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Essays on Exchange Rates, Monetary Policy, and Intermediation

By

JAEJUN YOON
DISSERTATION

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To my family, for their endless love and support.

Contents

Acknowledgments	xi
Abstract	xii
1 Understanding Exchange Rate Dynamics: The Fama Puzzle and Monetary Policy	1
1.1 Introduction	1
1.2 Related Literature	5
1.3 Documenting the Puzzles: New Stylized Facts	8
1.3.1 Data and Sample Selection	9
1.3.2 The Simultaneous Structural Breaks in Three Major Puzzles	10
1.3.3 The Asymmetric R-squared Puzzle	17
1.4 A Policy-Driven Explanation: Model and Estimation Results	21
1.4.1 Model Framework	21
1.4.2 Estimation Strategy: Simulated Method of Moments (SMM)	24
1.4.3 SMM Estimation Results	27
1.4.4 Investigating SMM Results: Underlying Mechanisms	29
1.4.5 Beyond the Fama Puzzle: Backus–Smith and Meese–Rogoff Puzzles	34
1.5 Empirical Tests of the Hypothesis	36
1.5.1 Augmented Fama Regression as a Robustness Check	36

1.5.2	Time-Varying Taylor Rule Estimation	39
1.6	Conclusion	41
2	State-Dependent Delayed Overshooting	42
2.1	Introduction	42
2.2	Related Literature	45
2.3	Empirical Strategy	46
2.3.1	Data Description and Macroeconomic State	46
2.3.2	State-Dependent Local Projections with KBO Decomposition	48
2.4	Main Empirical Results	50
2.4.1	Asymmetric Exchange Rate Dynamics	50
2.4.2	Asymmetric Excess Return Dynamics	50
2.5	Robustness Checks	53
2.5.1	Alternative Shock Identification: Purging Central Bank Information Effects	53
2.5.2	Alternative State Variable: National Financial Conditions Index	55
2.5.3	Post-Volcker Period Analysis	57
2.6	Monte Carlo Analysis of Truncated Sample Bias	59
2.6.1	Analytical Framework of the Truncated Instrument Bias	60
2.6.2	Quantitative Simulation Results	62
2.7	Implications	64
3	Bridging Local Housing and Global Dollar Markets: The Intermediary Channel	66
3.1	Introduction	66
3.2	Related Literature	69
3.3	Empirical Motivation	70
3.4	A Two-Period Model of Domestic Collateral and CIP Deviations	76

3.4.1	Stage 1: Intra-temporal Optimization and the Price Index	76
3.4.2	Stage 2: Inter-temporal Optimization and the Nominal Euler Equation	78
3.4.3	Macroeconomic Implication: Income-Driven Import Compression . .	79
3.4.4	Hedged Dollar Liquidity and the CIP Basis	79
3.4.5	The Financial Intermediary	80
3.4.6	Equilibrium	84
3.4.7	Regimes and Main Implications	85
3.5	A Full DSGE Model with Endogenous CIP Deviations	88
3.5.1	Households, Housing, and Domestic Credit	89
3.5.2	Consumption Aggregation and Prices	90
3.5.3	Firms and Price Setting	91
3.5.4	Covered Dollar Funding and the CIP Basis	92
3.5.5	Intermediary Balance Sheet and Net Worth	92
3.5.6	Dollar Liquidity and the VaR Constraint	94
3.5.7	Asset Pricing Conditions	96
3.5.8	Foreign Exchange Market Clearing and the Balance of Payments . . .	98
3.5.9	Housing Stress and Risk Repricing	99
3.5.10	Market Clearing and Equilibrium Definition	100
3.6	Quantitative Exercise: Dynamic Mechanism	102
3.6.1	Implementation	102
3.6.2	Calibration and Stochastic Simulation Design	103
3.6.3	Stochastic Simulation Design	104
3.6.4	Quantitative Analysis: Disconnect and Reconnect in General Equilib- rium	106
3.6.5	Sensitivity Analysis	108
3.7	Conclusion	111

A Appendix A 118

A.1	Data Description	118
A.2	Full Model	121
A.2.1	Households	121
A.2.2	Production and Pricing	122
A.2.3	Nominal Rigidities	122
A.2.4	Market Clearing and External Balance	123
A.2.5	Calibration	123
A.3	A Structural Break in Fama Puzzle	125
A.4	A Structural Break in Meese-Rogoff Puzzle	127
A.5	A Structural Break in the Backus-Smith Puzzle	131
A.6	Rolling Window Scatterplot for Individual Country	134
A.7	Rolling Window R2-beta	135
A.8	beta-R2 regression	137
A.9	Augmented Fama Regression	138
B	Appendix B	141
B.1	Full Model	141
B.1.1	Households	141
B.1.2	Production, Pricing, and Nominal Rigidities	142
B.1.3	Monetary Policy	142
B.1.4	International Asset Markets and the Endogenous UIP Wedge	142
B.1.5	Calibration	143
B.2	Diagnostic: Shock-Induced State Transitions	144
C	Appendix C	146
C.1	Endogenous macro-natural hedge	146
C.2	Full DSGE Model	147
C.2.1	Households	147

C.2.2	Consumption Aggregation and Prices	150
C.2.3	Firms and Price Setting	151
C.2.4	The Representative Financial Intermediary	151
C.3	Derivation of Optimal Portfolio Choice	155

List of Figures

1.1	Structural Breaks in Three Major Puzzles	11
1.2	The Asymmetric β - R^2	18
1.3	Binscatter Plot	20
1.4	Rolling window Fama Regression Heatmap	33
1.5	Rolling-Window Estimates of Augmented Taylor Rule Coefficients	39
2.1	State-Dependent Exchange Rate Dynamics and Friction-driven Amplification	51
2.2	State-Dependent Excess Return Dynamics and UIP Deviations	52
2.3	Robustness to Information Effects: AD (2024) NLP Shocks	54
2.4	Robustness to an Alternative State Variable: NFCI	56
2.5	Robustness to Subsample: Post-Volcker Period (RR Shocks)	58
2.6	Robustness to Subsample: Post-Volcker Period (AD Shocks)	59
2.7	Monte Carlo Simulation of Truncated Sample Bias	63
3.1	State-dependent relationship between housing price growth and CIP deviations	71
3.2	Three-Line Mechanism Overlay under an Uncertainty Shock	110
3.3	Three-Line Mechanism Overlay under a 3pp Realized NPL Shock	110
A.7.1	The Extended Asymmetric β - R^2	135
A.7.2	Shadow Rate Fama estimation	136
B.2.1	Shock-Induced State Transition Diagnostics	145

List of Tables

1.1	Summary of Stylized Facts: Pre- vs. Post-GFC Panel Moments	17
1.2	SMM Estimation Results: A Comparison of Policy Rules	27
1.3	Simulation Decomposition of Fama Coefficient and R^2	31
1.4	Augmented Panel Fama Regression Results: Pre- vs. Post-GFC	37
3.1	State-Dependent Spillover: Housing Prices and Signed CIP Deviations . . .	74
3.2	Key Calibration Values	104
3.3	Simulated Regression Diagnostics: The Disconnect and Reconnect	107
3.4	Stochastic Simulation Peak Diagnostics across Structural Shocks	109
A.1.1	Data for Fama and Meese-Rogoff Puzzles (Monthly, 1994m1–2019m12) . .	119
A.1.2	Data for Backus-Smith Puzzle (Quarterly, 1994q1–2019q4)	120
A.2.1	Model Calibration (Following IM21)	124
A.3.1	Pre-GFC: 1994m1 to 2007m12	125
A.3.2	Post-GFC: 2008m1 to 2019m12	125
A.3.3	Fama Regression Panel Estimation (1994m1-2019m12)	126
A.4.1	Out-of-Sample Forecast Performance vs. Random Walk (Clark-West Test)	127
A.4.2	Meese and Rogoff Regression : Pre-GFC (1995m1 - 2007m12)	128
A.4.3	Meese and Rogoff Regression : Post-GFC (2008m1 - 2019m12)	129
A.4.4	Structural Break Test for Predictability Model Regression (Panel FE) . . .	130
A.5.1	Backus-Smith Correlations: Pre- vs. Post-Global Financial Crisis	131

A.5.2	Panel Regression Analysis of the Backus-Smith Puzzle	133
A.9.1	Augmented Fama regression (Simulated data)	139
B.1.1	Model Calibration	144

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Abstract

This dissertation investigates how foreign exchange rate dynamics interact with macroeconomic policy regimes, market environments, and financial intermediaries. Recognizing that currency movements encapsulate complex fundamental information, this research analyzes exchange rate behavior from multi-dimensional angles.

Chapter 1 demonstrates that a structural shift in monetary policy towards financial stability following the 2008 Global Financial Crisis provides a unified explanation for the simultaneous evolution of the Fama, Meese-Rogoff, and Backus-Smith puzzles. Utilizing two-country New Keynesian model, I show that a financial-stability-augmented Taylor rule successfully replicates these post-GFC patterns. These results underscore that accounting for the systematic evolution of monetary policy rules is a prerequisite for understanding international asset prices.

Chapter 2 establishes that the delayed overshooting of exchange rates in response to monetary policy shocks is state-dependent, governed by the risk-bearing capacity of global financial intermediaries. Applying a state-dependent local projection framework and a Kitagawa-Blinder-Oaxaca decomposition, I find that a contractionary U.S. monetary policy shock generates protracted currency appreciation exclusively under a restrictive policy environment driven by a limits-to-arbitrage wedge.

Chapter 3 documents a state-dependent connection between domestic housing markets and global dollar funding markets, where disconnect breaks down during macro-financial crises. I develop a small open economy dynamic stochastic general equilibrium model featuring global intermediaries subject to an occasionally binding Value-at-Risk constraint. Non-linear stochastic simulations reveal that this stress-time reconnect is driven by forward-looking macro-financial uncertainty shocks that induce sharp risk repricing, systematic deleveraging, and an endogenous scarcity of global dollar liquidity.

Chapter 1

Understanding Exchange Rate

Dynamics: The Fama Puzzle and Monetary Policy

1.1 Introduction

This paper argues that the fundamental shift in monetary policy towards financial stability following the 2008 Global Financial Crisis (GFC) provides a unified explanation for the simultaneous evolution observed across three foundational puzzles in international macroeconomics: the Fama, the Backus-Smith, and the Meese-Rogoff puzzles. The GFC marked a turning point, augmenting the traditional focus on inflation stabilization with a mandate to maintain financial stability, often termed “Leaning Against the Wind” (LAW). I contend that this policy regime shift had profound implications, reshaping the dynamics underlying these long-standing puzzles. By documenting new empirical regularities showing the contemporaneous post-GFC changes across all three puzzles, this paper uses these anomalies as a laboratory to demonstrate how the evolution of monetary policy is the key driver altering the behavior of exchange rates and related macro-financial variables.

This paper’s investigation uses the Fama puzzle—the canonical challenge to Uncovered Interest Parity (UIP)—as a primary laboratory. The Fama regression (Fama, 1984), relating exchange rate changes to interest differentials, famously yields coefficients inconsistent with UIP. Crucially, I treat this regression not as a structural estimate (acknowledging its inherent endogeneity biases) but, following modern literature (Lilley et al., 2022), as a diagnostic tool. Its evolving coefficients and explanatory power (R^2) offer a lens into the changing relationship between interest rates and exchange rates. Therefore, the instability of the Fama regression results, particularly its post-GFC shift, is not a methodological flaw but the central empirical pattern my policy-driven structural framework aims to explain.

Recent literature has highlighted this evolving puzzle. Itskhoki and Mukhin (2021) provide a powerful explanation for the traditional “Fama puzzle”, showing how financial shocks can generate the negative coefficient observed in the pre-crisis era. More recently, however, Bussière et al. (2022) documents the emergence of a “New Fama Puzzle” in the post-GFC period, characterized by a reversal to a positive coefficient. This paper reconciles these findings, arguing that the structural shift in monetary policy governs this transition. My central contribution is demonstrating how augmenting a standard model with a policy rule reacting to financial conditions provides a unified explanation for both the old and new Fama puzzles. Furthermore, this structural shift is not an isolated phenomenon confined to the Fama puzzle. This paper documents two other foundational puzzles that also exhibit contemporaneous improvements over the exact same period. The Backus-Smith correlation (Backus and Smith, 1993) moves significantly toward its theory-consistent positive sign, indicating better international risk sharing, and the out-of-sample predictability of exchange rates improves, weakening the Meese-Rogoff puzzle (Meese and Rogoff, 1983; Engel and Wu, 2024). The simultaneity of these shifts across three distinct classic puzzles provides consistent support for a single common driver: the evolution of policy behavior.

This policy evolution hypothesis is not a mere theoretical construct but is grounded in the well-documented shift in central bank mandates and rhetoric following the GFC. The Bank

of England, for instance, was given a formal financial stability mandate under the Financial Services Act of 2012. Similarly, the European Central Bank enhanced its macroprudential role through the creation of the European Systemic Risk Board (ESRB). Even without a formal mandate change, the Federal Reserve repeatedly emphasized financial stability as a key consideration in its policy deliberations (Bernanke, 2013; Adrian and Liang, 2018). This paper argues that these acknowledged shifts in policy orientation had first-order consequences for international asset prices.

To formalize this mechanism and its impact on exchange rate dynamics, I adapt a two-country New Keynesian model with segmented financial markets, following Itskhoki and Mukhin (2021). The key innovation is to incorporate an augmented Taylor rule in which the central bank leans against the wind, reacting not only to inflation but also directly to financial conditions. The model’s core mechanism operates through the endogenous co-movement between the policy rate and financial shocks. In the pre-GFC regime when the policy follows a conventional Taylor rule, the policy rate responds to a financial shock only indirectly, through their effects on inflation. In contrast, in the post-GFC period, monetary authorities responded proactively to financial disturbances, dampening their impact on domestic variables and simultaneously altering the propagation of shocks. This regime shift provides a unified mechanism that can account for the concurrent evolution of the Fama, Backus–Smith, and Meese–Rogoff puzzles.

This paper’s key contribution is providing a unified theoretical and quantitative framework that simultaneously explains the evolution of three canonical international macroeconomic puzzles. This contribution encompasses three main elements:

First, I document the simultaneous evolution of three canonical puzzles in international macroeconomics. Using rolling-window regressions, I establish that the post-GFC period is characterized by a structural break that is not confined to the Fama puzzle. I show that the emergence of the “New Fama Puzzle” (a positive UIP slope) coincides with a signifi-

cant strengthening of the international risk-sharing condition (the Backus-Smith puzzle)—a finding that, to my knowledge, has not been previously documented— and a marked improvement in exchange rate predictability (the Meese-Rogoff puzzle). Documenting this simultaneity motivates my search for a unified explanation.

Second, I uncover a new stylized fact regarding the Fama regression: an “Asymmetric R-squared”. I document that the regression’s in-sample explanatory power (R^2) is paradoxically higher during the pre-GFC era, when the estimated slope is negative (interpreted as ‘wrong’). Conversely, the R^2 falls in the post-GFC period, even as the slope coefficient turns positive. This finding presents a paradox: standard intuition suggests that as the coefficient aligns with theory (UIP), the model’s explanatory power should increase, not collapse. The observed pattern also contrasts with traditional literature that often documents a negative coefficient associated with a near-zero R^2 (Itskhoki and Mukhin, 2021). This perplexing R^2 behavior, therefore, constitutes a key empirical challenge that my paper aims to resolve.

Third, I provide a theoretical and quantitative framework that rationalizes these new stylized facts. Theoretically, I demonstrate how the single shift to a LAW policy regime provides a common driver for all empirical findings. This policy-driven stabilization (1) realigns the interest rate differential with future exchange rate movements, thereby solving the Fama β puzzle, (2) mitigates volatility in consumption growth, thereby strengthening the Backus-Smith risk-sharing correlation, and (3) makes the policy response to financial shocks systematic, thereby allowing forecasting models to be better specified, which improves exchange rate predictability and weakens the Meese-Rogoff puzzle. The framework also rationalizes the Asymmetric Fama R-squared documented earlier. To quantitatively validate this unified mechanism, I estimate the model using the Simulated Method of Moments (SMM). The estimated model, driven only by the activation of the LAW rule, successfully replicates the simultaneous shift in all the documented empirical regularities. This demonstrates that the evolution of monetary policy is a single, sufficient driver for the complex, simultaneous shifts observed in international macro-financial data.

1.2 Related Literature

The Canonical Fama Puzzle and Its Explanations

The Uncovered Interest Parity (UIP) condition, a cornerstone of international finance, predicts that interest rate differentials should align with expected currency depreciation. However, a vast body of empirical evidence consistently rejects this hypothesis, leading to the well-known Fama puzzle (Engel, 2014; Fama, 1984). This anomaly is typically documented through the classic Fama regression, where the estimated slope coefficient relating exchange rate changes to interest differentials is often found to be negative, rather than the theoretically predicted value of unity, particularly in the pre-GFC era.

Numerous explanations have been proposed to reconcile this discrepancy. Risk-based accounts argue that the negative coefficient reflects a time-varying risk premium, where high-yield currencies compensate investors for bearing additional risk, leading to persistent excess returns (Verdelhan, 2010). Others highlight market imperfections or behavioral factors, suggesting that expectation errors, deviations from rational expectations, or infrequent portfolio adjustments by investors cause systematic deviations from UIP (Gourinchas and Tornell, 2004; Bacchetta and Wincoop, 2010a; Burnside et al., 2011; Candian and De Leo, 2023).

More recently, attention has shifted toward the role of financial frictions. Itskhoki and Mukhin (2021) provide a powerful framework arguing that financial shocks—stemming from factors like intermediary constraints and noise trader behavior—are key drivers of exchange rate dynamics and UIP deviations. Their model successfully generates a negative Fama coefficient, consistent with pre-GFC evidence, by showing how limited risk-bearing capacity in financial markets affects both interest rates and exchange rate movements. This line of research complements earlier findings documenting the breakdown of arbitrage during financial crises due to capital constraints and heightened risk aversion (Jeanne and Rose, 2002; Gabaix and Maggiori, 2015a).

The Evolving Puzzle Post-GFC

Recent research reveals that the Fama coefficient is not stable over time. Most notably, Bussière et al. (2022) document the emergence of a “New Fama puzzle,” observing that the coefficient turned persistently positive in the post-GFC period—a direct contradiction to its pre-GFC behavior. Some studies link this reversal to factors specific to the post-GFC environment, such as the constraints imposed by the zero lower bound (ZLB) and the adoption of unconventional monetary policies (Burnside, 2019; Engel, Kazakova, et al., 2022). Others note that the UIP relationship can vary across horizons (Chinn and Meredith, 2004; Chinn and Quayyum, 2012; Engel, 2016; Candian and De Leo, 2023; Lloyd and Marin, 2024).

However, explaining the sustained positive coefficient post-GFC remains difficult, particularly within frameworks like Itskhoki and Mukhin (2021) where financial shocks fundamentally drive UIP deviations towards a negative coefficient. Furthermore, existing literature has largely overlooked a key empirical finding documented in this paper: the paradoxical “Asymmetric R-squared” relationship. I show that the Fama regression’s explanatory power (R^2) was paradoxically higher during the pre-GFC period when the coefficient was ‘wrong’ (negative) and declined post-GFC even as the coefficient turned ‘right’ (positive). This asymmetry challenges standard statistical intuition and requires a deeper explanation beyond changes in shock composition or the mere presence of the ZLB.

Crucially, the instability of the Fama puzzle does not appear to be an isolated phenomenon. As this paper documents empirically, the post-GFC period also witnessed contemporaneous structural improvements in two other major international macroeconomic puzzles: the Backus-Smith puzzle (international risk-sharing) and the Meese-Rogoff puzzle (exchange rate predictability). The weakening of the Meese-Rogoff puzzle, implying improved exchange rate predictability based on fundamentals, has been noted by Engel and Wu (2024). The simultaneity of these shifts across three distinct puzzles strongly suggests a common under-

lying driver is at play.

Monetary Policy Shifts as a Unifying Explanation

This paper proposes that the structural shift in monetary policy behavior following the GFC is the key unifying mechanism capable of explaining the Fama puzzle’s evolution (including the coefficient reversal and the Asymmetric R-squared) and the concurrent changes in the Backus-Smith and Meese-Rogoff puzzles.

The idea that monetary policy can influence UIP tests is not new. Central banks set interest rates endogenously in response to macroeconomic and financial conditions, potentially biasing Fama regression estimates. As argued by McCallum (1994) and Chinn and Meredith (2004), such policy feedback mechanisms can distort UIP tests and could potentially explain the negative Fama coefficients observed historically.

What is central to my argument is the growing body of evidence identifying a distinct structural shift in monetary policy conduct after the 2008 GFC. Recent studies show that financial stability considerations—such as credit spreads, asset prices, and financial stress indicators—became significantly more prominent in central bank reaction functions alongside traditional targets like inflation and the output gap (Ajello et al., 2016; Baxa, Horvath, and Vašíček, 2013). This shift is reflected in changes in estimated Taylor Rule coefficients and the development of augmented Taylor Rule specifications that explicitly incorporate financial stability metrics (Dean and Schuh, 2021; Käfer, 2014; Laséen and Svensson, 2011). This paper provides further direct evidence for this shift through time-varying estimation of the Federal Reserve’s reaction function, showing a significant change in the response to financial conditions post-GFC.

This documented policy evolution occurred amid an ongoing debate about whether central banks should explicitly target financial stability (Bernanke and Gertler, 2001; Woodford, 2012). Regardless of the normative debate, my findings, along with the broader literature,

support the view that central banks did, implicitly or explicitly, incorporate financial conditions into their decision-making more systematically after the crisis, often described as “leaning against the wind”.

1.3 Documenting the Puzzles: New Stylized Facts

This section establishes the novel empirical findings that motivate the paper’s theoretical analysis. I proceed in two parts, corresponding to the first two contributions outlined in the introduction.

First, I document the simultaneous evolution of the three canonical puzzles (Fama, Backus-Smith, and Meese-Rogoff) around the 2008 GFC. Using rolling-window panel regressions, I demonstrate that the structural break in the Fama puzzle is not an isolated phenomenon. Rather, it coincides with a significant strengthening of the international risk-sharing condition (the Backus-Smith puzzle) and a marked improvement in exchange rate predictability (the Meese-Rogoff puzzle). Presenting these shifts as a single, contemporaneous event is a key empirical contribution, as it strongly suggests the presence of a single common driver capable of explaining all three.

Second, I uncover a new, paradoxical stylized fact regarding the Fama regression itself: the “Asymmetric R-squared”. I document that the regression’s in-sample explanatory power (R^2) is higher when the slope coefficient is negative and “wrong” (pre-GFC), and subsequently collapses as the slope turns positive and “correct” (post-GFC). This finding presents a new empirical challenge, as it contrasts sharply with the traditional literature that often associates the negative Fama coefficient with a near-zero R^2 (Chinn and Meredith, 2004; Itskhoki and Mukhin, 2021).

1.3.1 Data and Sample Selection

The empirical analysis in this paper focuses on a panel of 11 major economies: the United Kingdom, Japan, Canada, the Euro Area, Switzerland, Sweden, Norway, Australia, New Zealand, South Korea, and Mexico. The United States serves as the base country against which all bilateral exchange rates and interest rate differentials are measured. These countries are selected to align with the economies included in the Bank for International Settlements (BIS) 'narrow' effective exchange rate (EER) index. The BIS narrow basket identifies the economies whose currencies are significant in terms of global trade and financial market liquidity.

While the full sample period spans from 1994 through 2019, the core of the analysis focuses on two distinct subsamples to identify structural breaks: a pre-GFC period from 1994 to 2007 and a post-GFC period from 2008 to 2019. The rationale for this division is to isolate periods characterized by relatively stable and distinct monetary policy regimes. The pre-GFC sample begins in 1994, capturing the period after the stabilization of inflation-targeting frameworks in many central banks. The post-GFC period begins with the clear structural break induced by the 2008 crisis, capturing the subsequent shift in policy focus towards financial stability. The sample ends in 2019 to exclude the extraordinary shocks and unconventional policy responses associated with the COVID-19 pandemic, which are beyond the scope of this paper's analysis.

The dataset is constructed from several sources and frequencies. For the Fama and Meese-Rogoff regressions, I collect monthly data on bilateral spot exchange rates against the U.S. dollar and three-month interbank interest rates. This dataset is augmented with CPI-based year-over-year inflation rates, the NFCI index as a proxy for financial risk, and U.S. trade balance data for the predictability exercises. Furthermore, to test the Backus-Smith puzzle, I collect quarterly data on non-durable goods consumption for each country. For the Euro Area, these data are available starting from 1999. The real exchange rate is constructed

using the bilateral spot exchange rates and CPI data described above. Full details on the dataset, including sources and variable construction, are provided in Appendix A.1.

1.3.2 The Simultaneous Structural Breaks in Three Major Puzzles

Figure 1.1 provides the visual evidence for the paper’s first key empirical contribution, demonstrating the contemporaneous structural breaks across the three puzzles. The estimates plotted in each panel (Panel (a) for the Fama coefficient, Panel (b) for the Meese-Rogoff predictability statistic, and Panel (c) for the Backus-Smith risk-sharing coefficient) are generated from 5-year (60-month or 20-quarter) rolling-window panel regressions that include country fixed effects .

As the figure highlights, all three panels show a sharp, structural break coinciding with the GFC. Crucially, this shift is not a temporary artifact of the crisis data itself; the new coefficient levels (e.g., a positive Fama and Backus-Smith coefficient and significant CW statistic) persist long after the GFC window has passed (e.g., in windows from 2014-2019). This persistence strongly indicates a durable regime shift rather than a transient crisis effect.

