrency/LIBOR rates, both in per annum. A yield curve is upward-sloping if its slope is positive, inverted if negative, and flat when equals to 0. Newey-West standard errors are reported.

Figure 3.5: Excess return coefficients across yield curve sign combinations.



Notes: This figure shows the UIP condition tends to hold more closely during normal state when both the U.S. and U.K. yield curves are upward sloping. It also shows inverted U.S. yield curves are associated with more positive excess return. The maturity-specific excess return coefficients are computed in four yield curve sign combinations: positive-positive, positive-negative, negative-positive, and negative-negative. This pattern is consistent with the dominant role of U.S. financial conditions in global currency risk premia. Newey–West standard errors are reported.

3.5 APPENDIX FIGURES

Figure A1: Snapshot of Eurocurrency Rates from Financial Times on July 31st, 1975

EURO-CURRENCY INTEREST RATES*						
July 31 1975	Sterling	U.S. Dollar	Canadian Dollar	Dutch Guilder	W. German mark	Swiss franc
†Short term ...	7 ³ / ₄ -8 ¹ / ₂	5 ³ / ₄ -6 ¹ / ₄	5 ¹ / ₂ -6 ¹ / ₂	1-2	2-3	2-3
7 days notice..	10 ¹ / ₄ -10 ¹ / ₂	5 ⁵ / ₈ -6 ¹ / ₈	6-7	11 ¹ / ₂ -13 ¹ / ₄	25 ⁵ / ₈ -27 ⁵ / ₈	7 ¹ / ₂ -8
Month	10 ¹ / ₂ -11	6 ¹ / ₈ -6 ⁵ / ₈	7 ¹ / ₈ -7 ¹ / ₂	23 ¹ / ₄ -31 ¹ / ₄	31 ¹ / ₂ -55 ⁵ / ₈	53 ¹ / ₄ -4
Three month..	10 ⁷ / ₈ -11 ⁵ / ₈	6 ³ / ₄ -7 ¹ / ₄	7 ¹ / ₈ -7 ¹ / ₈	31 ¹ / ₄ -51 ¹ / ₂	33 ¹ / ₄ -4	41 ¹ / ₄ -1 ¹ / ₂
Six months	11 ³ / ₄ -12 ¹ / ₄	7 ³ / ₄ -8 ¹ / ₄	8 ¹ / ₅ -8 ⁵ / ₈	51 ¹ / ₈ -55 ⁵ / ₈	45 ⁵ / ₈ -47 ⁵ / ₈	51 ¹ / ₄ -1 ¹ / ₂
One year	12 ¹ / ₂ -15	8 ¹ / ₄ -9 ³ / ₄	8 ¹ / ₈ -8 ⁷ / ₈	65 ⁵ / ₈ -67 ⁵ / ₈	53 ¹ / ₄ -6	53 ¹ / ₄ -6

Figure A2: Maturity-specific Eurocurrency/LIBOR Rates in the United States.

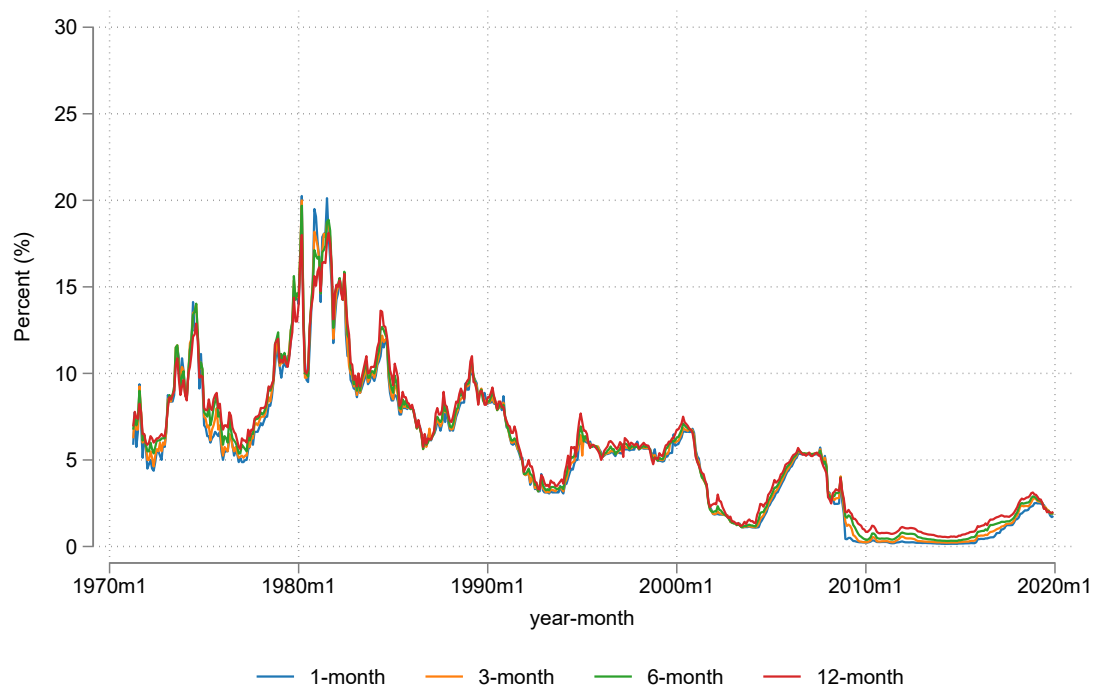


Figure A3: Maturity-specific Eurocurrency/LIBOR Rates in the United Kingdom.



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