Panel (a): The Fama Puzzle

Panel (a) of Figure 1.1 plots the rolling-window Fama coefficient, $\beta_{(\tau)}^f$. As is visually evident, the coefficient is persistently negative—consistent with the “Fama Puzzle”—in the pre-GFC period. However, it undergoes a sharp structural break around the 2008 GFC, rising to and persisting at positive values (Bussière et al., 2022). This coefficient, $\beta_{(\tau)}^f$, is estimated using a 3-month horizon ($k = 3$) over a 60-month rolling window ($H = 60$) from the following panel regression:

$$s_{i,t+k} - s_{i,t} = \alpha_{i,(\tau)}^f + \beta_{(\tau)}^f (i_{k,i,t} - i_{k,i,t}^*) + u_{i,t+k}, \quad t \in [\tau - H + 1, \tau]. \quad (1.1)$$

In this setup, for each rolling window of length H months ending at time τ , the re-

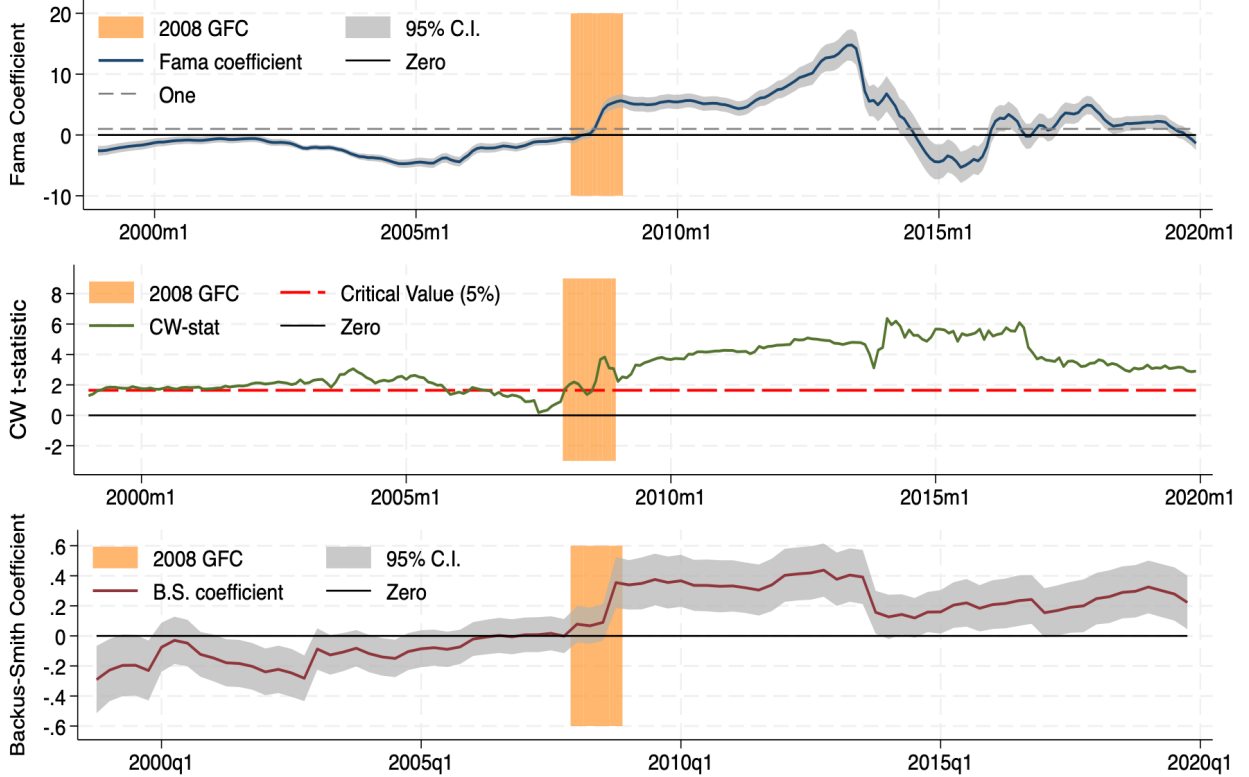


Figure 1.1: This figure documents the contemporaneous structural break in three canonical international macroeconomic puzzles. All panels plot estimates from rolling-window panel regressions for 11 economies relative to the U.S. The vertical orange band highlights the Global Financial Crisis.

Panel (a) plots the rolling Fama coefficient (β_t^f), which reverses from negative (the “Fama Puzzle”) to positive (UIP-consistent) precisely around the GFC. Panel (b) plots the rolling Clark-West (2007) t -statistic. The statistic rises from insignificant (near zero) to significantly positive (above the 5% critical value, red dashed line), indicating that exchange rate predictability improved (the “Meese-Rogoff Puzzle” weakened) after the GFC. Panel (c) plots the rolling Backus-Smith coefficient (ρ_t^{bs}), which turns from negative (the “Backus-Smith Puzzle”) to positive, suggesting an improvement in international risk sharing.

The sample period is 1994-2019. Panels (a) and (b) use monthly data with a 60-month rolling window. Panel (c) uses quarterly data with a 20-quarter rolling window. Gray shaded areas represent 95% confidence intervals. All regressions include country fixed effects.

gression pools observations across N countries indexed by i . The slope coefficient $\beta_{(\tau)}^f$ is re-estimated as the window advances through time. I set the baseline window length to $H = 60$ months, and standard errors are clustered by country within each window. The inclusion of country fixed effects allows me to control for unobserved, time-invariant country characteristics—such as average risk premia, structural financial market features, or central bank credibility—that could otherwise confound the estimated relationship between interest differentials and exchange rates.

This dynamic analysis extends the canonical Fama (1984) puzzle. The puzzle, which is traditionally tested using the following static (unconditional) Fama regression, stems from the Uncovered Interest Parity (UIP) condition. Suppose that an investor in the U.S. can invest her money into a domestic asset or a foreign asset. The return from the domestic asset investment is the interest rate with maturity k , $(1 + i_{k,t})$. Similarly, the return from investment in a foreign asset, in foreign currency, is $(1 + i_{k,t}^*)$. An investor can convert the foreign investment back into dollars using the spot exchange rate, represented as $(1 + i_{k,t}^*) \frac{S_{t+k}}{S_t}$. If we assume a world without transaction costs and financial market frictions, and with rational expectations and risk neutral investors, arbitrage would equate returns. To simplify the analysis, literature commonly use the log transformation of returns and exchange rates. Thus, the no-arbitrage condition can be expressed as the UIP condition as follows:

$$E_t [s_{t+k} - s_t] = (i_{k,t} - i_{k,t}^*) \quad (1.2)$$

where $s_{t+k} - s_t$ represents the log changes in the spot exchange rate, while $i_{k,t} - i_{k,t}^*$ denotes the interest rate differential over horizon k . If the U.S. interest rate is higher than the foreign interest rate ($i_{k,t} > i_{k,t}^*$), then the U.S. currency will depreciate, resulting in an increase in the log spot exchange rate ($s_{t+k} > s_t$). In essence, the UIP condition suggests that the interest rate differential is offset by changes in exchange rates. To empirically test this hypothesis, the Fama regression is typically estimated using ex-post realized data, which

compares actual future exchange rate changes to current interest rate differentials.

$$s_{t+k} - s_t = \alpha^f + \beta^f (i_{k,t} - i_{k,t}^*) + u_{t+k} \quad (1.3)$$

Here, the coefficient β^f measures the sensitivity of exchange rate changes to the interest rate differential. Empirical tests of the UIP condition using the Fama regression typically examine whether $\beta^f = 1$. A β^f significantly different from unity would suggest deviations from UIP, possibly due to risk premium, market imperfections, or other factors. Additionally, a non-zero intercept term, α^f , may reflect a systematic risk premium for a given country or an expectation error. Additionally, it is important to mention that the typical Fama regression is a simple reduced-form unconditional test. Rather than attempting to identify the “true” structural parameter, the goal of this exercise is to document how the coefficient from this standard, widely-cited specification has evolved over time. The subsample Fama regression results are consistent with findings in the existing literature: β^f is negative during the pre-GFC period and positive in the post-GFC period. Detailed results for individual countries, panel regressions, and formal structural break test are provided in Appendix A.3.

Panel (b): The Meese-Rogoff Puzzle

Panel (b) of Figure 1.1 plots the rolling Clark-West (2007) t -statistic. As is visually evident, the statistic persistently hovers near the critical value in the pre-GFC period, indicating that the fundamental models fail to significantly outperform a simple random walk (RW). This is consistent with the classic Meese-Rogoff puzzle, which refers to the seminal finding that fundamental-based exchange rate models consistently fail to produce more accurate OOS forecasts than a simple random walk. However, coinciding with the GFC, the t -statistic undergoes a sharp structural break, rising well above the 5% critical value (red dashed line) and persisting at a significant level. This result indicates a dramatic improvement in exchange rate predictability, suggesting a post-GFC weakening of the puzzle.

At each point τ on the graph, this statistic is computed from a 5-year rolling-window out-of-sample (OOS) forecasting exercise following Engel and Wu (2024). First, the fundamental model is estimated using a 5-year (60-month) window. This model is then used to produce a 1-month-ahead OOS forecast. The CW t -statistic plotted is then calculated from the resulting 5-year rolling window of these forecast errors, comparing them to the errors from an RW benchmark.

The fundamental forecasting model is specified as:

$$\Delta s_{i,t} = \alpha_{i,(\tau)} + \beta'_{(\tau)} \mathbf{X}_{i,t-1} + \epsilon_{i,t}, \quad t \in [\tau - H + 1, \tau]. \quad (1.4)$$

where $\Delta s_{i,t}$ is the log change in the nominal bilateral exchange rate relative to USD, $\alpha_{i,(\tau)}$ is a country-specific fixed effect, and $\mathbf{X}_{i,t-1}$ is a vector of explanatory variables from the previous period, which includes:

- Each country's real interest rate
- Each country's year-over-year inflation rate
- The lagged bilateral real exchange rate
- The US trade balance to GDP ratio
- Risk indices (e.g., NFCI)

To formally test the performance of model, I assess the predictability of exchange rates by comparing the forecast performance of the model against the random walk benchmark using the Clark and West (CW) test (Clark and West, 2006; Clark and West, 2007). The variable to be forecast is $y_{i,t} = \Delta s_{i,t}$. The null model is the random walk, for which the forecast of the exchange rate movement is zero, $\hat{y}_{i,t}^{rw} = 0$. The resulting forecast error for the random walk model is:

$$e_{i,t}^{rw} = y_{i,t} - \hat{y}_{i,t}^{rw} = \Delta s_{i,t}$$

For the alternative model specified in Equation 1.4, the OOS forecast (using parameters

estimated up to $t - 1$) is $\hat{y}_{i,t}^{alt} = \hat{\alpha}_{i,(\tau)} + \hat{\beta}'_{(\tau)} \mathbf{X}_{i,t-1}$. The resulting forecast error is:

$$e_{i,t}^{alt} = y_{i,t} - \hat{y}_{i,t}^{alt} = \Delta s_{i,t} - \hat{y}_{i,t}^{alt}$$

The CW test statistic is based on the adjusted mean squared prediction error (MSPE) difference, denoted by $d_{i,t}$:

$$d_{i,t} = \underbrace{(e_{i,t}^{rw})^2}_{\text{RW Squared Error}} - \left[\underbrace{(e_{i,t}^{alt})^2}_{\text{Model Squared Error}} - \underbrace{(\hat{y}_{i,t}^{alt} - 0)^2}_{\text{Adjustment Term}} \right]$$

A positive average value of $d_{i,t}$ indicates that the alternative model performs better than the random walk after adjusting for the noise introduced by estimating additional parameters. The CW t -statistic is calculated by regressing $d_{i,t}$ on a constant and testing if the constant is significantly greater than zero:

$$t_{CW} = \hat{\mu}_d / \widehat{se}(\hat{\mu}_d)$$

A significant positive CW statistic leads to the rejection of the null hypothesis, indicating that the structural model's forecasts are superior to the random walk. Detailed results for individual countries, panel CW estimation and formal structural break test are provided in Appendix A.4.

Panel (c): The Backus-Smith Puzzle

Finally, Panel (c) of Figure 1.1 plots the rolling Backus-Smith coefficient, $\rho_{(\tau)}$, which captures the paper's third stylized fact. Mirroring the other two panels, the coefficient is persistently close to zero or negative in the pre-GFC period, which is the essence of the "Backus-Smith Puzzle." This indicates a failure of the international risk-sharing condition. However, in perfect synchrony with the Fama and Meese-Rogoff breaks, the coefficient undergoes a structural shift during the GFC, rising to and persisting at positive values. This result can be

interpreted as an improvement in international risk sharing in the post-GFC period.

This coefficient $\rho_{(\tau)}$ is estimated from the following 5-year (20-quarter) rolling-window panel regression:

$$\Delta \log q_{i,t} = \alpha_{i,(\tau)} + \rho_{(\tau)} \Delta \log(c_{i,t}/c_{i,t}^*) + \epsilon_{i,t}, \quad t \in [\tau - H + 1, \tau]. \quad (1.5)$$

where $\Delta \log q_{i,t}$ is the quarterly change in the log real exchange rate and $\Delta \log(c_{i,t}/c_{i,t}^*)$ is the quarterly growth of relative non-durable consumption against the U.S.

This dynamic analysis extends the canonical Backus and Smith (1993). The puzzle stems from the theory of complete international risk sharing, which posits that the real exchange rate should equal the ratio of marginal utilities of consumption. A key implication of this theory is that relative consumption growth should be positively correlated with changes in the real exchange rate. Intuitively, this means a country experiencing a positive shock that boosts its consumption relative to its trading partners should see its currency depreciate in real terms as part of the risk-sharing mechanism. However, a large body of empirical work finds that this correlation is often near zero or even negative. Further analysis regarding this traditional correlation and formal structural break test are detailed in Appendix A.5.

Summary of Stylized Fact: The Simultaneous Breaks

To quantify the visual regime shift identified in Figure 1.1 and to summarize the paper's first set of stylized facts, Table 1.1 presents the key panel regression moments estimated separately on the two subsamples: Pre-GFC (1994-2007) and Post-GFC (2008-2019). These moments will serve as the target moments in the structural estimation section that follows.

The table provides robust, statistical confirmation of the visual evidence from the rolling-window analysis. First, the Fama coefficient (β^f) flips from significantly negative to significantly positive (-0.904 to 1.521). Second, the Backus-Smith coefficient (ρ^{bs}) mirrors this, turning from statistically insignificant (0.101) to strongly positive (0.692). Finally,

the Meese-Rogoff puzzle weakens, as the out-of-sample predictability measured by the CW t -statistic becomes highly significant (rising from 2.068 to 3.811).

Table 1.1: Summary of Stylized Facts: Pre- vs. Post-GFC Panel Moments

Puzzle / Moment	Specification	Pre-GFC (1994-2007)	Post-GFC (2008-2019)
1. Fama Puzzle			
Fama Coefficient (β^f)	Static Panel FE (Eq. 1.3)	-0.904*** (0.178)	1.521*** (0.452)
2. Meese-Rogoff Puzzle			
CW statistics (vs. RW)	Static Panel FE (Eq. 1.4)	2.068** (n/a)	3.811*** (n/a)
3. Backus-Smith Puzzle			
B.S. Coefficient (ρ^{bs})	Static Panel FE (Eq. 1.5)	0.101 (0.086)	0.692*** (0.104)

Notes: This table summarizes the key empirical moments from panel regressions for each puzzle. Panel 1 reports β^f from the static Fama regression (Eq. 1.3). Panel 2 reports the CW statistics (vs. a random walk). Panel 3 reports ρ^{bs} from the static Backus-Smith regression (Eq. 1.5). All models include country fixed effects. Standard errors are in parentheses. *** $p < 0.01$, ** $p < 0.05$.

1.3.3 The Asymmetric R-squared Puzzle

In Figure 1.2, I plot the Fama coefficients ($\beta_{(\tau)}^f$) and the corresponding explanatory power (within- $R_{(\tau)}^2$) from each regression estimated from the Equation (1.1), following a similar approach to Lilley et al. (2022), who highlight a reconnection between exchange rates and risk measures through the time path of $R_{(\tau)}^2$. The navy line, representing $\beta_{(\tau)}^f$ based on the preceding 60 months of data, was persistently negative throughout the pre-GFC period and turned generally positive during the post-GFC period. The red line, which traces $R_{(\tau)}^2$, also exhibited considerable time variation, with several notable peaks.¹

Interestingly, $R_{(\tau)}^2$ peaked above 0.15 in the mid-2000s, particularly when $\beta_{(\tau)}^f$ was negative. These elevated values suggest a relatively high degree of (in-sample) predictability in

¹Appendix Figure A.7.1 replicates this analysis using an extended sample (1975–2024). The same qualitative asymmetry between β^f and R^2 persists, indicating that the pattern is not specific to the pre-GFC episode.

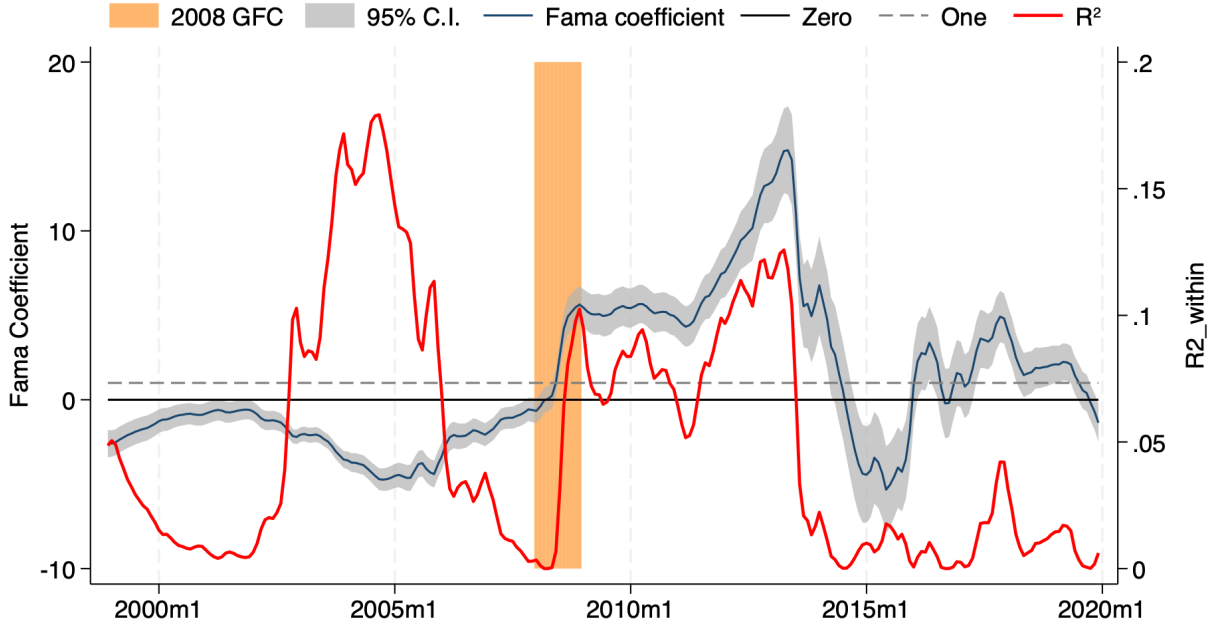


Figure 1.2: The Asymmetric β - R^2 Relationship: This figure plots the 60-month rolling-window Fama coefficient ($\beta_{(\tau)}^f$, navy line, left axis) and the corresponding within- R^2 (red line, right axis) from the panel regression (Eq. 1.1).

exchange rate changes based on interest rate differentials, even during periods typically characterized as inconsistent with the UIP condition. In contrast, during the post-GFC period, the $R_{(\tau)}^2$ overall level remained significantly lower than in earlier episodes, despite similarly large absolute values of $\beta_{(\tau)}^f$ estimates. This finding becomes even clearer after 2013, once the acute crisis window has passed. This disconnect implies that a sign alignment alone does not constitute evidence in favor of UIP and therefore does not resolve the Fama puzzle; there is an asymmetric relationship between the coefficients and the R^2 .

This pattern is particularly interesting because traditional Fama puzzle literature consistently reports that Fama regressions with short horizons (e.g. 1 to 3 months) have negative coefficients with low predictability (Chinn and Meredith, 2004; Valchev, 2020). From this prior, Itskhoki and Mukhin (2021) argues that these arbitrarily low R^2 values, potentially caused from the exogenous shocks, can explain the observed pattern. However, my finding challenges this view, demonstrating that the R^2 was not arbitrarily low during the pre-GFC

period. Instead, the regression’s explanatory power was paradoxically highest when the coefficient was most ‘wrong’.

Furthermore, this finding presents a theoretical puzzle because it directly contradicts the standard statistical relationship between a regression coefficient and its R^2 . Under the standard OLS assumptions, particularly the exogeneity of the regressor, the OLS estimator ($\hat{\beta}^f$) is unbiased, and the population R^2 derived from the Fama regression specification is a function of the square of the true coefficient, $(\beta^f)^2$. The derivation is straightforward based on the definition of R^2 as the ratio of explained variance to total variance:

$$\begin{aligned} R^2 &= \frac{\text{Var}(\mathbb{E}_t[\Delta s_{t+1}])}{\text{Var}(\Delta s_{t+1})} = \frac{\text{Var}(\hat{\beta}(i_t - i_t^*))}{\text{Var}(\Delta s_{t+1})} \\ &= \frac{\hat{\beta}^2 \cdot \text{Var}(i_t - i_t^*)}{\text{Var}(\Delta s_{t+1})} \end{aligned} \quad (1.6)$$

From equation (1.6), it is clear that, under the standard OLS assumptions (particularly exogeneity), the population R^2 is an increasing function of $(\beta)^2$. Theoretically, therefore, the magnitude of the estimated time-varying $R^2_{(\tau)}$ should increase symmetrically with the absolute value of the rolling coefficient, $|\beta^f_{(\tau)}|$, regardless of its sign (conditional on the variance of the interest rate differential).²

This theoretical relationship, however, contrasts with my empirical findings, which show a clear asymmetry. This discrepancy is the source of the “Asymmetric R-squared” phenomenon. It strongly suggests that the standard OLS assumptions are violated, implying that the sign of $\beta^f_{(\tau)}$ is intertwined with factors that alter the regression’s fit—such as changes in underlying structure. Thus, this asymmetric pattern reconfirms that the sign reversal of β^f is not an isolated phenomenon but is deeply connected to these underlying structural shifts. To further illustrate this asymmetry, Figure 1.3 presents binscatter plots of rolling

²The results of a regression of $R^2_{(\tau)}$ on $(\beta^f_{(\tau)})^2$ and the variance of the interest rate differential confirm a significantly flatter slope in the post-GFC period, even after controlling for the variance term. See Appendix A.8.

Fama coefficients and their associated $R^2_{(\tau)}$ values, grouped by period.

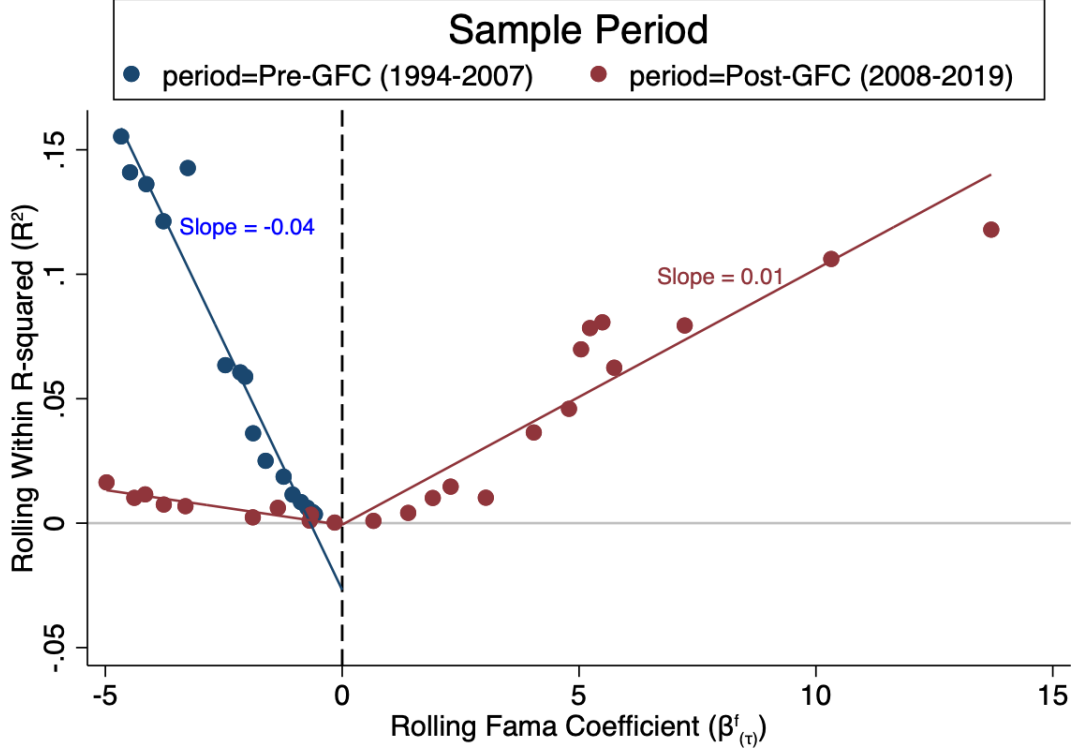


Figure 1.3: The Asymmetric β - R^2 Relationship (Binscatter) : This figure shows a binscatter plot of the rolling-window $\beta^f_{(\tau)}$ (x-axis) and within- $R^2_{(\tau)}$ (y-axis) from Figure 1.2. The plot is grouped by the Pre-GFC period (red) and Post-GFC period (blue). Slopes from linear regressions for each period are reported on the graph.

In the pre-GFC period (red), $\beta^f_{(\tau)}$ values lie in the negative domain, and their associated $R^2_{(\tau)}$ values form a steep, downward-sloping pattern (Slope = -0.04). This suggests that as the coefficient became more negative (i.e., $|\beta|$ increased), the R^2 increased—consistent with the β^2 theory, but only in the “wrong” direction. By contrast, in the post-GFC period (Blue), most Fama coefficients become positive, yet the $\beta^f_{(\tau)}$ - $R^2_{(\tau)}$ relationship is noticeably flatter, with a slope of only +0.01, indicating weaker predictive power despite the coefficient aligning with UIP in sign.

Taken together, these asymmetric patterns in the Fama regression—depending on the sign of the coefficients—suggest that the observed sign reversal does not represent a structural

resolution of the Fama puzzle. Rather, it implies a shift in underlying structural conditions that govern the estimated relationship between interest rate differentials and exchange rate movements.

1.4 A Policy-Driven Explanation: Model and Estimation Results

Building on the empirical findings in Section 1.3, which documented (i) the simultaneous post-GFC evolution of the 3 fundamental puzzles (Fama, Backus-Smith, and Meese-Rogoff), and (ii) the asymmetric behavior of the Fama regression’s explanatory power. This section develops a unified quantitative framework that can reproduce these regularities through a single mechanism — the regime shift in monetary policy toward financial-stability concerns.

To this end, I employ the two-country New Keynesian model, incorporating distinct monetary policy regimes. The analysis proceeds as follows: First, I briefly outline the model framework, highlighting the key financial friction and the alternative policy rules. Second, I detail the estimation strategy using the Simulated Method of Moments (SMM), specifying the target moments derived from the empirical findings. Finally, anticipating the main result, the SMM estimation strongly confirms the hypothesized policy shift: a conventional Taylor rule fails to match the post-GFC data, while an augmented rule incorporating a direct response to financial conditions successfully replicates the key empirical patterns across both the pre- and post-GFC periods.

1.4.1 Model Framework

I adapt on the two-country New Keynesian (NK) general equilibrium model of Itskhoki and Mukhin (2021), augmenting it with a modified monetary policy rule that allows the central bank to react to financial conditions. The model features two symmetric economies (Home

and Foreign), each with its own currency and monetary authority. The key departure from a standard open-economy NK model is the inclusion of segmented financial markets. This friction introduces financial shocks (ψ_t) that act as a time-varying wedge in the Uncovered Interest Parity (UIP) condition and are central to explaining the major exchange rate puzzles.

Beyond this core friction, the framework incorporates several standard features common in the literature: representative households maximizing discounted expected utility over consumption and labor subject to a budget constraint; aggregate consumption follows Kimball (1995) preferences, allowing for variable markups and nests the standard CES aggregator, incorporating home bias in consumption bundles; nominal rigidities via Calvo price stickiness and local currency pricing (LCP) for internationally traded goods. A detailed description of the full model setup is provided in Appendix B.1.

As emphasized, the core structure and calibration parameters largely follow Itskhoki and Mukhin (2021), and the key and only modification in this paper lies in the specification of the monetary policy rule. Specifically, I allow the central bank’s reaction function to differ across the pre- and post-GFC regimes to test the central hypothesis that a shift in policy focus towards financial stability explains the observed empirical patterns.

Segmented Financial Market

The central mechanism generating UIP deviations originates from the segmented financial market. In this framework (Itskhoki and Mukhin, 2021, Sec II.B), households trade only local-currency bonds, while international positions are intermediated by risk-averse arbitrageurs who absorb demand imbalances from households and noise traders. Crucially, noise traders introduce exogenous excess demand shocks (ψ_t) for foreign currency bonds. The arbitrageurs require a risk premium to accommodate these shocks, leading to deviations from UIP. Following Itskhoki and Mukhin (2021), I capture this friction in reduced form by introducing the stochastic shock, ψ_t , directly into the log-linearized UIP condition as a

wedge:

$$E_t[s_{t+1} - s_t] = i_t - i_t^* - \psi_t \quad (1.7)$$

where $i_t - i_t^*$ is the nominal interest rate differential. The financial shock (ψ_t) is assumed to follow a persistent AR(1) process:

$$\psi_t = \rho_\psi \psi_{t-1} + \epsilon_t^\psi, \quad \epsilon_t^\psi \sim iid(0, \sigma_\psi^2) \quad (1.8)$$

This UIP wedge ψ_t is the key transmission channel for financial shocks in the model. The full modified UIP condition derived by Itskhoki and Mukhin (2021, Lemma 1) also includes a term related to households' net foreign asset position, reflecting endogenous risk premia dynamics tied to international capital flows. For analytical clarity and to isolate the role of exogenous financial shocks and policy responses, I temporarily abstract from this bond position wedge in the subsequent theoretical discussion. The full quantitative model incorporating this channel is retained in the quantitative analysis.

Monetary Policy

The model includes monetary policy shocks, defined as deviations from the systematic Taylor rule. For analytical clarity, I limit the stochastic disturbances to these two shocks: monetary and financial shocks. In the baseline model, monetary policy follows a standard Taylor rule as follows :

$$i_t = \rho i_{t-1} + (1 - \rho)\phi_\pi \pi_t + \varepsilon_t^{mp}, \quad \varepsilon_t^{mp} \sim iid(0, \sigma_{mp}^2) \quad (1.9)$$

where π_t denotes inflation, and ε_t^{mp} is a monetary policy shock. The rule is log-linearized around a steady state, so π_t represents deviations from target inflation. The output gap is omitted here to isolate the core channel of endogeneity.

In an extended version of the model, the central bank also responds directly to financial shocks, reflecting the possibility of monetary policy “leaning against the wind” in the

presence of financial market disruptions. To formalize this behavioral change, I employ an augmented Taylor rule in which the central bank also responds to a financial stability variable. The modified Taylor rule introduces an additional policy target:

$$i_t = \rho_m i_{t-1} + (1 - \rho_m)(\phi_\pi \pi_t + \phi_X X_t) + \varepsilon_t^{mp}, \quad \varepsilon_t^{mp} \sim iid(0, \sigma_{mp}^2) \quad (1.10)$$

where X_t denotes the financial stability target. In the model, I use the change in financial shock, $\Delta\psi_t$, as the additional target variable. This term captures financial market pressure, including contemporaneous innovation (ε_t^ψ) and the persistent effects from past financial shocks from the $AR(1)$ structure of ψ_t . Intuitively, financial shocks negatively affect the economy by depressing investment and consumption while raising risk premia. Anticipating deterioration in financial markets, the monetary authority lowers the policy rate in response to the shock—particularly when traditional targets such as inflation or the output gap have not yet reflected the shock.

1.4.2 Estimation Strategy: Simulated Method of Moments (SMM)

I employ the Simulated Method of Moments (SMM) to quantitatively assess the model's ability to replicate the empirical stylized facts and to estimate the key parameters that govern the proposed mechanism. The core idea of SMM is to find the set of structural parameters (θ) that minimizes the distance between a vector of moments calculated from the empirical data (m_{data}) and the same vector of moments generated by simulating the model ($m_{sim}(\theta)$) (McFadden, 1989; Pakes and Pollard, 1989). The SMM estimator minimizes a quadratic objective function:

$$\hat{\theta}_{SMM} = \arg \min_{\theta} J(\theta) = [m_{data} - m_{sim}(\theta)]' W [m_{data} - m_{sim}(\theta)] \quad (1.11)$$

where W is a weighting matrix, which determines the relative importance assigned to matching each moment condition

Target Moments

The target moments primarily focus on moments from the Fama regression for several key reasons. First, and most importantly, the Fama regression provides a parsimonious, reduced-form diagnostic that directly reflects the joint behavior of interest rates and exchange rates—the very objects most sensitive to the hypothesized shift in monetary policy behavior. Second, these Fama moments (β^f , R^2) represent the canonical empirical benchmarks for the UIP literature. Successfully replicating their observed patterns is crucial for any model seeking to provide a credible explanation for these long-standing anomalies. Finally, by focusing the estimation on Fama-related moments, the moments associated with the Backus-Smith and Meese-Rogoff puzzles remain untargeted, providing a more stringent validation test of the proposed unified explanation. Specifically, the target moments vector m_{data} includes:

1. **Fama Regression Coefficient (β^f):** The central moment capturing the Fama puzzle and its reversal. I target the estimated coefficients from the static panel regressions (Table 1.1) for both the Pre-GFC and Post-GFC periods.
2. **Fama Regression R^2 :** To capture the overall low explanatory power associated with the Fama regression, I target the R^2 values from the same static panel regressions used for β^f .³
3. **Slope of the β^2 - R^2 :** This moment directly targets the “Asymmetric R-squared” phenomenon. I use the slopes estimated both the Pre-GFC and Post-GFC periods, as shown in Appendix A.8

³It is important to distinguish this target—the low overall R^2 from the long subsample regressions—from the time-varying and sometimes elevated $R^2_{(\tau)}$ observed in the rolling-window analysis (Figure 1.2). Low R^2 values in long-sample estimations are a robust feature in the literature (Itskhoki and Mukhin, 2021). Incorporating this moment ensures the model is consistent with this standard benchmark.

4. **Relative Volatility** ($std(i_t - i_t^*)/std(\Delta s_{t+1})$): Matching this ratio ensures the model generates realistic quantitative magnitudes alongside the targeted conditional moments from the Fama regression.

To assess the model’s broader implications, I evaluate its ability to match key moments that were not included in the estimation procedure. The first untargeted moment is the Backus-Smith coefficient, ρ^{bs} , which tests the model’s performance on the international risk-sharing puzzle by examining the co-movement of relative consumption and the real exchange rate. The second is the Clark and West (CW) out-of-sample statistic, which measures predictive power against a random walk and directly evaluates the model’s ability to address the Meese-Rogoff puzzle.

Estimated Parameters and Procedure

The estimation focuses on a small set of key parameters central to the paper’s hypothesis, while other parameters are calibrated following Itskhoki and Mukhin (2021). The vector of estimated parameters, θ , includes:

- Shock Volatility (σ_{mp}/σ_ψ): The standard deviation of monetary policy shocks relative to financial shocks.
- Policy Response to Financial Conditions (ϕ_ψ): The coefficient capturing the central bank’s reaction to financial shocks ($\Delta\psi_t$) in the augmented Taylor rule (Eq. 1.10). This parameter is fixed at zero under the conventional rule.

These parameters feature in the financial shock process (Eq. 1.8) and the Taylor rules (Eq. 1.9 and 1.10). Details on the SMM implementation follow. For the weighting matrix W , I use a diagonal matrix where each diagonal element is the inverse of the corresponding squared empirical moment ($W_{ii} = 1/(m_{data,i})^2$). This choice normalizes the moment errors, preventing the estimation from being dominated by moments with larger numerical values. For a given parameter vector θ , the simulated moments $m_{sim}(\theta)$ are generated using a Monte

Carlo procedure. I simulate the model for 150 periods, repeating this 500 times with different random shock sequences. The final simulated moment is taken as the median across these simulations to ensure robustness.⁴

1.4.3 SMM Estimation Results

I estimate two competing model specifications: (1) a restricted model assuming the conventional Taylor rule ($\phi_\psi = 0$) holds in both periods, testing if shock volatility changes alone can explain the data, and (2) the full model allowing ϕ_ψ to differ across regimes via the augmented Taylor rule.

Table 1.2: SMM Estimation Results: A Comparison of Policy Rules

	Data		Conventional Rule		Augmented Rule	
	Pre-GFC (1995-2007)	Post-GFC (2008-2019)	Pre-GFC (Model 1)	Post-GFC (Model 1)	Pre-GFC (Model 2)	Post-GFC (Model 2)
Estimated Structural Parameters (θ)						
σ_{mp}/σ_ψ (Rel. Shock Vol.)	-	-	0.483	0.372	0.518	0.096
ϕ_ψ (Policy Response)	-	-	(Fixed at 0)	(Fixed at 0)	7.056	-11.552
Targeted Moments (Model Fit)						
Fama Beta (β^f)	-0.904	1.521	-1.236	-1.832	-0.959	1.577
Fama R^2	0.009	0.005	0.009	0.014	0.013	0.005
$\beta^2 - R^2$ Slope	0.590	0.066	0.284	0.204	0.552	0.096
Rel. Vol ($i^d/\Delta s$)	0.101	0.054	0.069	0.063	0.107	0.039
Untargeted Moments (Model Validation)						
Backus-Smith Coeff. (ρ_{bs}) ^a	0.101	0.692	-0.296	-1.240	-0.600	7.459
OOS CW-Stat	2.068	3.811	4.908	5.006	4.801	5.254

Notes: The table compares empirical moments with simulated moments (median across 500 simulations) from two model specifications, estimated via SMM. Model 1 uses the conventional Taylor rule ($\phi_\psi = 0$). Model 2 uses the augmented Taylor rule where ϕ_ψ is estimated. Data moments for Fama Beta, Fama R^2 , and $\beta^2 - R^2$ Slope are from Section 1.3.

^a Backus-Smith panel regression coefficient ($\Delta q_t \sim \Delta(c_t - c_t^*)$) for the respective periods.

The SMM estimation results, presented in Table 1.2, provide strong support for the hypothesized monetary policy shift. Model 1, which assumes a conventional Taylor rule ($\phi_\psi = 0$), successfully replicates the negative Fama beta of the pre-GFC period (Model: -1.236 vs. Data: -0.904). However, it critically fails to capture the post-GFC dynamics. It incorrectly predicts a large negative beta (-1.832) when the data shows a large positive

⁴The numerical minimization is performed using MATLAB's fmincon optimizer with a multi-start procedure to enhance the likelihood of finding the global minimum.

beta (1.521). Furthermore, this model struggles to match other post-GFC moments, over-predicting the Fama R^2 and the $\beta^f - R^2$ Slope. This confirms that changes in shock volatilities alone are insufficient to explain the structural break.

In sharp contrast, Model 2, which incorporates the augmented Taylor rule, performs substantially better. The estimated parameters confirm the hypothesized policy shift: the response to financial shocks (ϕ_ψ) is large and positive pre-GFC (7.056) but switches sign to become large and negative post-GFC (-11.552). This policy shift allows the model to succeed precisely where the conventional rule failed. It replicates the sign reversal, generating a positive post-GFC coefficient (1.577) that is close to the empirical value (1.521). Moreover, the augmented model provides an excellent fit for other targeted moments. It captures the steep pre-GFC $\beta^f - R^2$ Slope (0.552 vs. 0.590) and the flat post-GFC slope (0.096 vs. 0.066), matching the "Asymmetric R^2 " pattern. It also provides a tight fit for the Fama R^2 (Pre : 0.013 vs. 0.009; Post: 0.005 vs. 0.005) and the relative volatility of interest rates (Pre: 0.107 vs. 0.101; Post : 0.039 vs. 0.054).

For the Backus-Smith puzzle, the augmented model correctly captures the direction of the shift, generating a negative coefficient pre-GFC (-0.600) and a positive coefficient post-GFC (7.459), qualitatively matching the empirical improvement in risk-sharing (0.101 to 0.692). However, while the model identifies the correct direction, it significantly overshoots the magnitude of the post-GFC coefficient. Regarding the Meese-Rogoff puzzle, the data shows an improvement in out-of-sample predictability (CW-Stat rising from 2.068 to 3.811). The full model is qualitatively consistent with this, showing a slight improvement (4.801 to 5.254).

Furthermore, the parameter estimates themselves provide an insight consistent with the Zero Lower Bound (ZLB) phenomenon. In the augmented model (Model 2), the estimated relative volatility of monetary policy shocks (σ_{mp}/σ_ψ) drops from 0.518 pre-GFC to 0.096 post-GFC. This implies that after the GFC, the variance of monetary policy shocks was less

than 10% of the variance of financial shocks. This finding suggests that the post-GFC era was dominated by large financial fluctuations while conventional monetary policy shocks became negligible. At the ZLB, central banks indeed have limited scope for traditional rate shocks, and economic dynamics are instead driven primarily by the financial shocks themselves and the systematic component of policy (our estimated ϕ_ψ response), precisely as the model estimates.

1.4.4 Investigating SMM Results: Underlying Mechanisms

This section delves into the economic mechanisms underlying the SMM estimation results presented in Table 1.2. The goal is to connect the quantitative findings to the theoretical framework, demonstrating how the estimated shift in the monetary policy regime provides a unified explanation for the observed changes in the Fama, Backus-Smith, and Meese-Rogoff puzzles after the Global Financial Crisis (GFC). The analysis confirms that the augmented Taylor rule (Model 2), by capturing the significant change in the central bank's response to financial shocks (ϕ_ψ), successfully replicates the key empirical patterns where the conventional rule (Model 1) fails.

The Fama Puzzle: Coefficient Reversal and Asymmetric R-squared

The SMM results show Model 2's remarkable success in matching the Fama puzzle dynamics, particularly the sign reversal of β^f and the asymmetric β^f - R^2 relationship. This success stems directly from the estimated policy shift altering the underlying structure of the Fama regression. To unpack this mechanism quantitatively, I decompose both the Fama coefficient and the regression R^2 into components reflecting the role of the policy-induced covariance term.

Decomposition of Fama coefficients

As outlined in Equation 1.10, endogeneity arises because the financial shock (ψ_t) influences both the regressor (i_t^d) and the error term (u_t) in the Fama regression (Equation 1.3). The

monetary policy response shapes the covariance between them, $\text{Cov}(i_t^d, u_t)$. The sign and magnitude of this covariance crucially determine the bias in the OLS estimate $\hat{\beta}^f$ relative to the true β as follows :

$$\hat{\beta} = \beta_{True} + \frac{\text{Cov}(i_t^d, u_t)}{\text{Var}(i_t^d)} \quad (1.12)$$

While β_{True} is unobservable, the direction of the bias can be inferred. When $\text{Cov}(i_t^d, u_t) < 0$, the OLS estimator is biased downward relative to its true value. This situation arises, for example, when a positive financial shock (which is part of the error term) leads to a policy rate hike (increasing i_t^d), thereby inducing a systematic negative covariance.

Decomposition of Fama R^2

The distorted Fama coefficient also affects the measured $\hat{R}^2$ relative to the theoretical R^2 under exogeneity (Equation 1.6). If the exogeneity condition does not hold, as is the case with a biased Fama coefficient, the estimated $\hat{R}^2$ takes the following form:

$$\begin{aligned} \hat{R}^2 &= \frac{\text{Var}(\mathbb{E}_t[\Delta s_{t+1}])}{\text{Var}(\Delta s_{t+1})} = \frac{\hat{\beta}^2 \cdot \text{Var}(i_t^d)}{\text{Var}(\Delta s_{t+1})} \\ &= \left(\beta_{True} + \frac{\text{Cov}(i_t^d, u_t)}{\text{Var}(i_t^d)} \right)^2 \frac{\text{Var}(i_t^d)}{\text{Var}(\Delta s_{t+1})} \\ &= \frac{\left[\beta_{True}^2 \text{Var}(i_t^d) + 2\beta_{True} \text{Cov}(i_t^d, u_t) + \frac{\text{Cov}(i_t^d, u_t)^2}{\text{Var}(i_t^d)} \right]}{\text{Var}(\Delta s_{t+1})} \\ &= \frac{\left[\beta_{True}^2 \text{Var}(i_t^d) + 2 \beta_{True} \text{Cov}(i_t^d, u_t) + \frac{\text{Cov}(i_t^d, u_t)^2}{\text{Var}(i_t^d)} \right]}{\left[\beta_{True}^2 \text{Var}(i_t^d) + 2 \beta_{True} \text{Cov}(i_t^d, u_t) + \text{Var}(u_t) \right]} \end{aligned} \quad (1.13)$$

Compared to Equation 1.6, the covariance terms create bias in the estimated $\hat{R}^2$. Therefore, the $\text{Cov}(i_t^d, u_t)$ is crucial for determining the direction of bias in both the Fama coefficient and its explanatory power. Additionally, the policy response could affect the variance of the residual ($\text{Var}(u_t)$), so the effect of the policy regime shift is ambiguous from the above equations. I therefore examine this relationship through my simulation model, where the model's setting is under control.

Table 1.3: Simulation Decomposition of Fama Coefficient and R^2 using SMM Parameters
(Median across 1000 Trials per Regime)

Component	Pre-GFC	Post-GFC
Panel 1: Simulation Setting		
Relative Shock Volatility (σ_{mp}/σ_ψ)	0.5178	0.0955
Policy Response to the Financial shock (ϕ_ψ)	7.0564	-11.5520
Panel 2: Fama Coefficient (Beta) Decomposition		
Baseline (β_{true}) (A)	1.0000	1.0000
Omitted Var Bias [$\text{Cov}(i^d, u)/\text{Var}(i^d)$] (B)	-2.0245	0.4026
Estimated OLS Beta	-0.8972	1.4175
Reconstructed Beta (A + B)	-1.0245	1.4026
Panel 3: Fama Regression R-Squared Decomposition		
A. Numerator		
Actual Explained Var ($\text{Var}(\Delta \hat{s}_{t+1})$)	2.5930	2.1061
<i>Decomposition</i>		
$\text{Var}(i^d) \cdot (\beta_{\text{true}})^2$	2.8689	0.6865
+ $2\beta_{\text{true}} \cdot \text{Cov}(i^d, u)$	-11.6161	0.5528
+ $\text{Cov}(i^d, u)^2 / \text{Var}(i^d)$	11.6295	1.6156
= Reconstructed Numerator (C)	2.8822	2.8549
B. Denominator		
Actual Var ($\text{Var}(\Delta s_{t+1})$)	252.4958	447.1769
<i>Decomposition</i>		
$\text{Var}(i^d \cdot \beta_{\text{true}})$	2.8689	0.6865
+ $2 \text{Cov}(i^d \cdot \beta_{\text{true}}, u)$	-11.6161	0.5528
+ $\text{Var}(u)$	261.7291	446.1438
= Reconstructed Denominator (D)	252.9825	447.3830
Estimated R^2 from Regression ($\text{Var}(\Delta \hat{s}_{t+1}) / \text{Var}(\Delta s_{t+1})$)	0.0101	0.0047
Reconstructed R^2 (C / D)	0.0103	0.0047

Table 1.3 presents the simulation results decomposing the Fama coefficient (β^f) and R^2 using the SMM-estimated parameters for each regime.

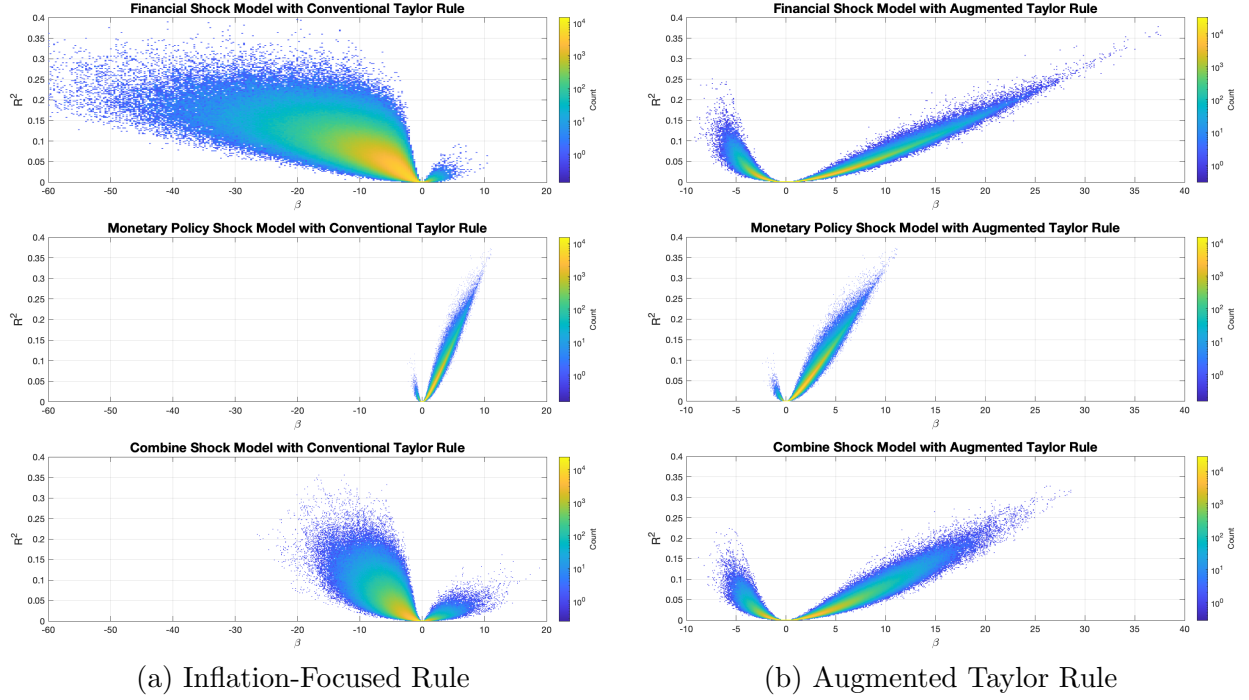
The change in the Fama coefficient is clearly explained. As shown in Panel 2, the bias in β^f is determined by the omitted variable bias term, $\text{Cov}(i^d, u)/\text{Var}(i^d)$. In the Pre-GFC regime, under the estimated policy response ($\phi_\psi = 7.0564$), this bias term has a large negative value (-2.0245). This bias overwhelms the true baseline β_{True} (1.0000), resulting in an OLS Beta of -0.8972. In the Post-GFC regime, the central bank's active response ($\phi_\psi = -11.5520$) reverses the sign of the covariance, making the bias positive (+0.4026). Consequently, the OLS Beta flips to 1.4175, successfully replicating the sign reversal puzzle.

The decline in R^2 is explained by a more complex, two-part mechanism. The simulation confirms the empirical fact: the ' R^2 ' drops from 0.0101 (1.01%) to 0.0047 (0.47%). First, the numerator in Panel 3-A slightly decreases from 2.5930 to 2.1061. Although the policy shift reverses the sign of the $2\beta_{true} \text{Cov}(i^d, u)$ term (from -11.6161 to +0.5528), which adds explanatory power, this effect is offset by a larger reduction in the variance from other components. Second, and more decisively, the denominator ($\text{Var}(\Delta s_{t+1})$) explodes by 77%, rising from 252.4958 to 447.1769. Panel 3-B shows this is almost entirely due to the variance of the residual ($\text{Var}(u)$) surging from 261.7291 to 446.1438. This illustrates that while the policy response successfully counteracts the bias on β^f , it does not eliminate the volatility from the financial shocks. This un-offset volatility is captured by the residual term, increasing the system's overall noise and driving down the R^2 .

Rolling-Window $\beta_{(\tau)}^f$ and $R_{(\tau)}^2$

I also conduct a rolling-window regression analysis using simulated data. The objective of this exercise is to replicate the empirical asymmetry observed earlier: periods with negative Fama coefficients (β^f) tend to be associated with elevated R^2 , whereas periods with positive coefficients often coincide with low R^2 . To isolate the pure effect of the policy regime change, I deviate from the SMM-estimated parameters for this exercise. Instead, I use

Figure 1.4: Rolling window Fama Regression Heatmap



a fixed calibration to demonstrate that a change in the policy rule alone is sufficient to generate the puzzle.

I generate 300 months of data, assuming a constant underlying shock process: the relative volatility ($\sigma_{mp}/\sigma_{\psi}$) is set to 0.3, following Itskhoki and Mukhin (2021). I then simulate the economy under two distinct policy rules shown in Figure 1.4: (a) the Pre-GFC "Inflation-Focused Rule" ($\phi_{\psi} = 0$) and (b) an "Augmented Taylor Rule." ($\phi_{\psi} = -5$) For each regime, I run 10,000 simulations, applying a 60-month rolling-window regression to each path to track the estimated (β^f, R^2) pairs.

The results in Figure 1.4 successfully replicate the empirical asymmetry. The heatmap for the "Inflation-Focused Rule" (a) shows that the simulated (β^f, R^2) pairs cluster densely in the region of a negative β^f and are associated with a steep slope for R^2 . In sharp contrast, the "Augmented Taylor Rule" (b) causes the distribution to shift dramatically. The pairs now cluster in the region of a positive β^f and a much flatter slope for R^2 .

This exercise successfully isolates the impact of the policy regime. It demonstrates that even with an identical underlying shock process, the change in the policy rule alone is the key mechanism responsible for generating the observed paradoxical relationship between the Fama coefficient and the slope of R^2 .

1.4.5 Beyond the Fama Puzzle: Backus–Smith and Meese–Rogoff Puzzles

The preceding analysis highlights how shifts in monetary policy behavior can systematically affect on the Fama regression. In this section, I show that the same mechanism extends naturally to the Backus-Smith and Meese-Rogoff puzzles, both of which exhibit structural improvements over the same period. These patterns reinforce the view that a regime change in monetary policy, rather than an alteration in the nature of shocks, is the unifying driver behind the evolving exchange rate dynamics.

The Backus-Smith Puzzle and Risk Sharing

In the framework of Itskhoki and Mukhin (2021), the negative relationship arises when exchange rate movements are primarily driven by financial shocks. A positive financial shock—interpreted as an excess demand for foreign-currency assets—causes an immediate depreciation of the domestic currency. Under an inflation-targeting policy rule, the central bank reacts by tightening interest rates to contain inflationary pressure, which depresses domestic consumption and induces a negative comovement between consumption and the real exchange rate.

After the Global Financial Crisis (GFC), however, the direction of the monetary response changed fundamentally. Central banks began to incorporate financial stability objectives into their policy reaction functions. When financial stress emerges, policy rates are lowered rather than raised. In this environment, a financial shock that depreciates the exchange rate

simultaneously raises consumption, producing the positive correlation. Hence, the Backus–Smith puzzle is not resolved by the disappearance of financial shocks but by the reversal of the monetary reaction to them. The same financial disturbances that once reduced consumption and risk sharing now elicit countercyclical policy easing, bringing the data closer to the theoretical benchmark of complete markets.

The Meese-Rogoff (Predictability) Puzzle

This improvement in predictability is also explained by the shift to a systematic, augmented policy rule. The mechanism is not that fundamentals suddenly ”mattered more,” but rather that the forecasting model became better specified. In the Pre-GFC period, exchange rate movements were heavily driven by unobserved financial shocks (ψ_t). Because the central bank’s response to these shocks was not systematic and indirect, the fundamental variables in the forecasting model (e.g., interest rate differentials, inflation) did not contain clear information about this dominant shock. The model was therefore misspecified, as it omitted a key driver, and its forecasts failed.

However, in the Post-GFC period, The augmented Taylor rule makes the central bank’s response to financial shocks systematic and predictable. When a financial shock occurs, it now triggers a reliable and observable policy response in the interest rate. Consequently, the observed fundamentals now serve as a valid signal for the underlying, unobserved financial shock. By including these fundamentals, the same forecasting model becomes better specified in the new regime. It implicitly captures the impact of the financial shocks via the systematic policy response.

A Unified Interpretation

Taken together, the Fama, Backus-Smith, and Meese-Rogoff puzzles reveal a coherent pattern. All three puzzles show significant improvement or reversal after the Global Financial Crisis, and all can be explainable by a single structural transformation-the shift of monetary

policy from an inflation-centric to a financial-stability-augmented regime.

Under this unified view, financial shocks remain the fundamental source of exchange rate volatility, as in Itskhoki and Mukhin (2021), but their macroeconomic propagation crucially depends on the policy reaction function. A pre-GFC regime that tightened policy in response to financial shocks amplified UIP deviations and weakened risk sharing, while a post-GFC regime that stabilized financial conditions reduced those distortions and reconnected exchange rates with fundamentals. This regime-dependent interaction between financial shocks and monetary policy thus provides a single, structural explanation for the evolving nature of the three canonical international macroeconomic puzzles.

1.5 Empirical Tests of the Hypothesis

1.5.1 Augmented Fama Regression as a Robustness Check

As established in the theoretical discussion, the standard Fama regression suffers from an endogeneity bias arising from the central bank’s systematic policy responses. To provide further corroborating evidence for my central hypothesis, this section examines whether the predictions of my structural model hold in a reduced-form empirical setting. Specifically, I augment the baseline Fama regression with proxies for the underlying drivers of policy, and test whether the key coefficients shift across regimes as predicted by the model.

I estimate the following fixed-effect panel Fama regression for the pre-GFC (1994–2007) and post-GFC (2008–2019) subsamples:

$$s_{i,t+3} - s_{i,t} = \alpha_i + \beta i_{i,t}^d + \gamma \pi_t^{\text{US}} + \theta \pi_{i,t}^{\text{Foreign}} + \phi [\log(VIX_{t+3}) - \log(VIX_t)] + u_{i,t} \quad (1.14)$$

where π denotes the annual inflation rate, and $\log(VIX)$ serves as a proxy for financial shocks. The VIX is widely regarded as a measure of global risk appetite and market sentiment. Incorporating it into exchange rate regressions follows the empirical strategies of

studies such as Kalemli-Özcan and Varela (2021) and Bussière et al. (2022). This specification closely mirrors the simulation setup in equation (A.9.1), enabling a direct comparison between simulated and empirical results with respect to coefficient patterns.

Table 1.4: Augmented Panel Fama Regression Results: Pre- vs. Post-GFC

Variable	Pre-GFC (1994m1-2007m12)				Post-GFC (2008m1-2019m12)			
	Baseline (1)	+ Infl. (2)	+ VIX (3)	Full (4)	Baseline (5)	+ Infl. (6)	+ VIX (7)	Full (8)
Dependent Variable: $\Delta s_{i,t+3}$								
Interest Rate Diff.	-0.904*** (0.178)	-1.583*** (0.189)	-0.903*** (0.178)	-1.582*** (0.189)	1.521*** (0.452)	3.763*** (0.651)	1.595*** (0.420)	2.822*** (0.610)
π^{US}		-0.0102 (0.0147)		-0.0101 (0.0147)		-0.0807*** (0.0207)		-0.0534** (0.0194)
π^{Foreign}		0.185*** (0.0330)		0.185*** (0.0330)		0.0450 (0.0364)		0.0507 (0.0339)
$\log(VIX_{t+3}) - \log(VIX_t)$			0.0303 (0.516)	0.0342 (0.503)			6.111*** (0.382)	5.970*** (0.386)
Constant	0.113 (0.127)	-15.83*** (1.634)	0.114 (0.128)	-15.83*** (1.634)	-0.332 (0.173)	14.00*** (3.297)	-0.281 (0.161)	6.666* (3.108)
N	1752	1752	1752	1752	1584	1584	1584	1584
R^2	0.015	0.066	0.015	0.066	0.007	0.021	0.146	0.151
Adj. R^2	0.00841	0.05935	0.00784	0.05881	0.00019	0.01315	0.13980	0.14332

Notes: The table presents results from augmented fixed-effect panel Fama regressions. For each period, 'Baseline' refers to the simple Fama regression; '+ Infl.' adds inflation controls; '+ VIX' adds the change in the VIX index; and 'Full' includes all variables. Standard errors clustered by country are in parentheses. * $p < 0.05$, ** $p < 0.01$, and *** $p < 0.001$ denote significance.

The empirical results in Table 1.4 provide strong corroborating evidence for the paper's central mechanism. Two key findings stand out. First, the baseline Fama coefficient on the interest rate differential flips from negative in the pre-GFC period (e.g., -0.904 in Column 1) to positive and highly significant in the post-GFC period (e.g., 1.521 in Column 5), mirroring the main stylized fact.

Second, and more importantly, the coefficient on the financial shock proxy, the change in the VIX, exhibits a dramatic structural break in its magnitude and statistical significance. In the pre-GFC period (Column 4), the coefficient is small and statistically insignificant (0.0342). This implies that financial stress, as measured by the VIX, had no clear, sys-

tematic relationship with exchange rate movements. In sharp contrast, in the post-GFC period (Column 8), the coefficient becomes large, positive, and highly significant (5.970). This empirical shift—from a statistically insignificant relationship to a strong, positive, and significant one—provides direct, reduced-form evidence that the transmission of financial shocks to the exchange rate fundamentally changed across the two policy regimes, strongly supporting this paper’s proposed mechanism. I provide results of regression from the simulated model in Appendix A.9, which are consistent with my empirical test.

1.5.2 Time-Varying Taylor Rule Estimation

The central mechanism of this paper rests on the hypothesis that the conduct of monetary policy fundamentally changed after the Global Financial Crisis, with a new emphasis placed on financial stability. To substantiate this claim with direct empirical evidence, this section estimates an augmented Taylor rule using a 40-quarter rolling-window regression. This approach allows me to track the evolution of the policy coefficients over a long sample period. While several financial stress indicators are available, I use the National Financial Conditions Index (NFCI) from the Chicago Fed as the primary proxy for financial shocks, given its comprehensive nature and longer time-series availability compared to other measures like the VIX. The findings, presented in Figure 1.5, offer compelling visual evidence of the policy shift.

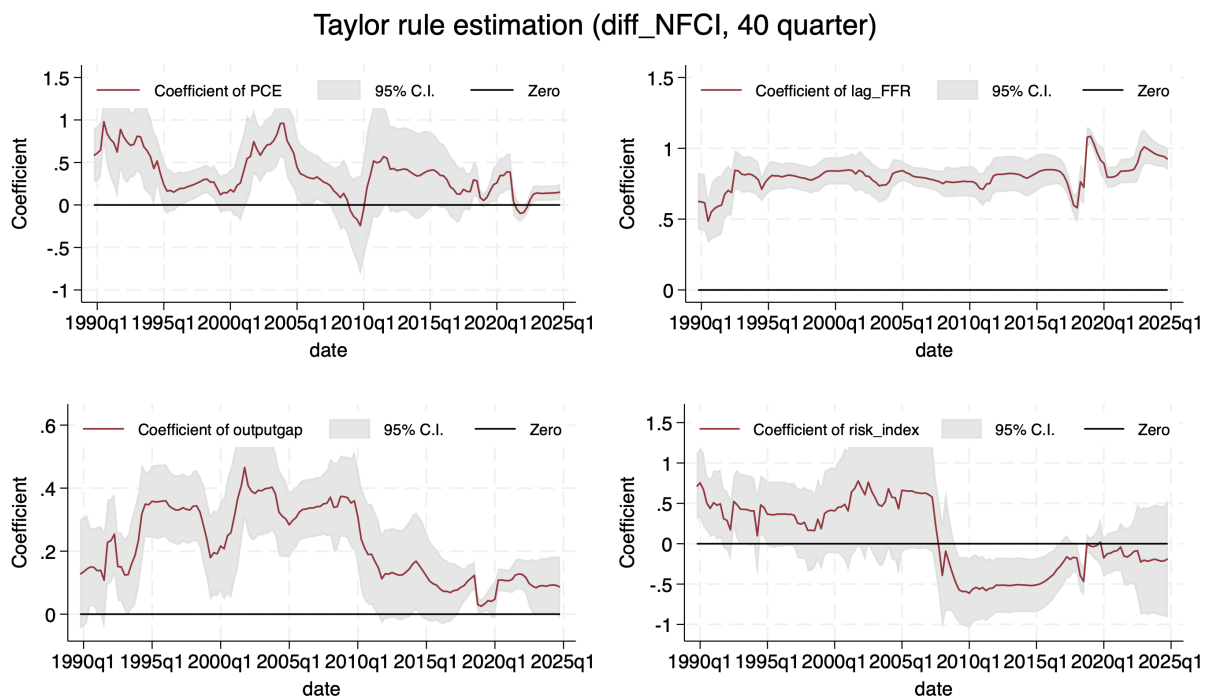


Figure 1.5: Rolling-Window Estimates of Augmented Taylor Rule Coefficients (NFCI augmented)

To conduct this analysis, I estimate a standard smoothed Taylor rule, augmented with the first difference of the NFCI as the proxy for financial market stress. The specification

for the rolling-window regression is as follows:

$$FFR_t = c_t + \beta_{p,t}FFR_{t-1} + \beta_{\pi,t}\pi_t + \beta_{y,t}\text{OutputGap}_t + \beta_{\psi,t}\Delta\text{NFCI}_t + \epsilon_t \quad (1.15)$$

The analysis uses U.S. quarterly data, with the rolling window capturing the evolving policy response from 1990 to 2024. The policy rate (FFR_t) is the effective Federal Funds Rate, inflation (π_t) is the year-over-year change in the Core PCE price index, and the output gap is from the Congressional Budget Office (CBO).

The results of this estimation, plotted in Figure 1.5, provide decisive evidence of a policy regime shift. The bottom-right panel, which displays the rolling coefficient on the financial risk index ($\beta_{\psi,t}$), is particularly revealing. Prior to the GFC, the coefficient is consistently positive, but often statistically insignificant. This indicates a hawkish pre-crisis response: a tightening of financial conditions was met with a higher policy rate, likely because the underlying drivers of financial stress were perceived as inflationary risks.

After the Global Financial Crisis, however, the coefficient undergoes a sharp and dramatic structural break. It plummets from positive territory to become negative throughout the post-GFC era. This reversal signals a fundamental shift in policy. In the post-GFC world, a tightening of financial conditions is no longer viewed through an inflation lens but is instead treated as a direct threat to economic stability. The Federal Reserve now responds to a financial instability by systematically and aggressively easing policy to support the financial system. This finding provides strong, direct empirical support for the paper’s central mechanism: the adoption of a financial stability-oriented policy response in the post-GFC era.

1.6 Conclusion

The findings of this paper provide a stark, contemporary example of the Lucas (1976) critique within international macroeconomics. The instability of reduced-form relationships like the Fama regression serves as a textbook illustration: its key coefficient underwent a structural break precisely because the underlying monetary policy regime changed towards incorporating financial stability. An economist relying solely on pre-GFC data would derive policy advice directly contradicting that derived from post-GFC data, highlighting the unreliability of reduced-form estimates when policy rules evolve. It underscores the necessity of structural models which, by separating deep behavioral parameters from the policy rule, can anticipate how empirical regularities shift in response to changing central bank objectives.

This policy-centric perspective offers a unifying lens for seemingly disparate phenomena. The documented post-GFC improvements in exchange rate predictability, attributed by Engel and Wu (2024) to enhanced central bank credibility, and the shifting patterns of UIP deviations linked to expectation errors (Bussière et al., 2022), can be understood as distinct facets of the single, structural policy shift identified here. A central bank prioritizing financial stability is naturally incentivized towards clearer communication (enhancing credibility), while the adoption of a new, complex policy rule necessitates a learning period where expectational errors can arise.

These findings ultimately underscore the necessity of accounting for policy endogeneity in modern macroeconomics. Standard empirical tests are fundamentally misspecified if they ignore the systematic evolution of policy rules. Future research could further explore causal identification and dynamic transmission channels under different regimes. In an era of active and evolving central banking, understanding policy is indeed a prerequisite for understanding the economy itself.

Chapter 2

State-Dependent Delayed Overshooting

2.1 Introduction

The canonical overshooting model (Dornbusch, 1976) predicts an immediate, sharp appreciation of the domestic currency following a monetary tightening, followed by a gradual depreciation to satisfy uncovered interest rate parity (UIP). However, classic empirical studies document a notable departure from this theoretical benchmark (Eichenbaum and Evans, 1995). Instead of an immediate jump, currencies frequently exhibit “delayed overshooting”: a domestic monetary contraction leads to a protracted appreciation that peaks months, or even years, later before the currency begins to depreciate.

I show that the monetary policy stance shapes this sluggish adjustment, generating state-dependent currency dynamics. Using a state-dependent local projection (SDLP) framework, I find that a monetary policy shock generates delayed overshooting exclusively under a contractionary policy environment. Under tight monetary conditions, the domestic currency exhibits a muted initial response but gradually appreciates, reaching its peak appreciation

approximately 15 months after the shock. In contrast, under an expansionary stance, the exchange rate adjusts rapidly, peaking within the first 5 months and closely aligning with the immediate jump predicted by standard theory.

I relate this state-dependency to the risk-bearing capacity of international financial intermediaries. The aggregate U.S. monetary policy stance drives global liquidity, altering the funding conditions of these intermediaries—a channel consistent with the Global Financial Cycle (Miranda-Agrippino and Rey, 2020). During contractionary periods, binding balance sheet constraints restrict major foreign exchange (FX) market participants, such as global dealer banks and hedge funds. These limits to arbitrage impede their ability to absorb policy shocks, yielding the delayed overshooting documented in the data. Conversely, during expansionary periods, intermediaries operate with ample liquidity, facilitating an adjustment that closely aligns with the theoretical benchmark.

Following Grimm et al. (2023), I measure the monetary policy stance using the deviation of the real policy rate from the natural rate of interest ($r_t - r_t^*$) (Holston, Laubach, and Williams, 2017). I construct this state variable as a backward-looking, 5-year moving average of the interest rate gap. High-frequency variations in policy rates reflect transient business cycles, but the risk-bearing capacity and balance sheet constraints of global intermediaries evolve slowly. A prolonged, cumulative exposure to a policy regime alters structural financial frictions and limits to arbitrage. The backward-looking moving average filters out high-frequency noise and yields a predetermined state variable isolated from contemporaneous policy shocks.

For exogenous monetary policy shocks, I use an updated version of the narrative series developed by C. Romer and D. Romer (2004) and Wieland and Yang (2020). Unlike high-frequency identification (HFI) data restricted to the post-1990 period, this extended narrative series encompasses both historically tight financial regimes and recent low-friction environments. To ensure the delayed dynamics are not an artifact of the central bank infor-

mation effect, I validate the baseline findings using narrative shocks constructed by Aruoba and Drechsel (2024). These shocks use natural language processing (NLP) to filter out the Fed’s private economic assessments, mitigating confounding informational noise.

I estimate the asymmetric exchange rate dynamics using a local projection (LP) framework (Jordà, 2005). Building on recent advances in State-Dependent LP (Cloyne, Jordà, and Taylor, 2023; Alessandri, Jordà, and Venditti, 2025), I apply a Kitagawa-Blinder-Oaxaca decomposition to the state-dependent impulse responses. This decomposition disentangles the unconditional macroeconomic response from the friction-driven amplification, quantifying the fraction of delayed overshooting directly attributable to the policy environment.

This framework reconciles conflicting empirical findings. Early studies documented the puzzle as a generalized phenomenon (Eichenbaum and Evans, 1995; Grilli and Roubini, 1996). Kim, Moon, and Velasco (2017) observed that delayed overshooting was concentrated during Paul Volcker’s tenure (1979–1987), concluding it was a historical anomaly. My state-dependent framework provides a structural rationale: the Volcker era was a highly contractionary regime characterized by severe intermediary constraints, which naturally induced the protracted currency adjustments identified by my model.

This perspective also explains why recent linear SVAR-IV studies using HFI data find no delayed overshooting (Rüth, 2020). HFI instruments are widely available only from the 1990s onward—a period dominated by the Great Moderation and its associated low financial frictions. Estimating a linear, time-invariant model with this truncated instrument set omits the dynamics of high-friction eras, biasing the average response toward low-friction adjustments.

To evaluate this sample truncation bias, I conduct a Monte Carlo simulation based on a state-dependent Data Generating Process (DGP) inspired by the segmented financial market framework of Itskhoki and Mukhin (2021). The economic intuition is straightforward: a contractionary shock impairs the risk-bearing capacity of global intermediaries, raising the

shadow cost of capital for arbitrageurs and widening the UIP wedge. Using this simulation, I show that using truncated instruments in a linear SVAR yields a regime-specific bias that masks the underlying delay. Delayed overshooting remains a robust feature of the data, obscured only by linear econometric models estimated over structurally benign historical samples.

2.2 Related Literature

This paper bridges three strands of literature in international macroeconomics: empirical exchange rate puzzles, limits to arbitrage, and state-dependent identification.

Delayed Overshooting and UIP Puzzles. The canonical benchmark for exchange rate dynamics remains the overshooting model of Dornbusch (1976), which predicts an immediate currency jump followed by a depreciation path consistent with Uncovered Interest Rate Parity (UIP). Eichenbaum and Evans (1995) famously documented the empirical failure of this prediction, showing that contractionary monetary shocks lead to a protracted, hump-shaped appreciation. This delayed adjustment violates UIP, generating predictable excess returns. Engel (2016) characterizes this anomaly as a short-run versus long-run excess return puzzle: high-interest-rate currencies feature short-run excess returns but long-run predictability reversals. I provide a structural rationale for this anomaly, showing it is not an unconditional feature of the data but a regime-specific outcome.

Limits to Arbitrage and Financial Frictions. To explain UIP deviations, international macroeconomics has pivoted from frictionless models to frameworks emphasizing financial intermediaries. Bacchetta and Wincoop (2010b) show how infrequent portfolio decisions generate sluggish, delayed currency dynamics. More recently, Itskhoki and Mukhin (2021) provide an open-economy framework with segmented financial markets, where a UIP wedge emerges endogenously due to the restricted risk-bearing capacity of specialized financial intermediaries. My empirical findings provide direct macroeconomic evidence for

this mechanism. Consistent with the Global Financial Cycle (Miranda-Agrippino and Rey, 2020), I show that when U.S. monetary policy tightens, limits to arbitrage bind, generating the persistent UIP wedge that delays currency adjustment.

Methodological Advances in Identification. This paper builds on recent innovations that move beyond linear models. Studies using high-frequency identification (HFI) in linear SVAR frameworks, such as R  th (2020), generally find immediate overshooting. I provide evidence that these linear methods suffer from sample truncation bias. Instead, I adopt the continuous state-dependent local projection (SDLP) methodologies pioneered by Alessandri, Jord  , and Venditti (2025) and Cloyne, Jord  , and Taylor (2023). Furthermore, I address the central bank information effect by using the NLP-purged narrative shocks developed by Aruoba and Drechsel (2024). Integrating these tools isolates the structural financial wedge from the underlying frictionless macro response.

2.3 Empirical Strategy

2.3.1 Data Description and Macroeconomic State

I restrict the panel to the G10 currencies against the US dollar (AUD, CAD, CHF, EUR, GBP, JPY, NOK, NZD, SEK). Focusing on these advanced economies isolates limits-to-arbitrage mechanisms. Emerging market exchange rates face idiosyncratic constraints—such as capital controls, managed floats, and direct FX interventions—that fall outside the scope of this study. In deep and liquid G10 markets, protracted adjustments reflect structural constraints among global financial intermediaries rather than regulatory barriers.

To capture the slow-moving, structural accumulation of financial frictions, I construct the state variable, $\tilde{S}_t$, as a backward-looking, 5-year (60-month) moving average of the real interest rate gap. A prolonged, cumulative exposure to a policy regime alters deep financial frictions. The backward-looking trailing average filters out high-frequency endogenous noise

and avoids look-ahead bias, ensuring the macroeconomic state is predetermined and isolated from contemporaneous shocks.

For exogenous monetary policy shocks, the baseline analysis utilizes the narrative series constructed by C. Romer and D. Romer (2004) and extended by Wieland and Yang (2020). While the original series is orthogonal to the Federal Reserve’s Greenbook forecasts, recent literature indicates it may retain predictable responses to macroeconomic conditions omitted from the central bank’s internal projections (Coibion, 2012; Ramey, 2016). To avoid this problem, I project the raw shock series onto a set of observable macroeconomic controls, specifically six lags of U.S. industrial production growth and U.S. inflation. The residual from this projection serves as the baseline continuous monetary policy shock, ε_t . This procedure ensures the identifying variation is orthogonal to the predictable components of the U.S. business cycle.

Unlike High-Frequency Identification (HFI) data constrained to the post-1990 period, this extended narrative series provides broad historical coverage. The sample encompasses both the high-friction Volcker disinflation era and the subsequent low-friction Great Moderation, mitigating regime-specific bias. For robustness, I incorporate the narrative shocks developed by Aruoba and Drechsel (2024), which use Natural Language Processing (NLP) to filter out the Fed’s private economic assessments. To ensure the results are not driven by the choice of the interest rate gap, I also employ the Chicago Fed National Financial Conditions Index (NFCI). I construct a 5-year trailing average of the NFCI to provide a stable alternative proxy for structural limits to arbitrage.

The full sample period spans from January 1976 to December 2007. I restrict the sample at this point to exclude the Zero Lower Bound (ZLB) and unconventional monetary policies. These post-2008 environments introduced distinct market microstructures that confound traditional limits-to-arbitrage mechanisms.

I construct a monthly panel dataset of bilateral spot exchange rates and three-month

interbank rates. To evaluate deviations from UIP, I construct the cumulative excess return on the U.S. dollar dynamically on a month-by-month basis. Let e_t denote the log spot exchange rate (expressed as U.S. dollars per unit of foreign currency) and i_t (i_t^*) denote the U.S. (foreign) one-month interest rate yield. The h -month cumulative excess return, ER_{t+h} , is computed as:

$$ER_{t+h} = \sum_{j=0}^{h-1} (i_{t+j} - i_{t+j}^*) - (e_{t+h} - e_t) \quad (2.1)$$

This recursive formulation accurately replicates the ex-post realized return of a rolling-over carry trade strategy. The dataset incorporates industrial production and the Consumer Price Index (CPI) to control for country-specific business cycles.

2.3.2 State-Dependent Local Projections with KBO Decomposition

To evaluate the asymmetric responses of monetary policy shocks across macroeconomic regimes, I employ a state-dependent local projection (LP) framework (Jordà, 2005). Unlike standard structural vector autoregressions that impose symmetric dynamic restrictions, LP provides a flexible method for estimating impulse response functions (IRFs) across nonlinear environments. Building on recent literature (Cloyne, Jordà, and Taylor, 2023; Alessandri, Jordà, and Venditti, 2025), I adapt the Kitagawa-Blinder-Oaxaca (KBO) decomposition into this continuous panel time-series framework to formally disentangle the unconditional macroeconomic response from the friction-driven amplification.

Let $y_{i,t+h} - y_{i,t-1}$ denote the cumulative percentage change in the dependent variable (e.g., log spot exchange rate and excess return), and let ε_t represent the continuous exogenous monetary policy shock. The continuous, predetermined state variable is denoted by $\tilde{S}_{t-1}$, which captures the centered 5-year moving average of the US monetary policy stance ($S_{t-1} -$

$\bar{S}$). I estimate the following interacted panel LP model:

$$y_{i,t+h} - y_{i,t-1} = \alpha_{i,h} + \beta_h \varepsilon_t + \gamma_h \tilde{S}_{t-1} + \theta_h [\varepsilon_t \times \tilde{S}_{t-1}] + \Phi_h W_{i,t-1} + u_{i,t+h} \quad (2.2)$$

Here, $\alpha_{i,h}$ denotes country fixed effects. $W_{i,t-1}$ is a vector of predetermined macroeconomic controls, including lagged exchange rate changes, industrial production, and inflation. To preserve degrees of freedom, the state variable is interacted exclusively with the monetary policy shock.

A potential econometric concern is structural bias driven by endogenous state transitions. If the monetary policy shock alters the subsequent trajectory of the state, the decomposition confounds initial-state heterogeneity with transition-mediated effects. The 5-year moving-average state substantially mitigates this concern. Appendix B.2 shows that, unlike shorter windows, the baseline 5-year measure does not exhibit a statistically detectable response to monetary policy shocks over the projection horizon. The 5-year moving average provides the minimum structural smoothing required to keep the macroeconomic state strictly isolated from contemporaneous policy innovations. Furthermore, 5 years rolling window state variable is highly persistent with near unity autocorrelation (0.9995) and statistically invariant to policy shocks over the projection horizon.

The dynamic response of the exchange rate to a monetary policy shock, conditional on the state S , is given by $\hat{\beta}_h + S \cdot \hat{\theta}_h$. To plot the regime-specific dynamics, I evaluate this marginal effect at specific values. For instance, the response in a tight financial regime is constructed as $\hat{\beta}_h + S_{tight} \hat{\theta}_h$, while the response in a loose regime is $\hat{\beta}_h - S_{loose} \hat{\theta}_h$ (where S is evaluated at ± 1 standard deviation). This explicitly isolates how tighter financial conditions alter the trajectory of the exchange rate relative to the unconditional baseline. I employ Driscoll and Kraay (1998) standard errors to account for serial correlation and cross-sectional dependence.

2.4 Main Empirical Results

2.4.1 Asymmetric Exchange Rate Dynamics

Figure 2.1 plots the cumulative impulse response of the exchange rate following a one-standard-deviation contractionary monetary policy shock. I evaluate the response at a contractionary monetary policy stance (tight regime, +1 standard deviation of the centered state variable $\tilde{S}_{t-1}$) and an expansionary stance (loose regime, -1 standard deviation).

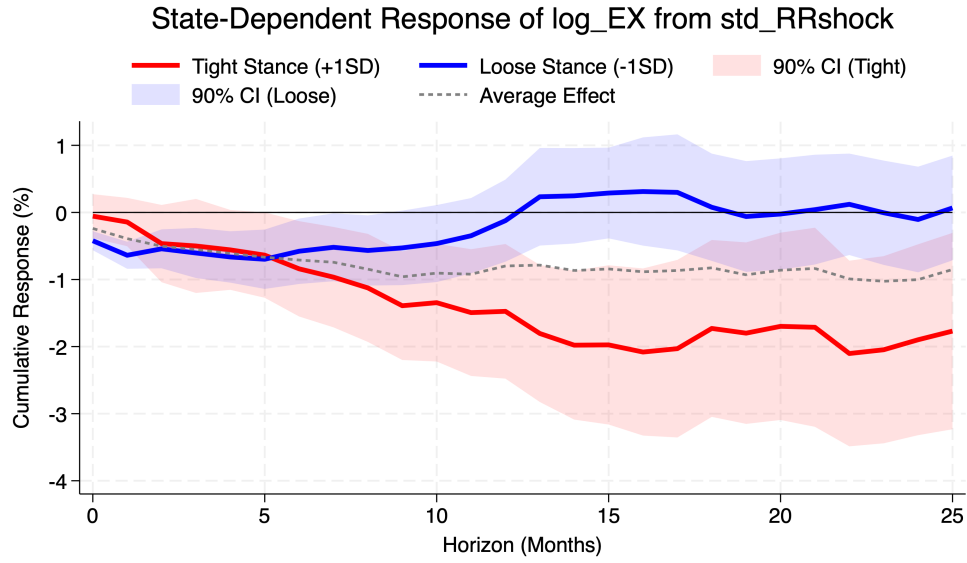
The exchange rate exhibits severe state-dependent asymmetry. Under a contractionary monetary policy stance (red solid line), the exchange rate exhibits a muted initial response, followed by a gradual appreciation that peaks at approximately 1.5 percent 15 months after the initial shock. In contrast, under an expansionary stance (blue solid line), the exchange rate adjusts more rapidly, peaking within the first 5 months before gradually depreciating.

Panel (b) isolates the indirect effect, which quantifies the additional exchange rate response attributable to the constrained state. The indirect effect is statistically indistinguishable from zero on impact. As the horizon extends, it diverges, contributing an additional 1.0 to 1.2 percentage points of cumulative appreciation between months 13 and 17. The persistent negative wedge drives the delayed overshooting peak, supporting structural limits to arbitrage as the primary driver of delayed overshooting under tight macroeconomic conditions.

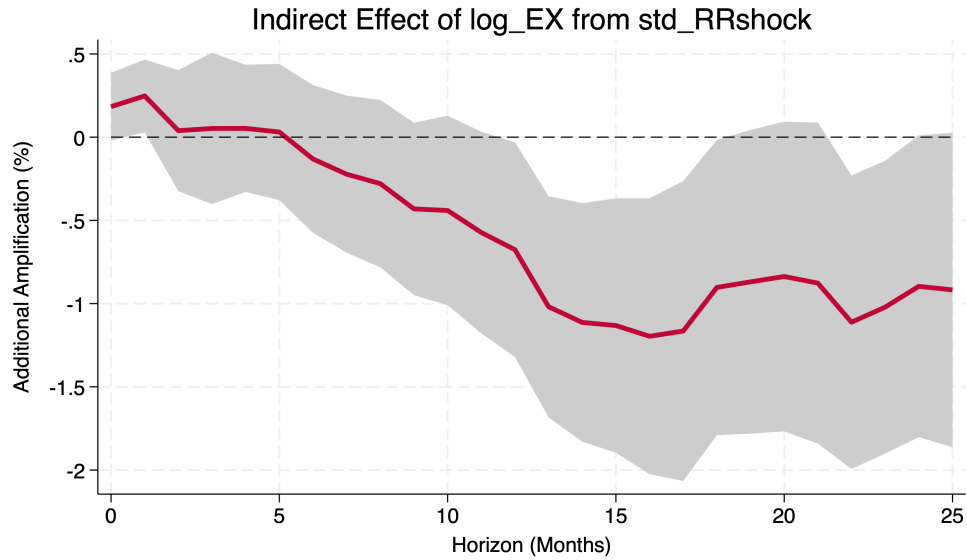
2.4.2 Asymmetric Excess Return Dynamics

Under the standard UIP condition, an exogenous monetary policy shock should not generate predictable excess returns; the exchange rate should jump on impact and subsequently depreciate to offset the interest rate differential. Figure 2.2 documents the asymmetric response of the cumulative excess return.

Under an expansionary monetary policy stance, the cumulative excess return remains statistically indistinguishable from zero across the majority of the projection horizon. When

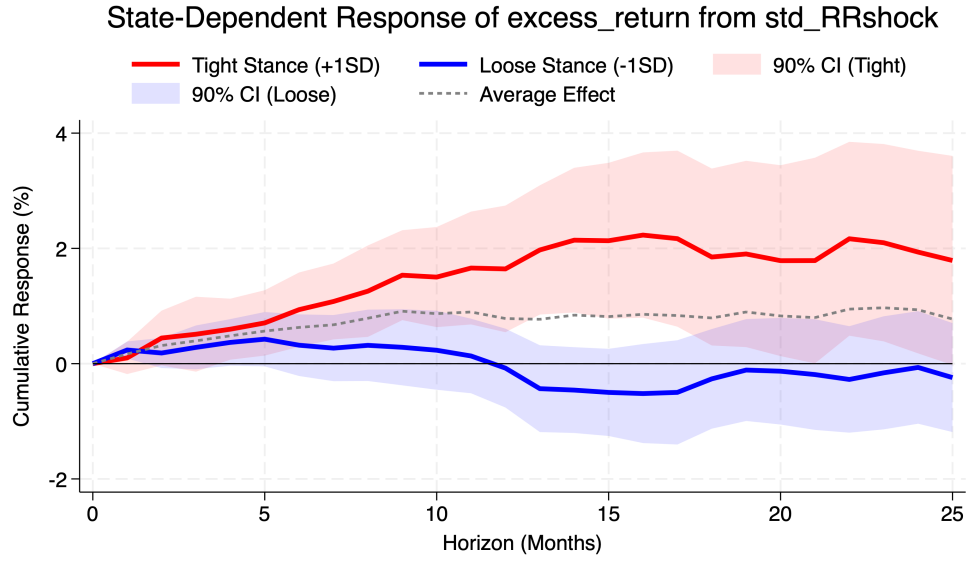


(a) State-Dependent Total Effects

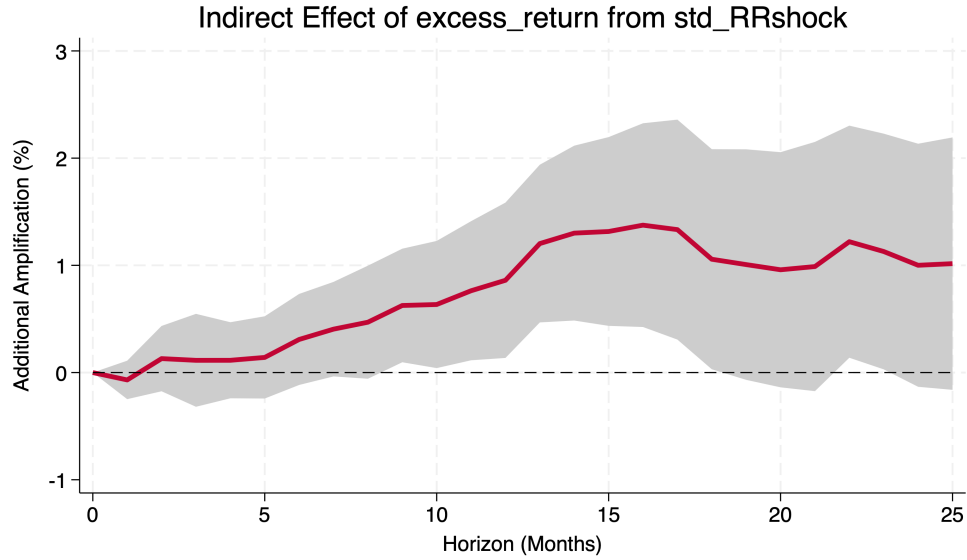


(b) Friction-driven Amplification (Indirect Effect)

Figure 2.1: **State-Dependent Exchange Rate Dynamics and Friction-driven Amplification.** Panel (a) displays the cumulative impulse response of the log exchange rate to a one-standard-deviation contractionary Romer and Romer (2004) shock. The red solid line indicates the response evaluated at a tight stance (+1 SD), and the blue solid line indicates a loose stance (-1 SD). Panel (b) isolates the indirect effect, which quantifies the additional exchange rate response mechanically attributable to the constrained macroeconomic state. Shaded areas indicate 90% and 95% confidence intervals.



(a) State-Dependent Total Effects



(b) Friction-driven Amplification (Indirect Effect)

Figure 2.2: **State-Dependent Excess Return Dynamics and UIP Deviations.** Panel (a) displays the cumulative impulse response of the realized excess return on the US dollar to a one-standard-deviation contractionary Romer and Romer (2004) shock. The red solid line indicates the response evaluated at a tight stance (+1 SD), and the blue solid line indicates a loose stance (-1 SD). Panel (b) isolates the indirect effect. Shaded areas indicate 90% and 95% confidence intervals.

aggregate liquidity is abundant, unconstrained financial intermediaries readily absorb policy shocks, facilitating an adjustment consistent with the UIP benchmark.

Conversely, under a contractionary stance, the shock generates a persistent and predictable excess return that peaks at approximately 2.1 percentage points. Panel (b) shows that the indirect effect accounts for an additional 1.0 to 1.3 percentage points of this deviation. The presence of this sustained excess return indicates that the delayed exchange rate appreciation translates into unexploited arbitrage profits. This dynamic is consistent with an environment where global financial intermediaries face binding balance sheet constraints, preventing them from closing the return differential.

2.5 Robustness Checks

To ensure the baseline findings are robust to alternative specifications and sample periods, I conduct three exercises. First, I test the sensitivity of the results to an alternative monetary policy shock. Second, I replace the baseline macroeconomic state variable with a direct measure of aggregate financial conditions. Finally, I estimate the model on a post-Volcker subsample to confirm that the state-dependent delayed overshooting is not a historical artifact of the 1980s.

2.5.1 Alternative Shock Identification: Purging Central Bank Information Effects

Identifying monetary policy shocks is often confounded by the central bank information effect (Nakamura and Steinsson, 2018). If a contractionary policy surprise inadvertently signals the Federal Reserve’s pessimistic outlook on macroeconomic fundamentals, market participants might incorporate this information gradually. This learning process could induce a sluggish adjustment in asset prices, independent of structural financial frictions.

To evaluate whether the observed delayed overshooting is an artifact of this informational

channel, I re-estimate the models using the narrative shock series developed by Aruoba and Drechsel (2024). Their methodology uses natural language processing (NLP) to filter internal economic assessments from the Federal Reserve’s historical documents, isolating the exogenous policy surprise. Figure 2.3 presents the state-dependent responses and the corresponding indirect effects using this NLP-purged shock.

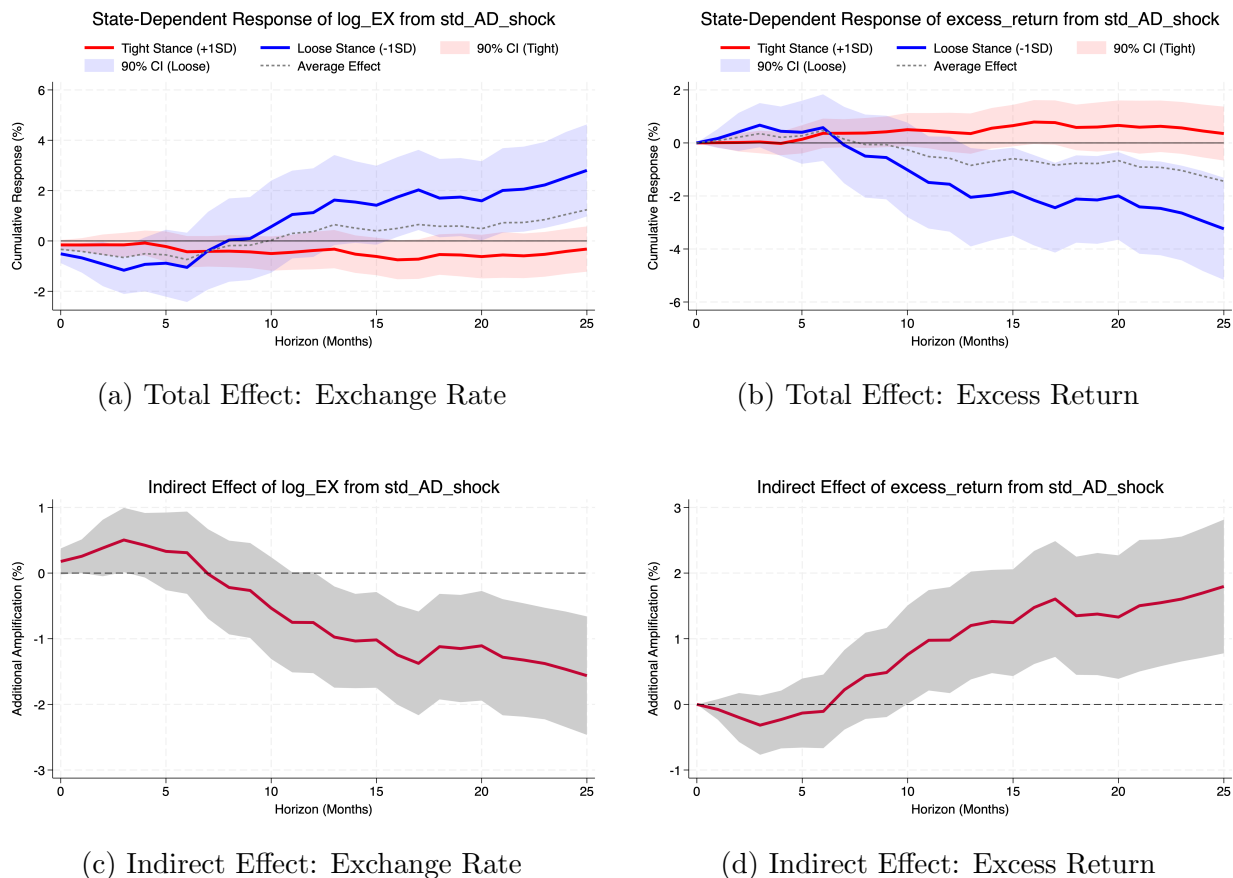


Figure 2.3: **Robustness to Information Effects: AD (2024) NLP Shocks.** The figure displays the state-dependent local projections using the NLP-purged monetary policy shock by Aruoba and Drechsel (2024). Panels (a) and (c) present the total and indirect effects for the cumulative exchange rate response. Panels (b) and (d) present the corresponding effects for the cumulative excess return. The red solid lines indicate the response evaluated at a tight stance (+1 SD), and the blue solid lines indicate a loose stance (-1 SD). Shaded areas denote 90% confidence intervals.

Panels (a) and (b) display the state-dependent total effects for the exchange rate and excess return, respectively. Compared to the baseline estimates derived from the C. Romer and D. Romer (2004) series, the estimates are less precise, yielding wider confidence bands.

This wider uncertainty stems from two sample constraints: the NLP-based instrument is restricted to the post-1982 period (excluding the early Volcker era), and the filtration process removes systematic variation from the instrument, reducing estimation efficiency.

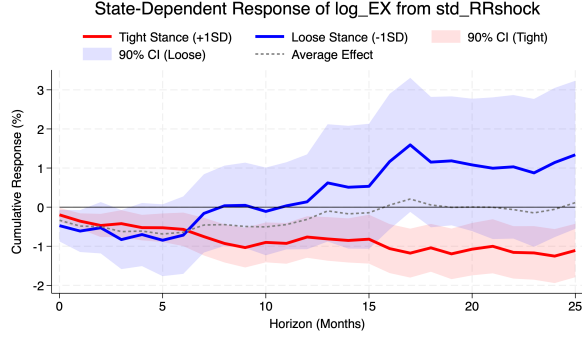
Despite the wider confidence intervals for the aggregate responses, the structural wedge generated by the macroeconomic state remains consistent. Panels (c) and (d) isolate the indirect effect for both variables. Under a contractionary regime, the indirect effect generates a protracted appreciation (Panel c) and a sustained predictable excess return (Panel d) over the medium-term horizon. This confirms that the delayed overshooting puzzle is driven by limits to arbitrage rather than the delayed absorption of central bank information.

2.5.2 Alternative State Variable: National Financial Conditions Index

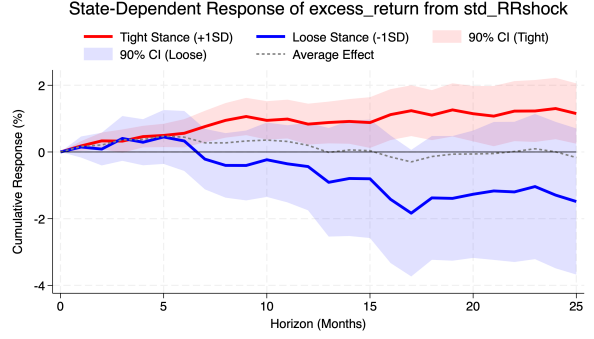
The baseline analysis uses the 5-year trailing average of the interest rate gap ($r - r^*$) to capture the structural accumulation of financial frictions. Because this measure relies on unobserved estimates of the natural rate of interest, I verify the findings by substituting the baseline state variable with the Chicago Fed National Financial Conditions Index (NFCI). The NFCI aggregates over 100 indicators of risk, liquidity, and leverage, providing a direct measure of the risk-bearing capacity of financial intermediaries.

I construct the alternative state variable as a 5-year backward-looking moving average of the NFCI to isolate the slow-moving accumulation of limits to arbitrage, filtering out transient financial market noise. Historically, this smoothed NFCI series exhibits a positive correlation with the baseline monetary policy stance (see Appendix B), confirming that prolonged contractionary policy regimes translate into tight aggregate financial conditions. Figure 2.4 presents the state-dependent local projections using this NFCI state variable and the baseline C. Romer and D. Romer (2004) shock.

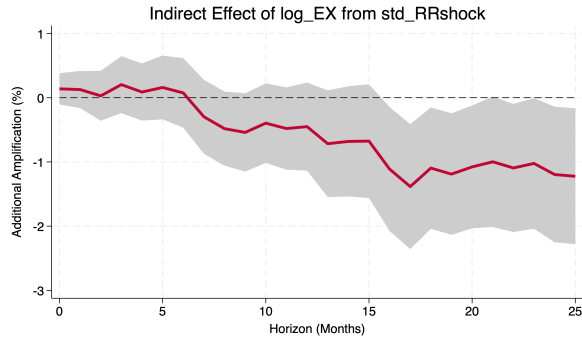
The empirical patterns are consistent with the baseline findings. In Panel (a), under



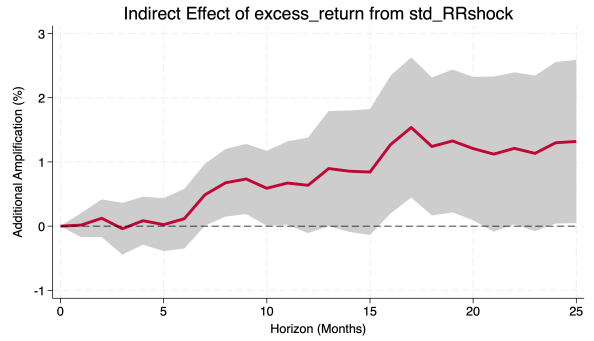
(a) Total Effect: Exchange Rate



(b) Total Effect: Excess Return



(c) Indirect Effect: Exchange Rate



(d) Indirect Effect: Excess Return

Figure 2.4: **Robustness to an Alternative State Variable: NFCI.** The figure displays the state-dependent local projections using the 5-year trailing average of the Chicago Fed National Financial Conditions Index (NFCI) as the conditioning macroeconomic state. The baseline C. Romer and D. Romer (2004) shock is utilized. The left column presents the total and indirect effects for the exchange rate, and the right column presents the corresponding effects for the excess return. The red solid lines indicate the response evaluated at tight financial conditions (+1 SD), and the blue solid lines indicate loose financial conditions (-1 SD). Shaded areas denote 90% confidence intervals.

a tight NFCI regime (red solid line), the exchange rate response remains muted on impact and gradually appreciates, reaching a delayed peak approximately 17 to 20 months after the initial shock. Under loose financial conditions (blue solid line), the initial appreciation is smaller, and the exchange rate subsequently depreciates.

Panel (c) isolates the friction-driven wedge for the exchange rate. Evaluated at the tight NFCI regime, the indirect effect contributes a persistent negative wedge that accounts for 1.2 to 1.4 percentage points of additional delayed appreciation around month 17. For the cumulative excess return, tight financial conditions generate a predictable deviation from UIP (Panel b). Panel (d) attributes this violation to limits to arbitrage; the indirect effect exhibits an upward trajectory, reaching approximately 1.5 percentage points over the medium term.

By confirming that a direct measure of aggregate financial friction yields a similar amplification channel, this exercise substantiates that the delayed overshooting puzzle is driven by structural limits to arbitrage that bind during restrictive macroeconomic environments.

2.5.3 Post-Volcker Period Analysis

Kim, Moon, and Velasco (2017) argue that the delayed overshooting puzzle is a historical anomaly. Because classic empirical findings rely heavily on data from the early 1980s, critics attribute protracted currency dynamics to the extreme monetary tightening and macroeconomic volatility of the Volcker disinflation. To show that state-dependent delayed overshooting is not an artifact of the Volcker era, I estimate the model on the Great Moderation sample (January 1990 to December 2007) using both the C. Romer and D. Romer (2004) narrative shocks and the Aruoba and Drechsel (2024) NLP-purged shocks.

The state-dependent asymmetry remains evident in the post-1990 subsample using the Romer and Romer shocks (Figure 2.5). Under a contractionary policy stance, the exchange rate still adjusts sluggishly. The indirect effect isolates a persistent limits-to-arbitrage wedge

that drives predictable UIP deviations over the medium term. The local projection estimates yield wider confidence intervals than the full baseline. This loss of precision reflects the Great Moderation itself: lower macroeconomic volatility and smaller policy shocks reduce the identifying variation available to estimate the friction-driven wedge. The indirect effect, however, maintains its characteristic shape: it remains muted on impact and builds over the following year, capturing the mechanics of delayed overshooting.

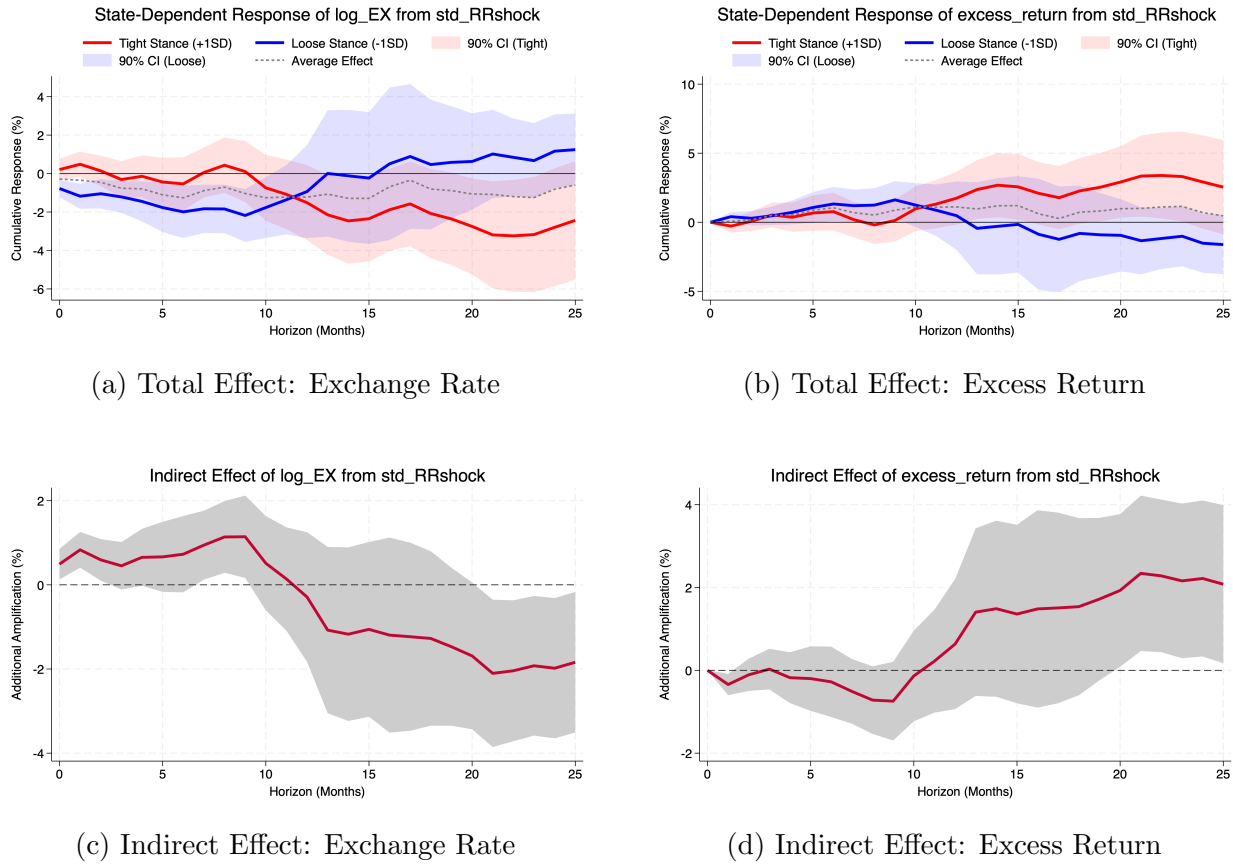


Figure 2.5: **Robustness to Subsample: Post-Volcker Period (RR Shocks).** The figure displays the state-dependent local projections using the baseline C. Romer and D. Romer (2004) shock restricted to the Great Moderation period (January 1990 to December 2007). The red solid lines indicate the response evaluated at a tight stance (+1 SD), and the blue solid lines indicate a loose stance (-1 SD). Shaded areas denote 90% confidence intervals.

These post-Volcker dynamics survive purging the central bank information effect (Figure 2.6). Using the Aruoba and Drechsel (2024) shocks, the indirect effect still generates delayed appreciation under tight financial conditions. Delayed overshooting is therefore not an iso-

lated anomaly of the early 1980s. It is a persistent structural pattern governed by aggregate financial conditions, operating consistently even during the tranquil Great Moderation.

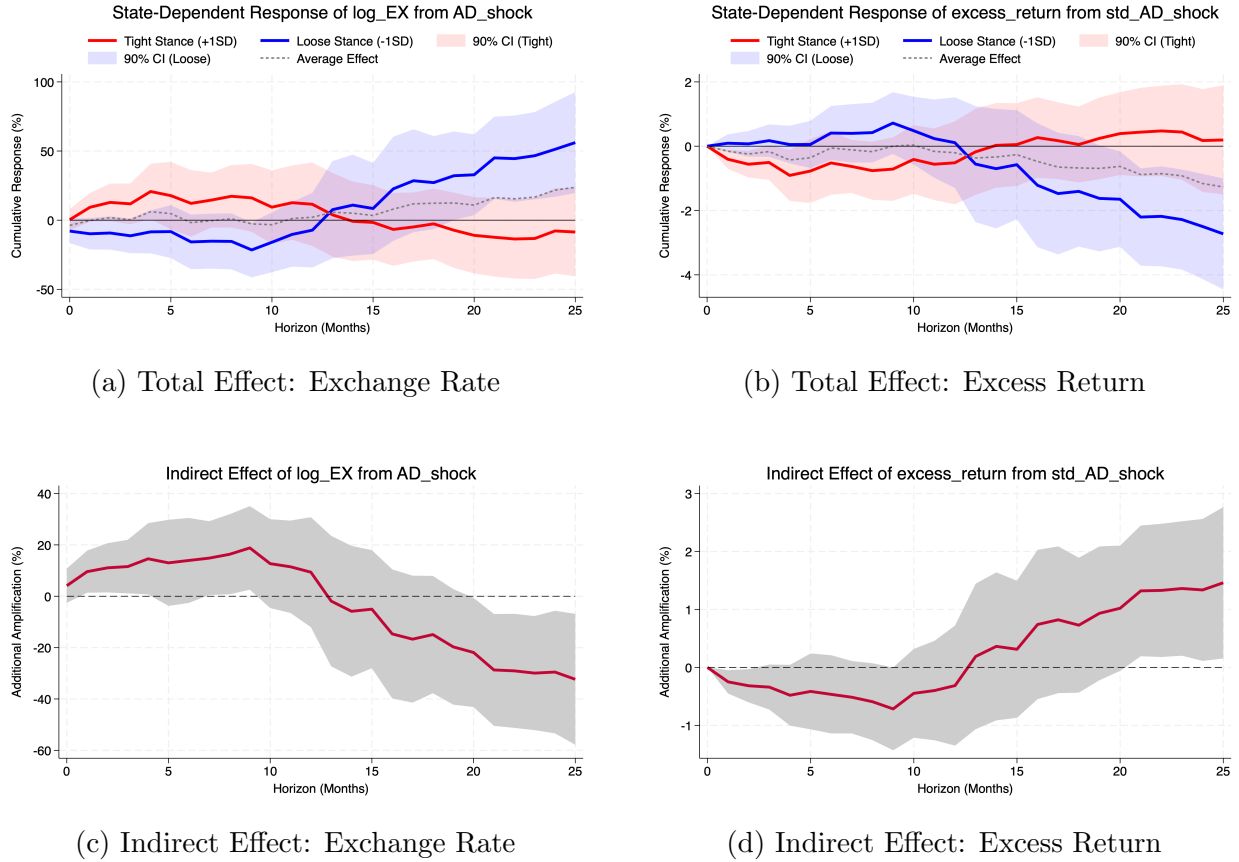


Figure 2.6: **Robustness to Subsample: Post-Volcker Period (AD Shocks).** The figure displays the state-dependent local projections using the NLP-purged monetary policy shock by Aruoba and Drechsel (2024) restricted to the Great Moderation period (January 1990 to December 2007). The red solid lines indicate the response evaluated at a tight stance (+1 SD), and the blue solid lines indicate a loose stance (-1 SD). Shaded areas denote 90% confidence intervals.

2.6 Monte Carlo Analysis of Truncated Sample Bias

The empirical evidence establishes that delayed overshooting is a state-dependent phenomenon. This challenges recent literature that estimates linear structural vector autoregressive (SVAR) models using high-frequency identification (HFI) techniques.

Studies such as R  th (2020) rely on proxy-SVAR (SVAR-IV) frameworks, instrumenting

reduced-form VAR residuals with high-frequency changes in interest rate futures. While this approach effectively isolates unanticipated policy surprises, the required tick-level data restricts these instruments strictly to the post-1990 period. This modern sample coincides with the Great Moderation—an era characterized by abundant global liquidity and low financial frictions. Applying a linear, regime-invariant SVAR to this truncated instrument set poses an econometric problem: it inadvertently extrapolates low-friction dynamics to the broader historical population.

2.6.1 Analytical Framework of the Truncated Instrument Bias

Before evaluating the quantitative simulation, it is necessary to formalize the econometric mechanism that drives the sample truncation bias in linear SVAR-IV models. In a standard structural VAR framework, a researcher first estimates a reduced-form VAR, $Y_t = B(L)Y_{t-1} + u_t$, to recover the one-step-ahead prediction errors, u_t . Because macroeconomic variables respond simultaneously to news, the elements of u_t are contemporaneously correlated and lack a direct economic interpretation. Instead, they are linear combinations of the unobserved, mutually orthogonal structural shocks, ε_t .

Suppose the true data generating process is state-dependent, transitioning between a historical high-friction regime ($S_t = A$) and a modern low-friction regime ($S_t = B$). The mapping from structural shocks to reduced-form residuals is governed by a regime-specific contemporaneous impact matrix $\Phi^{(S_t)}$:

$$u_t = \Phi^{(S_t)} \varepsilon_t, \quad \text{for } S_t \in \{A, B\} \quad (2.3)$$

Let $\varepsilon_{1,t}$ denote the structural monetary policy shock. The parameter of interest is the relative impulse response on impact of a target variable i (e.g., the exchange rate) normalized by a policy indicator r (e.g., the interest rate). The true relative response for regime S_t is given by the ratio of the respective elements in the first column of the impact matrix:

$$\theta^{(S_t)} = \Phi_{i,1}^{(S_t)} / \Phi_{r,1}^{(S_t)}.$$

In a pooled SVAR-IV estimation, researchers employ an external instrument Z_t that satisfies the relevance condition ($\mathbb{E}[Z_t \varepsilon_{1,t}] = \alpha \neq 0$) and the exogeneity condition ($\mathbb{E}[Z_t \varepsilon_{j,t}] = 0$ for $j \neq 1$). However, modern HFI instruments are systematically recorded only during the modern era ($S_t = B$). During the historical era ($S_t = A$), the instrument is unobserved, effectively yielding $Z_t = 0$.

The structural impact vector is identified via the unconditional sample covariance between the reduced-form residuals and the external instrument. As the sample size grows, the probability limit of this covariance decomposes via the Law of Total Expectation:

$$\mathbb{E}[u_t Z_t] = \Pr(S_t = A) \mathbb{E}[u_t Z_t \mid S_t = A] + \Pr(S_t = B) \mathbb{E}[u_t Z_t \mid S_t = B] \quad (2.4)$$

Because the instrument lacks variation during the historical regime ($Z_t = 0$), the first term evaluates to zero. Substituting the structural mapping $u_t = \Phi^{(B)} \varepsilon_t$ into the surviving second term yields:

$$\mathbb{E}[u_t Z_t] = \Pr(S_t = B) \cdot \Phi^{(B)} \mathbb{E}[\varepsilon_t Z_t \mid S_t = B] = (\Pr(S_t = B) \alpha) \Phi_{\bullet,1}^{(B)} \quad (2.5)$$

The standard SVAR-IV Wald estimator computes the relative response by taking the ratio of these covariances. Consequently, the asymptotic limit of the pooled SVAR-IV estimator for variable i is:

$$\text{plim } \hat{\theta}_i^{IV} = \frac{(\Pr(S_t = B) \alpha) \Phi_{i,1}^{(B)}}{(\Pr(S_t = B) \alpha) \Phi_{r,1}^{(B)}} = \frac{\Phi_{i,1}^{(B)}}{\Phi_{r,1}^{(B)}} = \theta^{(B)} \quad (2.6)$$

Equation (4) formalizes the sample truncation bias. The scaling factor $\Pr(S_t = B) \alpha$ cancels out in the ratio. The pooled SVAR-IV estimator does not recover a weighted unconditional average of the historical and modern dynamics; rather, it restricts the estimated

structural parameter to converge strictly to the low-friction transmission vector ($\theta^{(B)}$).

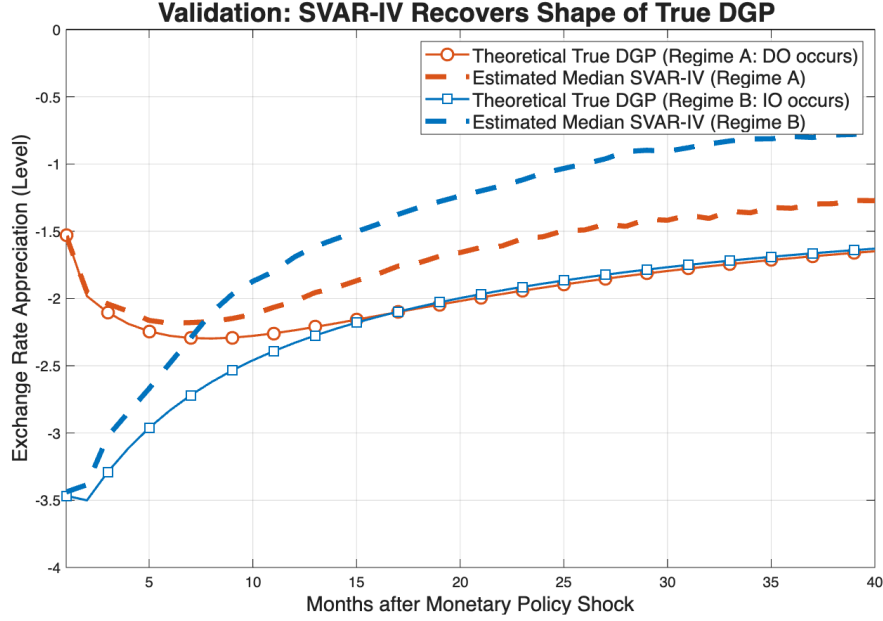
2.6.2 Quantitative Simulation Results

To quantify the magnitude of this truncation bias on dynamic impulse responses, I conduct a Monte Carlo simulation based on a state-dependent Data Generating Process (DGP) inspired by the segmented financial market framework of Itskhoki and Mukhin (2021). Because the primary focus of this chapter is empirical, the formal theoretical micro-foundation is developed in Chapter 3, and the full system of dynamic equations is relegated to Appendix B.1. The economic intuition is straightforward: a contractionary shock impairs the risk-bearing capacity of global intermediaries, raising the shadow cost of capital for arbitrageurs and widening the UIP wedge.

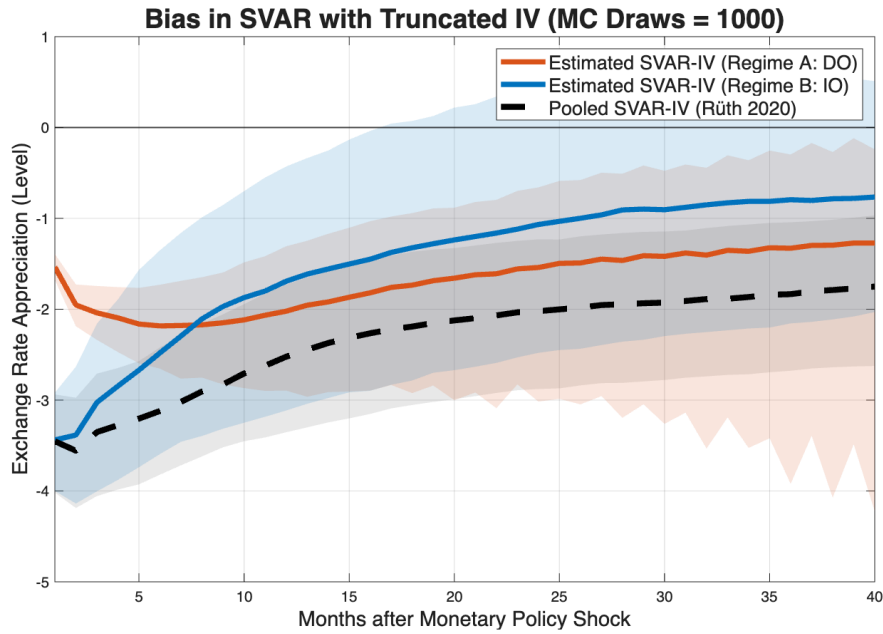
The simulation mimics the data constraints faced by recent empirical studies. The DGP generates a sample transitioning from a high-friction regime (Regime A) to a low-friction regime (Regime B). To replicate the reliance on modern HFI data, the exogenous instrument is deliberately truncated so that it is observable only during Regime B. I then estimate a standard, time-invariant linear SVAR-IV model over the entire pooled sample.

Figure 2.7 presents the results of the Monte Carlo exercise. Panel (a) validates the theoretical accuracy of the SVAR-IV estimator. When applied separately to the correct subsamples, the estimator accurately recovers the true underlying DGP for both the high-friction environment (Regime A, displaying delayed overshooting) and the low-friction environment (Regime B, displaying immediate overshooting).

Panel (b) illustrates the methodological bias induced by the truncated instrument in a pooled estimation. Because the instrument is observed exclusively during the low-friction era, the estimation algorithm identifies the structural vector corresponding to Regime B, exactly as predicted by Equation (4). As a result, the pooled model (black dashed line) yields an immediate overshooting response, omitting the friction-driven delayed dynamics



(a) Validation: SVAR-IV Recovers True DGP in Sub-samples



(b) Methodological Illusion: Pooled SVAR-IV with Truncated IV

Figure 2.7: **Monte Carlo Simulation of Truncated Sample Bias.** Panel (a) plots the theoretical true Data Generating Process (DGP) alongside the SVAR-IV estimates when applied correctly to separate sub-samples. Panel (b) displays the estimated impulse response from a pooled linear SVAR-IV model utilizing an instrument that is available only during the low-friction period (Regime B). The black dashed line indicates that the pooled estimator mechanically converges to the low-friction dynamics, masking the delayed overshooting present in Regime A.

present in the first half of the simulated sample.

Using this simulation, I show that using truncated instruments in a linear SVAR yields a regime-specific bias that masks the underlying delay. Delayed overshooting remains a robust feature of the data, obscured only by linear econometric models estimated over structurally benign historical samples.

2.7 Implications

The empirical evidence that delayed overshooting is a state-dependent phenomenon imposes concrete constraints on how central banks evaluate the transmission of monetary policy. Moving beyond unconditional averages, the findings yield three specific policy and methodological implications.

First, from the perspective of the U.S. Federal Reserve, the state-dependency alters the timing of the exchange rate channel. Under contractionary environments, the U.S. dollar's peak appreciation is delayed by approximately 15 to 20 months following a policy shock. Since an appreciating dollar transmits disinflationary pressure through lower import prices, this sluggish adjustment implies that the full macroeconomic effect of a rate hike operates with a significant lag. Policymakers relying on the immediate, muted exchange rate response may systematically underestimate the tightening already in the pipeline, increasing the risk of over-tightening during constrained financial conditions.

Second, the protracted exchange rate dynamics complicate the policy response of foreign economies facing U.S. monetary spillovers. When the Federal Reserve tightens during an already restrictive global cycle, trading partners and emerging markets experience prolonged, rather than immediate, currency depreciation. For foreign central banks targeting inflation or financial stability, this indicates that imported inflation and capital outflow pressures will persist over a multi-quarter horizon. Consequently, front-loaded policy responses, such as immediate foreign exchange interventions, may prove premature and ineffective against a

friction-driven, slowly building currency trend.

Finally, the findings highlight a methodological limitation in recent empirical literature that estimates linear structural vector autoregressions (SVAR) using high-frequency identification (HFI). Pioneering studies in both open-economy (e.g., R  th, 2020) and closed-economy macroeconomics (e.g., Gertler and Karadi, 2015; Nakamura and Steinsson, 2018) rely on high-frequency financial instruments that are predominantly available only post-1990. Because this truncated sample heavily overlaps with the low-friction environment of the Great Moderation, applying these instruments within a linear, regime-invariant framework mechanically estimates the frictionless, immediate adjustment. The simulation in Section 6 demonstrates that this methodological design implicitly extrapolates structurally benign dynamics to the broader historical sample, systematically masking the delayed transmission inherent in tight macroeconomic states.

Chapter 3

Bridging Local Housing and Global Dollar Markets: The Intermediary Channel

3.1 Introduction

Structurally, housing markets and dollar funding markets operate in distinct economic spheres. Housing is local, illiquid, and tied to domestic collateral values. It is not traded in global financial markets, and its price is shaped by domestic credit conditions, local income risk, and the supply of real estate. Covered interest parity (CIP) deviations, by contrast, are prices in one of the most globalized segments of financial markets. They measure violations of a no-arbitrage condition linking spot exchange rates, forward exchange rates, and interest rates. Since the global financial crisis, a large literature interprets these deviations as evidence of limits to arbitrage and scarce intermediary balance sheet capacity in dollar funding markets (Du, Tepper, and Verdelhan, 2018).

This chapter starts from a simple empirical observation: these two markets are discon-

nected in normal times, but reconnect in periods of global financial stress. Across a panel of advanced open economies, housing price growth has little systematic relationship with CIP deviations when financial conditions are calm. When global risk rises, however, the relationship turns negative. Countries experiencing weaker housing markets also experience larger dollar funding premia. The same domestic collateral variable that appears irrelevant in normal times becomes informative when intermediary risk-bearing capacity is scarce.

This state-dependence challenges standard frameworks for two reasons. First, housing is not itself the asset being priced in the FX swap market. A decline in domestic real estate values does not mechanically change the payoff of a currency forward contract. Second, standard open-economy models typically separate domestic collateral frictions from international arbitrage wedges. Housing affects borrowing, consumption, and domestic demand; CIP deviations are often treated as exogenous financial wedges or as the outcome of global dollar scarcity. The evidence in this chapter suggests that these two objects are jointly determined in stress states. A local collateral shock becomes a global funding-market price only when it loads onto the balance sheet of the marginal intermediary.

To document this state-dependence, I estimate panel regressions that interact housing price growth with measures of global financial stress. The results indicate that the housing-CIP relationship is state-dependent. In normal times, housing price growth has an insignificant effect on CIP deviations. In stress periods, however, a decline in housing prices is associated with a wider CIP deviation, robust to controlling for domestic macroeconomic conditions and time fixed effects.

Theoretically, I first develop a parsimonious two-period small open economy model featuring a constrained intermediary, based on Gabaix and Maggiori (2015b). Households borrow against domestic income and collateral. An intermediary supplies domestic credit and holds foreign safe assets. The intermediary is subject to a Value-at-Risk style balance sheet constraint, so the shadow value of risk capacity depends on both domestic credit exposure

and dollar liquidity positions.

The model has two regimes. When the intermediary is well capitalized, the balance sheet constraint is slack. The intermediary absorbs domestic mortgage risk and foreign exchange order flow elastically, so domestic housing fluctuations do not move CIP deviations. This is the normal-times disconnect. When intermediary net worth falls below a threshold, the constraint binds. A deterioration in domestic mortgage exposure tightens the intermediary's balance sheet, and increases the shadow value of dollar liquidity. In this state, the intermediary demands foreign safe assets despite their high shadow cost, generating a CIP deviation. The same domestic shock therefore has an insignificant impact in normal times but an amplified impact in stress states.

To evaluate this mechanism quantitatively, I embed the two-period model's intuition into an infinite-horizon Small Open Economy New Keynesian model. The framework features heterogeneous households to capture domestic housing market frictions, following Iacoviello (2005). Patient savers supply deposits, while impatient borrowers take out mortgage loans subject to housing collateral constraints. A global financial intermediary, based on Gertler and Karadi (2011), connects these domestic credit markets to international dollar funding markets. This intermediary manages both domestic mortgages and hedged dollar assets subject to an occasionally binding Value-at-Risk constraint.

I solve the occasionally binding constraint using the piecewise-linear approach of Guerrieri and Iacoviello (2015) to capture the nonlinear macroeconomic dynamics. Consistent with the two-period model, in normal times, the intermediary remains unconstrained, allowing the domestic housing market to evolve independently of the global dollar funding premium. However, a sufficiently large decline in housing prices depletes intermediary net worth and triggers the VaR constraint. This state-transition generates a macro-financial feedback loop: the binding constraint simultaneously forces aggregate deleveraging in the domestic economy and restricts the supply of dollar liquidity, endogenously widening the CIP basis and

depreciating the exchange rate.

The chapter is organized as follows. Section 2 reviews the related literature. Section 3 documents the motivating empirical evidence, establishing the state-dependent relationship between domestic housing markets and the global dollar premium. Section 4 constructs a two-period parsimonious model to analytically clarify the transmission channel. Section 5 extends this insight into a full general equilibrium DSGE model, and Section 6 presents the quantitative exercise to evaluate the model’s ability to replicate the empirical threshold dynamics. Section 7 concludes.

3.2 Related Literature

This chapter connects three closely related strands of literature. The first studies deviations from covered interest parity (CIP) and global dollar funding markets. In frictionless international financial markets, CIP holds by arbitrage: borrowing in one currency, swapping into another, investing at the foreign rate, and locking in the forward exchange rate delivers the same return as domestic lending. The persistent violations documented by Du, Tepper, and Verdelhan (2018) indicate that this arbitrage is severely limited in practice. Their evidence points to intermediary balance sheet costs, regulatory constraints, and quarter-end pressures as central determinants of the basis. This chapter builds on that insight but shifts the focus from the existence of the basis to its macro-financial state-dependence. I investigate whether domestic collateral conditions explain the time-varying magnitude of dollar funding premia. The evidence indicates a strict state-dependence: local collateral shocks correlate with global pricing anomalies only during stress episodes.

The second literature studies exchange rates and financial prices through the balance sheets of intermediaries. Gabaix and Maggiori (2015b) emphasize that international financiers absorb imbalances in global asset markets, implying that capital flows affect prices through their impact on intermediary balance sheets. Adrian and Shin (2014) and He and

Krishnamurthy (2013) similarly highlight the role of intermediary risk capacity in generating nonlinear financial amplification. This chapter applies that logic to a specific open-economy arbitrage price (CIP basis).

Third, this chapter bridges the macroeconomic literature on the exchange rate disconnect with models of collateral constraints. Itskhoki and Mukhin (2021) demonstrate that the disconnect between exchange rates and macro fundamentals is a natural consequence of exogenous financial shocks in segmented markets. Complementing this, Lilley et al. (2022) documents that exchange rates remain disconnected in normal times but systematically reconnect with global risk factors during financial crises. While these open-economy frameworks treat the financial wedge as exogenous, closed-economy and sudden-stop models (Iacoviello, 2005; Mendoza, 2010) demonstrate how declines in local collateral values tighten borrowing constraints and drive severe macroeconomic downturns. This chapter integrates these two distinct views. It provides a microfoundation for the international financial wedge by showing that both the normal-times disconnect and the crisis-state reconnect emerge endogenously when domestic frictions interact with the risk-bearing capacity of global intermediaries.

3.3 Empirical Motivation

This section documents the empirical motivation for the state-dependent behavior between domestic asset markets and the foreign exchange market. Rather than claiming a clean causal identification, the empirical exercise here serves to discipline the model along two core dimensions: the normal-time asset-FX disconnect and the stress-time reconnect. These empirical stylized facts establish the quantitative targets that any micro-founded intermediary model must satisfy.

Figure 3.1 illustrates the central motivating fact of this chapter. The figure plots year-over-year real asset price growth (using housing prices as a proxy for domestic collateral and

corporate balance sheet strength) against one-year CIP deviations vis-à-vis the US dollar for G10 economies from 2000 to 2023. Blue circles denote normal months, while red triangles denote months of severe global financial stress, measured using the upper tail (top 20%) of the VIX. In normal periods, the slope is close to flat for most economies. During stress periods, however, the slope steepens: domestic asset price declines coincide with wider CIP deviations. This contrast is especially visible for open economies such as Canada, Norway, New Zealand, Sweden, and the United Kingdom.

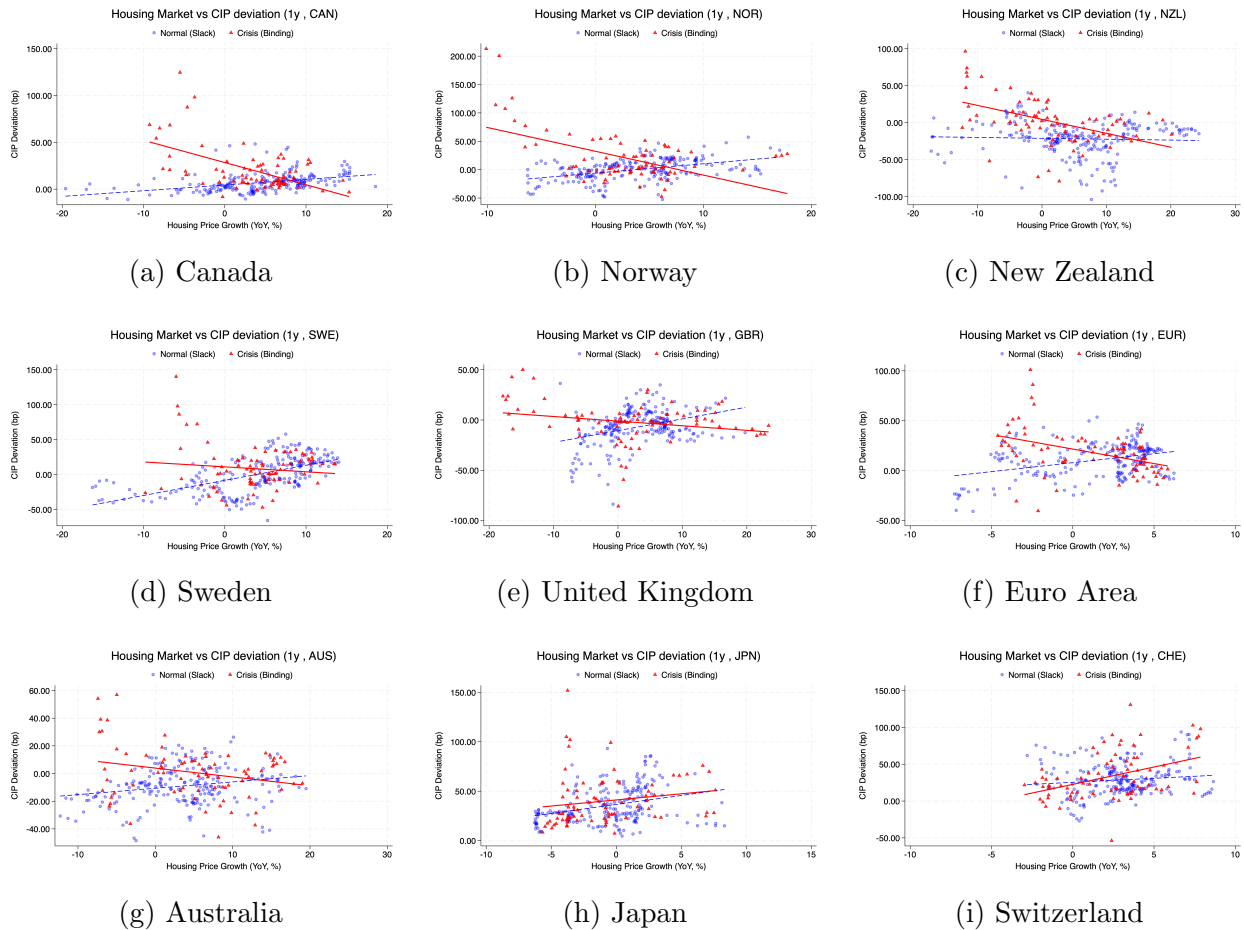


Figure 3.1: State-dependent relationship between housing price growth and CIP deviations. Blue circles denote normal periods, while red triangles denote periods of elevated global financial stress. The slope steepens during stress periods for typical open economies, while remaining flat or inverted for safe-haven currencies.

The behavior of safe-haven currencies provides further insight into this mechanism. Japan and Switzerland do not display the same crisis-state pattern. This divergence is

informative: if the channel operates through intermediary demand for dollar liquidity during domestic stress, it should be strongest for currencies whose domestic downturns are associated with capital outflows and dollar funding pressure. Conversely, the effect should be weaker, absent, or even reversed for safe-haven currencies that attract capital inflows when global risk rises. This cross-country heterogeneity points toward a balance sheet channel rather than a direct macroeconomic linkage.

To formalize this observation, I estimate panel regressions of the following form:

$$\begin{aligned}
b_{i,t}^{CIP} = & \alpha_i + \tau_t + \beta_1 \Delta q_{i,t} + \beta_2 VIXStress_t + \beta_3 (\Delta q_{i,t} \times VIXStress_t) \\
& + \beta_4 (\Delta q_{i,t} \times Safe_i) + \beta_5 (VIXStress_t \times Safe_i) + \beta_6 (\Delta q_{i,t} \times VIXStress_t \times Safe_i) \\
& + \Gamma X_{i,t-1} + \varepsilon_{i,t}.
\end{aligned} \tag{3.1}$$

Here, $b_{i,t}^{CIP}$ is the signed three-month government-bond monthly averaged CIP deviation for country i (Du, Tepper, and Verdelhan, 2018), $\Delta q_{i,t}$ is monthly year-over-year real asset price growth (using BIS Residential Property Prices as the proxy), $VIXStress_t$ is an indicator equal to one when the VIX is above its sample 80th percentile, and $Safe_i$ is an indicator for Japan and Switzerland. The safe-haven main effect is absorbed by country fixed effects, but its interactions are identified in the all-country specifications. The coefficient of interest for non-safe-haven currencies is β_3 .

The baseline sample excludes safe-haven currencies, defined as Japan and Switzerland. This choice is consistent with the model, which is designed to capture the transmission from domestic housing weakness to external funding pressure in non-safe-haven small open economies. Japan and Switzerland are retained in the right-hand columns as an inclusive robustness check. In these specifications, safe-haven interactions are included so that the main housing-stress coefficient is identified separately from safe-haven currency dynamics.

Country fixed effects absorb persistent differences in currency markets, financial structure, and average basis levels. I report specifications with both calendar-year fixed effects

and year-quarter fixed effects. The control vector $X_{i,t-1}$ includes lagged monthly domestic industrial production growth, lagged monthly domestic CPI inflation, the lagged domestic 3-month maturity interest rate, the lagged US short-term interest rate, lagged US CPI inflation, and lagged US industrial production growth. These controls are included to separate the housing-CIP relation from ordinary domestic macroeconomic conditions and common US monetary or real activity shocks. The dependent variable is winsorized at the upper and lower one percent tails.

Table 3.1: State-Dependent Spillover: Housing Prices and Signed CIP Deviations

	Baseline: Excluding Safe Havens		Robustness: All Countries	
Dependent Var.: Signed CIP	(1) Year FE	(2) Quarter FE	(3) Year FE	(4) Quarter FE
Housing Growth ($\Delta q_{i,t}$)	0.380	0.455	0.526*	0.549*
	(0.342)	(0.395)	(0.271)	(0.283)
VIX Stress ($Stress_t$)	9.994**	8.726***	9.787**	7.587**
	(3.565)	(2.338)	(4.212)	(3.155)
Housing Growth $\times$ VIX Stress	-1.378**	-1.214*	-1.234**	-0.870
	(0.474)	(0.571)	(0.429)	(0.485)
Time fixed effects	Year	Year-quarter	Year	Year-quarter
Macro controls	O	O	O	O
Country fixed effects	O	O	O	O
Safe havens included	X	X	O	O
Safe-haven interactions	X	X	O	O
Observations	2,355	2,355	2,903	2,903
Adjusted R^2	0.473	0.520	0.463	0.523

Note: The dependent variable is the signed three-month government-bond CIP deviation. $Stress_t$ is a dummy equal to one when the VIX is in the top quintile of the sample distribution. Macro controls include lagged domestic industrial production growth, domestic CPI inflation, domestic interest rates, U.S. CPI inflation, U.S. industrial production growth, and U.S. interest rates. Safe-haven currencies are Japan and Switzerland. In Columns (3) and (4), safe-haven interactions include the interaction between housing growth, VIX stress, and the safe-haven dummy. Standard errors clustered by country are in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table 3.1 shows that the relationship between domestic housing markets and global FX markets is sharply state-dependent. In normal times, housing growth is not negatively associated with CIP deviations in the baseline sample. The coefficient on housing growth is positive and statistically insignificant in Columns (1) and (2). This finding establishes the

first motivating fact: a baseline disconnect exists, wherein local housing conditions do not map into global funding premia during normal periods.

This relationship fundamentally alters during periods of global stress. In Column (1), which incorporates country fixed effects and annual time fixed effects, the interaction between housing growth and VIX stress is negative and statistically significant at the 5 percent level. Specifically, a one percentage point decline in housing growth is associated with a 1.38 basis point increase in the signed CIP deviation during stress periods. Column (2) introduces a stricter specification by replacing annual time fixed effects with year-quarter fixed effects, thereby absorbing higher frequency global fluctuation and seasonality. The interaction coefficient remains stable, moving only slightly from -1.378 to -1.214 , and retains its statistical significance at the 10 percent level. This establishes the second core fact: a stress-time reconnect occurs, linking domestic asset weakness to wider global funding premia.

Columns (3) and (4) expand the sample to include safe-haven currencies (Japan and Switzerland), allowing the housing-stress relationship to differ for these specific nations. While the primary interaction remains negative, the specification with quarterly fixed effects becomes less precise.

Overall, this evidence establishes the quantitative targets that drive the theoretical framework. The linkage between domestic housing markets and the CIP basis is weak in normal times but materializes strongly when global risk is elevated. The results indicate that this reconnect is a structural phenomenon.

3.4 A Two-Period Model of Domestic Collateral and CIP Deviations

This section develops a two-period small open economy model that links domestic collateral values to CIP deviations. This toy model shows the intuitive mechanism behind the empirical pattern: local housing markets and global dollar funding markets interact closely only when intermediary balance sheet capacity is scarce.

Time is indexed by $t = 0, 1$. There is a Home country and a foreign dollar funding market. The Home economy contains a representative household and a global financial intermediary. The household borrows domestically to finance current expenditure. The intermediary supplies domestic credit and holds a hedged dollar liquidity position obtained through the FX swap market. The intermediary is risk neutral but subject to a balance sheet constraint. The key state variable is the intermediary's initial net worth N_0 .

3.4.1 Stage 1: Intra-temporal Optimization and the Price Index

In each period $t \in \{0, 1\}$, the household derives utility from a real consumption basket C_t , which is a Cobb-Douglas aggregate of non-tradable goods ($C_{NT,t}$), domestic tradable goods ($C_{H,t}$), and foreign tradable goods ($C_{F,t}$):

$$C_t = [(C_{NT,t})^{\chi_t} (C_{H,t})^{a_t} (C_{F,t})^{\iota_t}]^{\frac{1}{\theta_t}} \quad (3.2)$$

where χ_t , a_t , and ι_t are non-negative preference parameters representing the weights of each good, and $\theta_t \equiv \chi_t + a_t + \iota_t$.

Let Exp_t denote the household's total nominal expenditure in period t . For a given level of real consumption $C_t = 1$, the household minimizes the total cost of achieving this utility.

The expenditure minimization problem is formulated as:

$$\min_{C_{NT,t}, C_{H,t}, C_{F,t}} p_{NT,t}C_{NT,t} + p_{H,t}C_{H,t} + p_{F,t}C_{F,t} \quad (3.3)$$

subject to the constraint $[(C_{NT,t})^{\chi_t} (C_{H,t})^{a_t} (C_{F,t})^{\iota_t}]^{\frac{1}{\theta_t}} = 1$.

Solving this standard Lagrangian problem yields the optimal demand functions, which dictate that the household allocates a constant fraction of its total nominal expenditure Exp_t to each good:

$$p_{NT,t}C_{NT,t} = \frac{\chi_t}{\theta_t} Exp_t \quad (3.4)$$

$$p_{H,t}C_{H,t} = \frac{a_t}{\theta_t} Exp_t \quad (3.5)$$

$$p_{F,t}C_{F,t} = \frac{\iota_t}{\theta_t} Exp_t \quad (3.6)$$

Substituting these optimal demand functions back into the definition of the real consumption basket C_t , we can define the ideal Consumer Price Index (CPI), denoted by P_t , as the minimum cost of purchasing one unit of the consumption basket:

$$P_t \equiv \theta_t \left[\left(\frac{p_{NT,t}}{\chi_t} \right)^{\chi_t} \left(\frac{p_{H,t}}{a_t} \right)^{a_t} \left(\frac{p_{F,t}}{\iota_t} \right)^{\iota_t} \right]^{\frac{1}{\theta_t}} \quad (3.7)$$

This allows us to establish the fundamental accounting identity linking the real consumption index to nominal expenditure:

$$Exp_t = P_t C_t \implies C_t = \frac{Exp_t}{P_t} \quad (3.8)$$

3.4.2 Stage 2: Inter-temporal Optimization and the Nominal Euler Equation

Taking the intra-temporal price index P_t as given, the household solves the inter-temporal consumption smoothing problem. The lifetime expected utility is logarithmic:

$$U = \ln C_0 + \beta \mathbb{E}_0[\ln C_1] \quad (3.9)$$

Substituting Equation (3.8) into the utility function, we obtain the indirect utility function expressed purely in terms of nominal expenditure and the price index:

$$U = (\ln Exp_0 - \ln P_0) + \beta \mathbb{E}_0 [\ln Exp_1 - \ln P_1] \quad (3.10)$$

At $t = 0$, the household supplements its exogenous nominal income Y_0 by borrowing a nominal amount $B_{H,0}$ from the domestic financial intermediary. The loan carries a pre-determined nominal gross interest rate $R_{B,0}$ and is subject to an aggregate macroeconomic default shock $\delta_1 \in (0, 1)$ realized at $t = 1$. The dynamic nominal budget constraints are:

$$Exp_0 = Y_0 + B_{H,0} \quad (3.11)$$

$$Exp_1 = Y_1 - (1 - \delta_1)R_{B,0}B_{H,0} \quad (3.12)$$

The household maximizes the indirect utility function (Equation 3.10) with respect to $B_{H,0}$, subject to the budget constraints. Crucially, because the household is a price-taker, the price indices (P_0 and P_1) drop out entirely from the first-order derivative. The optimization yields the nominal Euler equation:

$$\frac{1}{Exp_0} = \beta \mathbb{E}_0 \left[\frac{(1 - \delta_1)R_{B,0}}{Exp_1} \right] \quad (3.13)$$

3.4.3 Macroeconomic Implication: Income-Driven Import Compression

This two-stage budgeting formulation prevents the dimension mismatch often found in open-economy models that mix real utility optimization with nominal financial contracts. More importantly, it microfound the *income-driven import compression* mechanism.

Recalling the optimal import demand from Equation (3.6) and the budget constraint from Equation (3.11), the domestic-currency value of imports is explicitly tied to the household's borrowing capacity:

$$p_{F,0}C_{F,0} = \frac{\iota_0}{\theta_0}Exp_0 = \frac{\iota_0}{\theta_0}(Y_0 + B_{H,0}) \quad (3.14)$$

Equation (3.14) establishes a direct partial-equilibrium link between domestic credit conditions and international trade flows. Because the household allocates a constant share of its expenditure to foreign goods, any contraction in borrowing ($B_{H,0} \downarrow$) reduces total nominal expenditure ($Exp_0 \downarrow$), proportionally lowering import demand. As shown later in Section 3.4.6, when intermediary balance sheet constraints restrict credit supply, the resulting domestic deleveraging endogenously compresses imports. This income-driven import compression operates alongside the standard expenditure-switching effect, linking domestic financial frictions to the balance of payments

3.4.4 Hedged Dollar Liquidity and the CIP Basis

Before defining the intermediary's portfolio problem, we formally define the covered interest parity (CIP) deviation in the foreign exchange swap market. Let S_0 denote the spot exchange rate, measured as units of Home currency per dollar, and let F_0 denote the one-period forward exchange rate. The foreign dollar gross interest rate is R_0^* , and the Home risk-free gross rate is R_D .

A hedged dollar position delivers a Home-currency gross return:

$$R_{X,0} = R_0^* \frac{F_0}{S_0}. \quad (3.15)$$

Covered interest parity holds when $R_{X,0} = R_D$. I define the exact CIP basis as the difference between the domestic safe rate and the covered foreign return:

$$b_0 \equiv R_D - R_0^* \frac{F_0}{S_0} = R_D - R_{X,0}. \quad (3.16)$$

With this sign convention, a larger b_0 corresponds to a larger dollar funding premium, or equivalently, a larger convenience yield on hedged dollar liquidity.

3.4.5 The Financial Intermediary

The intermediary enters period 0 with net worth N_0 . It chooses domestic loans $A_{H,0}$, hedged dollar liquidity X_0 , and risk-free debt D_0 . Its balance sheet is

$$A_{H,0} + X_0 = N_0 + D_0. \quad (3.17)$$

Debt is funded at the Home risk-free rate R_D . Terminal payoff is

$$\Pi_1 = A_{H,0} R_{H,1} + X_0 R_{X,0} - D_0 R_D. \quad (3.18)$$

Where $R_{H,1} \equiv (1 - \delta_1) R_{B,0}$ denotes the realized return on the domestic loan portfolio. Substituting the balance sheet into the payoff gives the expected profit objective

$$\mathbb{E}_0[\Pi_1] = A_{H,0} (\mathbb{E}_0[R_{H,1}] - R_D) + X_0 (R_{X,0} - R_D) + N_0 R_D. \quad (3.19)$$

The intermediary is subject to a Value-at-Risk constraint. To define the risk contribution

of each position, the model distinguishes the contractual return on the hedged dollar asset from the state-contingent liquidity service of holding spot dollar assets. The contractual return on X_0 is strictly fixed by the FX swap in Equation (3.15), which immunizes the intermediary from directional currency risk at maturity. However, the intermediary still faces severe interim liquidity strains. Under modern regulatory frameworks such as the Uncleared Margin Rules (UMR), a sharp depreciation of the Home currency triggers immediate variation margin (VM) cash outflows on the intermediary's short forward book. Concurrently, domestic corporate hedgers and importers face sudden dollar funding shortages.

To capture this institutional reality in a tractable reduced-form, let ℓ_1 denote the state-contingent liquidity service of the spot dollar position. This term represents the shadow value of holding spot dollar liquidity (e.g., US Treasuries) that can be pledged or liquidated to buffer these margin and funding drains. A natural proxy assumes that this shadow value is proportional to the gross rate of Home-currency depreciation, which governs the intensity of the funding pressure:

$$\ell_1 = \varphi R_0^* \frac{S_1}{S_0}, \quad \varphi > 0, \quad (3.20)$$

where φ is a positive sensitivity parameter. Therefore, the liquidity service ℓ_1 explicitly enters the intermediary's risk management problem as a valuable liquidity shield rather than a private contractual payoff.

The intermediary's risk exposure is summarized by

$$\mathcal{R}(A_{H,0}, X_0) = \kappa \sqrt{\text{Var}_0(A_{H,0}R_{H,1} + X_0\ell_1)}, \quad (3.21)$$

where $\kappa > 0$ is the required risk coverage multiple. The VaR constraint is

$$N_0 \geq \mathcal{R}(A_{H,0}, X_0). \quad (3.22)$$

Since $\mathcal{R}(\cdot)$ is homogeneous of degree one in portfolio positions, Euler's theorem implies

that it can be locally represented around a given stress portfolio $\mathcal{S}_0$ by its marginal risk contributions:

$$N_0 \geq \theta_H A_{H,0} + \theta_X X_0. \quad (3.23)$$

The coefficients θ_H and θ_X are the local derivatives of the VaR requirement:

$$\theta_H \equiv \left. \frac{\partial \mathcal{R}(A_{H,0}, X_0)}{\partial A_{H,0}} \right|_{\mathcal{S}_0} = \frac{\kappa}{\sqrt{\text{Var}_0}} (A_{H,0} \sigma_H^2 + X_0 \text{Cov}_0(R_{H,1}, \ell_1)), \quad (3.24)$$

$$\theta_X \equiv \left. \frac{\partial \mathcal{R}(A_{H,0}, X_0)}{\partial X_0} \right|_{\mathcal{S}_0} = \frac{\kappa}{\sqrt{\text{Var}_0}} (X_0 \sigma_\ell^2 + A_{H,0} \text{Cov}_0(R_{H,1}, \ell_1)), \quad (3.25)$$

where $\sigma_H^2 = \text{Var}_0(R_{H,1})$, $\sigma_\ell^2 = \text{Var}_0(\ell_1)$, and Var_0 denotes the variance term inside Equation (3.21).

Assumption 1 (State-dependent dollar liquidity hedge). *During financial stress, domestic credit losses structurally coincide with severe dollar funding pressures. While this negative relationship is taken as given for the intermediary's partial equilibrium problem, in Section 3.4.6, a domestic credit contraction compresses imports and forces the spot exchange rate to depreciate to clear the balance of payments. This depreciation mechanically activates the variation margin drains and corporate dollar shortages, sharply raising the shadow value of spot dollar liquidity ($\ell_1 \uparrow$). Formally, the covariance between domestic credit returns and the liquidity service of the spot dollar position is strictly negative:*

$$\text{Cov}_0(R_{H,1}, \ell_1) < 0. \quad (3.26)$$

The stress portfolio also satisfies the boundary condition:

$$\frac{\sigma_\ell^2}{|\text{Cov}_0(R_{H,1}, \ell_1)|} < \frac{A_{H,0}}{X_0} < \frac{\sigma_H^2}{|\text{Cov}_0(R_{H,1}, \ell_1)|}. \quad (3.27)$$

Under Assumption 1, Equations (3.24)–(3.25) imply

$$\theta_H > 0, \quad \theta_X < 0. \quad (3.28)$$

The first inequality states that domestic credit exposure consumes risk capacity. A negative θ_X implies that spot dollar liquidity provides a conditional macro-natural hedge, locally relaxing the intermediary's VaR constraint in crisis states. Because the variation margin and funding pressures (and thus the shadow value of liquidity, ℓ_1) spike precisely when domestic collateral defaults and the currency depreciates, holding X_0 locally reduces the intermediary's required risk capacity around the stress portfolio.

Let $\mu_0 \geq 0$ be the multiplier on Equation (3.23). The intermediary's Lagrangian is

$$\mathcal{L} = \mathbb{E}_0[\Pi_1] + \mu_0 (N_0 - \theta_H A_{H,0} - \theta_X X_0). \quad (3.29)$$

The first-order conditions are

$$\mathbb{E}_0[R_{H,1}] - R_D = \mu_0 \theta_H, \quad (3.30)$$

$$R_0^* \frac{F_0}{S_0} - R_D = \mu_0 \theta_X. \quad (3.31)$$

Using the definition of the basis in Equation (3.16), the second condition can be written as

$$b_0 = -\mu_0 \theta_X = \mu_0 |\theta_X|. \quad (3.32)$$

Equation (3.32) is the central pricing equation of the model. The CIP basis is zero when the balance sheet constraint is slack and positive when the constraint binds. In the binding regime, the basis explicitly measures the shadow value of spot dollar liquidity required to buffer regulatory and funding pressures.

3.4.6 Equilibrium

The spot dollar market clears through the balance of payments. Let $Z_0 S_0$ denote the Home-currency value of export revenue, where $Z_0 > 0$ is an exogenous export volume. The term $m(Y_0 + A_{H,0})$ represents the Home-currency value of import demand, where Let $m \equiv \iota_0/\theta_0$ represent the constant import expenditure share derived in Equation (3.14). Let Q_0 denote non-fundamental dollar hedging demand, measured in Home currency. This demand comes from institutional hedgers and liquidity traders, such as corporate hedgers, life insurers, or passive global investors, whose currency-hedging needs are only imperfectly responsive to current prices. The intermediary absorbs the residual demand for hedged dollar liquidity:

$$X_0 = Z_0 S_0 - m(Y_0 + A_{H,0}) - Q_0. \quad (3.33)$$

This condition plays two roles. First, it links the real side of the open economy to dollar liquidity demand: a contraction in domestic credit lowers expenditure and compresses imports. Second, it allows non-fundamental dollar demand Q_0 to affect the exchange rate only when the intermediary cannot elastically absorb it.

Definition 1 (Equilibrium). *Given the exogenous states and parameters $\{Y_0, Y_1, N_0, R_D, R_0^*, Z_0, Q_0, \delta_1\}$, a competitive equilibrium consists of a set of prices $\{R_{B,0}, S_0, F_0, b_0\}$, allocations $\{B_{H,0}, A_{H,0}, X_0, D_0, Exp_0, Exp_1\}$, and a multiplier μ_0 such that:*

1. **Household optimization:** the Euler equation (3.13) and budget constraints (3.11)–(3.12) hold.
2. **Intermediary optimization:** the first-order conditions (3.30)–(3.31) and the balance sheet identity (3.17) hold.
3. **Domestic credit market clearing:** $A_{H,0} = B_{H,0}$.
4. **Foreign exchange market clearing:** $X_0 = Z_0 S_0 - m(Y_0 + A_{H,0}) - Q_0$.
5. **VaR constraint and complementary slackness:** the linearized risk constraint

satisfies primal feasibility and complementary slackness:

$$N_0 - \theta_H A_{H,0} - \theta_X X_0 \geq 0, \quad \mu_0 \geq 0, \quad \text{and} \quad \mu_0 (N_0 - \theta_H A_{H,0} - \theta_X X_0) = 0.$$

3.4.7 Regimes and Main Implications

Let $(A_{H,0}^*, X_0^*)$ denote the optimal allocation that would prevail if the intermediary were unconstrained. The minimum net worth required to support this first-best allocation defines the critical threshold $\underline{N}$:

$$\underline{N} \equiv \theta_H A_{H,0}^* + \theta_X X_0^* = \theta_H A_{H,0}^* - |\theta_X| X_0^*. \quad (3.34)$$

If $N_0 \geq \underline{N}$, the balance sheet constraint is slack. If $N_0 < \underline{N}$, the constraint binds.

Proposition 1 (Normal-time disconnect). *If $N_0 \geq \underline{N}$, then $\mu_0 = 0$ and the CIP basis is zero:*

$$b_0 = 0. \quad (3.35)$$

In this regime, marginal changes in domestic collateral risk do not affect the CIP basis.

Proof. When the constraint is slack ($N_0 \geq \underline{N}$), complementary slackness implies $\mu_0 = 0$. Equation (3.32) then gives $b_0 = 0$. Since the basis depends on domestic collateral risk only through μ_0 , local collateral shocks have no first-order effect on b_0 as long as the constraint remains slack. $\square$

Proposition 2 (Crisis-state reconnect). *If $N_0 < \underline{N}$, the balance sheet constraint binds, $\mu_0 > 0$, and the CIP basis is positive:*

$$b_0 = \mu_0 |\theta_X| > 0. \quad (3.36)$$

Moreover, any shock that raises the shadow value of intermediary balance sheet capacity

increases the dollar funding premium.

Proof. When $N_0 < \underline{N}$, the unconstrained allocation violates Equation (3.23), so the constraint binds and $\mu_0 > 0$. Under Assumption 1, $\theta_X < 0$. Equation (3.32) therefore implies $b_0 = \mu_0|\theta_X| > 0$. The comparative static with respect to μ_0 follows immediately:

$$\frac{\partial b_0}{\partial \mu_0} = |\theta_X| > 0. \quad (3.37)$$

□

This is the crisis-state reconnect. A domestic housing decline raises expected domestic credit losses, increasing the required compensation for holding domestic loans through Equation (3.30). When the intermediary is already constrained ($N_0 < \underline{N}$), this deterioration raises the shadow cost μ_0 . Because μ_0 also prices hedged dollar liquidity through Equation (3.31), the same domestic collateral shock widens the CIP basis.

The model also clarifies why the empirical relationship is nonlinear. δ_1 represent the aggregate default rate on domestic credit. Because domestic loans are predominantly collateralized by real estate, an exogenous decline in local housing prices directly manifests as a higher δ_1 .

Crucially, housing price declines do not move CIP deviations unconditionally. They widen the basis only when they affect the marginal value of intermediary balance sheet capacity. This yields the following threshold prediction:

$$\frac{\partial b_0}{\partial \mathbb{E}_0[\delta_1]} = \begin{cases} 0, & N_0 \geq \underline{N}, \\ |\theta_X| \frac{\partial \mu_0}{\partial \mathbb{E}_0[\delta_1]} > 0, & N_0 < \underline{N}. \end{cases} \quad (3.38)$$

This state-dependent derivative is the structural counterpart to the empirical specification in Section 3.3, which tests the negative interaction between local housing price growth and global financial stress.

Proposition 3 (Hedging-demand amplification and import compression). *In the binding regime, exogenous dollar demand Q_0 directly depreciates the spot exchange rate through the constrained dollar market. Concurrently, domestic credit contraction partially offsets this depreciation pressure via import compression and the freeing of intermediary risk capacity.*

Proof. When the constraint binds, Equation (3.23) pins down the required dollar liquidity position as:

$$X_0 = \frac{\theta_H A_{H,0} - N_0}{|\theta_X|}. \quad (3.39)$$

Substituting Equation (3.39) into the dollar market clearing condition (3.33) and rearranging for S_0 yields the equilibrium mapping for the spot exchange rate:

$$S_0 = \frac{1}{Z_0} \left[m(Y_0 + A_{H,0}) + Q_0 + \frac{\theta_H A_{H,0} - N_0}{|\theta_X|} \right]. \quad (3.40)$$

Taking partial derivatives with respect to Q_0 and $A_{H,0}$ gives:

$$\frac{\partial S_0}{\partial Q_0} = \frac{1}{Z_0} > 0, \quad \frac{\partial S_0}{\partial A_{H,0}} = \frac{1}{Z_0} \left(m + \frac{\theta_H}{|\theta_X|} \right) > 0. \quad (3.41)$$

Since a domestic credit contraction ($A_{H,0} \downarrow$) mechanically lowers the right-hand side, it mitigates depreciation pressure ($S_0 \downarrow$). $\square$

Proposition 3 highlights two distinct forces operating during a financial crisis. First, because the constrained intermediary cannot elastically absorb exogenous hedging demand, non-fundamental dollar demand shocks (Q_0) pass directly into the spot exchange rate. Second, domestic deleveraging provides a stabilizing general equilibrium feedback through two channels: it reduces import demand (m), and it frees intermediary risk capacity by shrinking domestic credit exposure ($\frac{\theta_H}{|\theta_X|}$). Thus, the real economy acts as a macroeconomic shock absorber: the costly domestic credit contraction buffers the exchange rate against severe depreciation.

Finally, it is straightforward to verify that the negative covariance between domestic credit returns and the dollar liquidity service, taken as given in the intermediary’s partial equilibrium problem (Assumption 1), holds endogenously in the general equilibrium. A formal proposition and proof can be provided as an appendix extension of the macro-natural hedge mechanism.

3.5 A Full DSGE Model with Endogenous CIP Deviations

The two-period model isolates the core mechanism: domestic collateral stress affects dollar funding premia only when the intermediary’s balance sheet constraint binds. This section embeds this mechanism into a dynamic, infinite-horizon small open economy framework with housing, endogenous default risk, nominal rigidities, and an explicit FX swap market. CIP deviations are directly observable in FX swap markets and constitute a standard measure of limits to arbitrage in global bank balance sheets (Du, Tepper, and Verdelhan, 2018).

The full model preserves two central ingredients from the two period model. First, domestic mortgage exposure consumes intermediary risk capacity. Second, hedged dollar liquidity provides a macro-natural hedge that relaxes the VaR constraint during stress states. Under our maintained assumptions, the marginal risk contributions satisfy

$$\theta_{H,t} > 0, \quad \theta_{X,t} < 0. \quad (3.42)$$

This asymmetry remains the key structural channel through which domestic housing stress translates into a wider CIP basis.

Time is discrete and indexed by $t = 0, 1, 2, \dots$. The home economy is populated by patient savers, impatient borrowers, intermediate and final goods firms, and a representative financial intermediary. The population size is normalized to unity, with patient households

comprising a share γ and impatient households comprising $1 - \gamma$. Patient households supply deposits to the intermediary. Impatient households borrow against housing collateral. The intermediary supplies domestic mortgage credit and channels hedged dollar funding through the FX swap market. The foreign nominal risk-free rate R_t^* is taken as strictly exogenous. To streamline the exposition, the main text focuses on the core economic mechanisms and the occasionally binding financial friction. A complete derivation of the model, including all first-order conditions and the full system of non-linear equilibrium equations, is provided in Appendix C.2.

3.5.1 Households, Housing, and Domestic Credit

Patient households maximize

$$\mathbb{E}_0 \sum_{t=0}^{\infty} \beta_S^t \left[\log C_{S,t} + \xi_{H,t} \log H_{S,t} - \frac{N_{S,t}^{1+\eta}}{1+\eta} \right], \quad (3.43)$$

subject to the real budget constraint

$$C_{S,t} + d_t + q_t(H_{S,t} - H_{S,t-1}) + \frac{\kappa_H}{2}(H_{S,t} - H_{S,t-1})^2 = w_t N_{S,t} + \frac{R_{D,t-1}}{\pi_t} d_{t-1} + \frac{Div_t}{\gamma} + \Pi_t^F. \quad (3.44)$$

Here d_t denotes real deposits, $R_{D,t}$ is the gross nominal domestic deposit rate, π_t is gross CPI inflation, q_t is the real housing price, and Div_t denotes aggregate dividends from the intermediary distributed proportionally to patient households.

Impatient households maximize an analogous utility function with a lower discount factor ($\beta_B < \beta_S$). They borrow from the intermediary subject to a standard loan-to-value (LTV) collateral constraint:

$$b_t \leq m_L \mathbb{E}_t \left[\frac{q_{t+1} \pi_{t+1}}{R_{B,t}} H_{B,t} \right], \quad (3.45)$$

where b_t denotes real borrowing, $R_{B,t}$ is the gross nominal mortgage rate, and m_L is the regulatory LTV limit. The intermediary holds the aggregate sum of these loans as domestic

mortgage assets, mapping household borrowing to the intermediary balance sheet:

$$A_{H,t} = (1 - \gamma)b_t. \quad (3.46)$$

Mortgage loans are subject to aggregate default risk. The realized real return on domestic mortgage assets is

$$R_{H,t} = (1 - \delta_t) \frac{R_{B,t-1}}{\pi_t}, \quad (3.47)$$

where the default rate δ_t rises when housing collateral values fall below their steady-state real housing price, $\bar{q}$:

$$\delta_t = \frac{\bar{\delta}}{\bar{\delta} + (1 - \bar{\delta}) \exp \left[\kappa_d \left(\frac{q_t}{\bar{q}} - 1 \right) \right]} + Shock_t^{NPL}. \quad (3.48)$$

where $Shock_t^{NPL}$ represents an exogenous deterioration in aggregate credit quality. This logistic formulation establish the macroeconomic link between local housing market declines and domestic credit risk.

3.5.2 Consumption Aggregation and Prices

Aggregate consumption C_t is a CES composite of Home ($C_{H,t}$) and Foreign ($C_{F,t}$) goods:

$$C_t = \left[(1 - \alpha)^{\frac{1}{\tau}} C_{H,t}^{\frac{\tau-1}{\tau}} + \alpha^{\frac{1}{\tau}} C_{F,t}^{\frac{\tau-1}{\tau}} \right]^{\frac{\tau}{\tau-1}}, \quad (3.49)$$

where $\alpha \in (0, 1)$ is the degree of openness (the import share) and $\tau > 0$ is the elasticity of substitution between Home and Foreign goods.

The Consumer Price Index (CPI) is defined as:

$$P_t = \left[(1 - \alpha) P_{H,t}^{1-\tau} + \alpha P_{F,t}^{1-\tau} \right]^{\frac{1}{1-\tau}}. \quad (3.50)$$

Assuming the Law of One Price holds at the dock, the domestic price of foreign goods is $P_{F,t} = S_t P_t^*$, where S_t is the nominal exchange rate and P_t^* is the exogenous foreign price level. Standard CES algebra yields the following demand functions:

$$C_{H,t} = (1 - \alpha) \left(\frac{P_{H,t}}{P_t} \right)^{-\tau} C_t, \quad C_{F,t} = \alpha \left(\frac{P_{F,t}}{P_t} \right)^{-\tau} C_t. \quad (3.51)$$

Importantly, the foreign goods demand equation captures the import compression channel: a deleveraging-induced contraction in aggregate consumption C_t reduces the real demand for imported goods $C_{F,t}$.

3.5.3 Firms and Price Setting

To allow financial shocks to have persistent real effects, I introduce nominal rigidities. Firms operate in a monopolistically competitive environment subject to Rotemberg pricing adjustment costs. They produce differentiated domestic goods using a linear technology $Y_t = A_t N_t$, where A_t is exogenous productivity and N_t is aggregate labor.

The real marginal cost is $mc_t = \frac{w_t/A_t}{p_{H,t}}$, where $p_{H,t} \equiv P_{H,t}/P_t$ is the relative price of domestic goods. Since patient households own the firms, profits are discounted using their stochastic discount factor $\Lambda_{S,t,t+1} \equiv \beta_S \frac{C_{S,t}}{C_{S,t+1}}$. Profit maximization yields the non-linear New Keynesian Phillips Curve (NKPC) for domestic inflation $\pi_{H,t}$:

$$(\pi_{H,t} - \bar{\pi})\pi_{H,t} = \mathbb{E}_t \left[\Lambda_{S,t,t+1} \frac{Y_{t+1}}{Y_t} (\pi_{H,t+1} - \bar{\pi})\pi_{H,t+1} \right] + \frac{\varepsilon}{\phi_{adj}} \left(mc_t - \frac{\varepsilon - 1}{\varepsilon} \right), \quad (3.52)$$

where ε is the elasticity of substitution among domestic varieties, and ϕ_{adj} governs the degree of price stickiness. The relative price of domestic goods evolves according to $p_{H,t} = p_{H,t-1} \frac{\pi_{H,t}}{\pi_t}$.

Finally, aggregate real dividends distributed to patient households are:

$$\Pi_t^F = p_{H,t} Y_t - w_t N_t - \frac{\phi_{adj}}{2} (\pi_{H,t} - \bar{\pi})^2 Y_t. \quad (3.53)$$

3.5.4 Covered Dollar Funding and the CIP Basis

The intermediary also supplies hedged dollar funding through the FX swap market. Let S_t denote the spot nominal exchange rate and F_t the one-period forward rate. Define the gross forward premium as:

$$f_t \equiv \frac{F_t}{S_t}. \quad (3.54)$$

The intermediary buys the foreign safe asset in the spot market and locks in the domestic-currency payoff through the forward contract. Hence, the contractual real return on the hedged dollar position chosen at t and paid at $t + 1$ is:

$$R_{FX,t+1} = \frac{R_t^* f_t}{\pi_{t+1}}. \quad (3.55)$$

The forward premium f_t is endogenous. It is the equilibrium price that clears the covered funding market when intermediaries face balance sheet constraints.

The exact nominal CIP basis is defined as:

$$b_t^{CIP} = R_{D,t} - R_t^* f_t. \quad (3.56)$$

A positive value of b_t^{CIP} indicates that the domestic nominal safe rate exceeds the covered foreign nominal return. Crucially, in this model, the basis is not imposed as an exogenous wedge. It is determined endogenously by the intermediary's optimality condition for supplying hedged dollar funding.

3.5.5 Intermediary Balance Sheet and Net Worth

The intermediary holds domestic mortgage assets and hedged dollar assets. Its total assets are

$$A_{Total,t} = A_{H,t} + rer_t A_{FX,t}, \quad (3.57)$$

where rer_t is the real exchange rate and $A_{FX,t}$ is the foreign asset position measured in foreign goods. The realized real portfolio return is

$$R_{port,t} = \frac{A_{H,t-1}R_{H,t} + rer_{t-1}A_{FX,t-1}R_{FX,t}}{A_{Total,t-1}}. \quad (3.58)$$

Aggregate net worth evolves according to the law of motion:

$$N_{FI,t} = \sigma_{bank}ROE_t N_{FI,t-1} + W_{inj}, \quad (3.59)$$

where σ_{bank} is the constant fraction of internal earnings retained by the infinitely-lived intermediary, W_{inj} is an exogenous lump-sum equity injection from patient households to prevent the banking sector from shrinking to a zero-net-worth state, and the return on equity (ROE_t) is

$$ROE_t = \left(R_{port,t} - \frac{R_{D,t-1}}{\pi_t} - \frac{\psi}{2}A_{Total,t-1} \right) \phi_{t-1} + \frac{R_{D,t-1}}{\pi_t}. \quad (3.60)$$

Here, $\phi_t = A_{Total,t}/N_{FI,t}$ is the intermediary's leverage ratio. The quadratic total cost function $\frac{\psi}{2}A_{Total,t-1}^2$ translates into a linear penalty $\frac{\psi}{2}A_{Total,t-1}$ on the return on equity, capturing ordinary balance sheet management costs that scale with the size of intermediation.

Following Gertler and Karadi (2011), let ν_t denote the expected marginal value of a unit of net worth at the end of period t . This continuation value satisfies the recursive equation:

$$\nu_t = \mathbb{E}_t [\Lambda_{S,t,t+1}\Omega_{t+1}ROE_{t+1}], \quad \Omega_{t+1} = 1 - \sigma_{bank} + \sigma_{bank}\nu_{t+1}, \quad (3.61)$$

where $\Lambda_{S,t,t+1}$ is the patient household's stochastic discount factor and Ω_{t+1} represents the shadow marginal value of a unit of internal funds.

3.5.6 Dollar Liquidity and the VaR Constraint

The intermediary's risk constraint is written in terms of local marginal risk contributions.

Let

$$w_{H,t} = \frac{A_{H,t}}{A_{Total,t}}, \quad w_{X,t} = \frac{rer_t A_{FX,t}}{A_{Total,t}}. \quad (3.62)$$

The model distinguishes the contractual return on the hedged dollar asset from the state-contingent liquidity service of holding spot dollar assets. The contractual return is strictly fixed by the swap contract in Equation (3.55), which immunizes the intermediary from directional currency risk at maturity. However, the intermediary still faces severe interim liquidity strains. Under modern regulatory frameworks such as the Uncleared Margin Rules (UMR), a sharp depreciation of the domestic currency triggers immediate variation margin (VM) cash outflows on the intermediary's short forward book. Concurrently, domestic importers and corporates face sudden dollar funding shortages.

To capture this institutional reality in a tractable manner, the state-contingent liquidity service of spot dollar assets is modeled in a reduced-form as:

$$\ell_t = \phi_\ell \frac{R_{U,t}}{R^*}, \quad R_{U,t} = \frac{rer_t}{rer_{t-1}} R_{t-1}^*. \quad (3.63)$$

This term proxies the shadow value of holding spot dollar liquidity (e.g., US Treasuries) that can be immediately pledged or liquidated to buffer these state-contingent funding outflows. It explicitly enters the intermediary's risk management problem as a valuable liquidity service, rather than a private contractual payoff.

The perceived portfolio risk is

$$\sigma_{p,t}^2 = w_{H,t}^2 \sigma_{H,t}^2 + w_{X,t}^2 \sigma_\ell^2 + 2w_{H,t}w_{X,t} \text{Cov}_t(R_{H,t+1}, \ell_{t+1}). \quad (3.64)$$

A defining feature of this macroeconomic environment is that domestic credit returns and

dollar liquidity services are negatively correlated in stress states:

$$\text{Cov}_t(R_{H,t+1}, \ell_{t+1}) < 0. \quad (3.65)$$

Consistent with the previous two-period setup, this negative covariance emerges endogenously in the general equilibrium. When local housing values fall, domestic credit returns decline because default risk rises. Concurrently, the resulting capital outflows force the home currency to depreciate. This depreciation mechanically activates the variation margin drains on the forward book and elevates the corporate sector's funding demands, sharply increasing the shadow value of dollar liquidity ($\ell_t \uparrow$). Because domestic credit losses and dollar liquidity pressures reliably coincide during crises, this endogenous negative covariance gives spot dollar assets a macro-natural hedge role on the intermediary's balance sheet. This reduced-form mechanism aligns with empirical evidence that dollar safe assets provide convenience yields that locally relax effective financing constraints (Jiang, Krishnamurthy, and Lustig, 2021).

However, for the quantitative dynamic simulation using the OccBin algorithm, which relies on piecewise linear approximations, computing time-varying conditional covariances is computationally infeasible. Following standard macro-finance approximations, I evaluate the endogenous covariance $\text{Cov}_t(R_{H,t+1}, \ell_{t+1})$ at its steady-state implied value and treat it as a constant parameter ($\text{cov_H_ell} < 0$) in the local VaR loadings.

The local VaR loadings are

$$\theta_{H,t} = \frac{\kappa_{VaR}}{\sigma_{p,t}} \left(w_{H,t} \sigma_{H,t}^2 + w_{X,t} \text{Cov}_t(R_{H,t+1}, \ell_{t+1}) \right), \quad (3.66)$$

$$\theta_{X,t} = \frac{\kappa_{VaR}}{\sigma_{p,t}} \left(w_{X,t} \sigma_{\ell}^2 + w_{H,t} \text{Cov}_t(R_{H,t+1}, \ell_{t+1}) \right). \quad (3.67)$$

As verified in the quantitative analysis, this macro-natural hedge ensures that domestic

mortgage exposure tightens the constraint while hedged dollar liquidity locally relaxes it:

$$\theta_{H,t} > 0, \quad \theta_{X,t} < 0. \quad (3.68)$$

The loadings are local marginal VaR weights computed at the current portfolio shares. Conditional on these weights, the VaR boundary is linearized in the level exposures that enter the intermediary's balance sheet. The intermediary's balance sheet constraint is

$$\nu_t N_{FI,t} \geq \theta_{H,t} A_{H,t} + \theta_{X,t} rer_t A_{FX,t}. \quad (3.69)$$

Dividing by $N_{FI,t}$ gives the equivalent leverage form,

$$\nu_t \geq \phi_t (\theta_{H,t} w_{H,t} + \theta_{X,t} w_{X,t}), \quad \phi_t = \frac{A_{Total,t}}{N_{FI,t}}. \quad (3.70)$$

Thus the constraint restricts the scale of risk exposures relative to intermediary net worth. It preserves the marginal VaR interpretation of $\theta_{H,t}$ and $\theta_{X,t}$, while ensuring that depleted net worth tightens the feasible balance sheet. Let $\mu_t \geq 0$ denote the multiplier on this constraint. A higher μ_t means that intermediary risk capacity is scarce.

3.5.7 Asset Pricing Conditions

The intermediary chooses $A_{H,t}$ and $A_{FX,t}$ taking prices and the local VaR loadings as given. The relevant Lagrangian is:

$$\mathcal{L}_t = (1 + \mu_t) \nu_t N_{FI,t} + \Upsilon_t - \mu_t (\theta_{H,t} A_{H,t} + \theta_{X,t} rer_t A_{FX,t}). \quad (3.71)$$

Differentiating with respect to the level exposure in domestic mortgage assets yields the first-order condition:

$$\mathbb{E}_t \left[\Lambda_{t,t+1}^S \Omega_{t+1} \left(R_{H,t+1} - \frac{R_{D,t}}{\pi_{t+1}} - \psi A_{Total,t} \right) \right] = \frac{\mu_t}{1 + \mu_t} \theta_{H,t}. \quad (3.72)$$

Similarly, the first-order condition for the hedged dollar asset is:

$$\mathbb{E}_t \left[\Lambda_{t,t+1}^S \Omega_{t+1} \left(\frac{R_t^* f_t}{\pi_{t+1}} - \frac{R_{D,t}}{\pi_{t+1}} - \psi A_{Total,t} \right) \right] = \frac{\mu_t}{1 + \mu_t} \theta_{X,t}. \quad (3.73)$$

Since R_t^* , $R_{D,t}$, and f_t are known at date t , Equation (3.73) prices the forward premium directly. Factoring out the predetermined variables, it can be rewritten as:

$$(R_t^* f_t - R_{D,t}) \mathbb{E}_t \left[\frac{\Lambda_{t,t+1}^S \Omega_{t+1}}{\pi_{t+1}} \right] - \psi A_{Total,t} \mathbb{E}_t [\Lambda_{t,t+1}^S \Omega_{t+1}] = \frac{\mu_t}{1 + \mu_t} \theta_{X,t}.$$

Substituting the definition of the exact nominal basis, $b_t^{CIP} = R_{D,t} - R_t^* f_t$, gives the closed-form solution for the CIP deviation:

$$b_t^{CIP} = - \frac{\psi A_{Total,t} \mathbb{E}_t [\Lambda_{t,t+1}^S \Omega_{t+1}] + \frac{\mu_t}{1 + \mu_t} \theta_{X,t}}{\mathbb{E}_t \left[\frac{\Lambda_{t,t+1}^S \Omega_{t+1}}{\pi_{t+1}} \right]}. \quad (3.74)$$

This expression proves that a binding constraint ($\mu_t > 0$) endogenously widens the CIP basis whenever the hedged dollar position provides a macro-natural hedge ($\theta_{X,t} < 0$). The exact same shadow value of balance sheet capacity (μ_t) also raises the domestic credit spread through Equation (3.72). Thus, domestic credit markets and global dollar funding markets are tightly connected by the common constraint of the intermediary. This mechanism provides a structural DSGE foundation for the empirical co-movement between domestic credit spreads and CIP deviations documented by Liao (2020).

3.5.8 Foreign Exchange Market Clearing and the Balance of Payments

Exogenous hedgers and liquidity traders demand forward foreign currency through FX swaps, as in segmented FX market models with noise traders (Itskhoki and Mukhin, 2021). These agents represent institutional actors, such as corporate hedgers, life insurance companies, and passive global investors, whose currency-hedging or dollar-liquidity needs are mechanical and only imperfectly responsive to contemporaneous exchange rates or interest differentials. Thus, $Q_{N,t}$ denotes non-fundamental liquidity demand rather than retail or sentiment-driven trading. The exogenous component of forward demand follows an AR(1) process:

$$Q_t^{exog} = \bar{Q} + \rho_Q(Q_{t-1}^{exog} - \bar{Q}) + \varepsilon_{Q,t}. \quad (3.75)$$

Total forward demand is elastic with respect to the CIP basis:

$$Q_{N,t} = Q_t^{exog} + \kappa_{CIP} b_t^{CIP}, \quad \kappa_{CIP} > 0. \quad (3.76)$$

The representative intermediary supplies these swaps. To remain fully hedged against currency risk, it must hold the corresponding spot foreign asset position. Therefore, the forward market clearing condition pins down the intermediary's spot external position:

$$NFA_t = A_{FX,t} = Q_{N,t}. \quad (3.77)$$

These equations explicitly link forward-market intermediation to the spot external position. A wider CIP basis induces a larger forward position from these price-elastic hedgers, which in turn strictly requires the intermediary to acquire additional spot foreign assets today to hedge its swap book.

The country's balance of payments (BOP) identity is given by:

$$rer_t NFA_t = rer_t R_{t-1}^* NFA_{t-1} + NX_t. \quad (3.78)$$

Net exports, expressed in units of the domestic consumption basket, are:

$$NX_t = p_{H,t} EX_t - rer_t C_{F,t}, \quad (3.79)$$

where export demand and import demand are derived from standard CES aggregators:

$$EX_t = \alpha \left(\frac{p_{H,t}}{rer_t} \right)^{-\tau} Y^*, \quad C_{F,t} = \alpha rer_t^{-\tau} C_t. \quad (3.80)$$

The spot real exchange rate clears the resource requirement created by swap-market hedging. When inelastic forward hedging demand forces intermediaries to accumulate spot foreign assets ($NFA_t \uparrow$), the domestic economy must finance this position through the trade balance. A real depreciation ($rer_t \uparrow$) boosts exports, compresses imports, and drives net exports into surplus, thereby generating the required foreign resources.

3.5.9 Housing Stress and Risk Repricing

In practice, the balance sheet pressure facing financial institutions is driven as much by forward-looking risk assessments as by the immediate realization of default losses. Rather than relying solely on ex-post default volumes, the risk management protocols of global intermediaries are respond to the perceived volatility of their assets. Following the literature on risk and uncertainty shocks (Christiano, Motto, and Rostagno, 2014), I model the housing market contraction as a catalyst for the endogenous repricing of domestic credit risk.

This formulation more accurately reflects the empirical reality that intermediaries proactively compress their risk capacity in response to heightened macro-financial uncertainty. Specifically, the perceived variance of domestic mortgage returns, $\sigma_{H,t}^2$, is indexed linearly to

the state of NPL stress:

$$\sigma_{H,t}^2 = \bar{\sigma} \cdot (1 + \chi_{NPL} Shock_t^{NPL}), \quad (3.81)$$

where $Shock_t^{NPL}$ is an exogenous shock process that raises actual defaults and simultaneously triggers forward-looking risk repricing.

By endogenizing the risk assessment process, the model links collateral depreciation to a tightening of the VaR constraint. This ensures that the subsequent widening of the CIP basis is driven by the structural scarcity of risk-bearing capacity, consistent with the behavior of global banks during periods of acute financial stress.

3.5.10 Market Clearing and Equilibrium Definition

To close the model, all domestic markets must clear. The aggregate resource constraint for the domestic economy equates total domestic output to domestic consumption of home goods, exports, and resource costs generated by financial and nominal frictions:

$$Y_t = C_{H,t} + EX_t + \frac{\psi}{2} A_{Total,t-1}^2 + \frac{\phi_{adj}}{2} (\pi_{H,t} - \bar{\pi})^2 Y_t + \Xi_{H,t}, \quad (3.82)$$

where $\Xi_{H,t}$ represents the aggregate quadratic housing adjustment costs incurred by both patient and impatient households.

The domestic housing market clears with a fixed aggregate supply, normalized to unity:

$$\gamma H_{S,t} + (1 - \gamma) H_{B,t} = 1. \quad (3.83)$$

The domestic financial markets clear such that the intermediary's total deposit liabilities equal the aggregate savings of patient households ($d_t = \gamma d_{S,t}$), and total domestic mortgage assets equal the aggregate borrowing of impatient households ($A_{H,t} = (1 - \gamma) b_t$).

Finally, monetary policy closes the nominal block. The central bank sets the nominal

interest rate $R_{D,t}$ following a standard Taylor rule focused on domestic inflation $\pi_{H,t}$:

$$\log\left(\frac{R_{D,t}}{\bar{R}_D}\right) = \rho_R \log\left(\frac{R_{D,t-1}}{\bar{R}_D}\right) + (1 - \rho_R)\phi_\pi \log\left(\frac{\pi_{H,t}}{\bar{\pi}}\right) + \varepsilon_{R,t}. \quad (3.84)$$

where $\varepsilon_{R,t}$ stands for exogenous monetary policy shock.

Definition of Competitive Equilibrium. Given the exogenous stochastic processes for foreign demand (Y^*), the foreign interest rate (R_t^*), domestic productivity (A_t), and price-elastic forward hedging demand (Q_t^{exog}), a competitive equilibrium is a sequence of prices and aggregate quantities such that:

1. Patient and impatient households maximize their respective expected utilities subject to their budget and collateral constraints, yielding the consumption Euler equations, labor supply conditions, and the binding LTV constraint.
2. Intermediate goods firms maximize profits subject to Rotemberg pricing adjustment costs, yielding the New Keynesian Phillips Curve (NKPC) for domestic inflation.
3. The representative financial intermediary maximizes its franchise value subject to the occasionally binding VaR constraint (3.69), yielding the asset pricing conditions for domestic credit (3.72) and hedged dollar funding (3.73).
4. The global FX swap market clears (3.77), determining the endogenous forward premium and the exact CIP basis (3.74).
5. The balance of payments identity holds (3.78), determining the spot real exchange rate through trade flows (3.79) and (3.80).
6. All domestic markets clear, satisfying the aggregate resource constraint (3.82), the housing market clearing (3.83), and domestic credit market clearing, while the central bank follows the monetary policy rule (3.84).

3.6 Quantitative Exercise: Dynamic Mechanism

The full DSGE model is used as a dynamic mechanism exercise rather than as an estimated quantitative model. The purpose is to ask whether the static logic developed above survives once the economy contains forward-looking households, housing collateral, nominal rigidities, external adjustment, and an occasionally binding intermediary constraint. The exercise therefore focuses on the direction and sequencing of the responses. A domestic mortgage-risk or uncertainty shock should move the intermediary from the slack region into the constrained region; once the constraint binds, the shadow value of balance sheet capacity should connect domestic mortgage risk to the covered dollar funding market.

3.6.1 Implementation

The Dynare implementation follows the equations in the previous section and solves the model with the piecewise-linear method of Guerrieri and Iacoviello (2015). The state variable that governs regime switching is the intermediary's continuation value relative to the VaR boundary,

$$\nu_t \geq \phi_t(\theta_{H,t}w_{H,t} + \theta_{X,t}w_{X,t}).$$

The constraint binds when ν_t falls below this boundary. In the slack regime, the multiplier satisfies $\mu_t = 0$, so the intermediary prices domestic mortgage assets and hedged dollar assets without an additional balance-sheet wedge. In the binding regime, $\mu_t > 0$, the intermediary's Euler equations include the marginal VaR costs. The forward premium f_t then adjusts endogenously so that the covered foreign asset is priced by the same scarce balance-sheet capacity that also prices domestic credit exposure.

The simulation separates two domestic mortgage-risk disturbances. The realized-loss shock enters the default block as an increase in $Shock_t^{NPL}$, with persistence $\rho_{Default}$. Economically, this shock is interpreted as a deterioration in the realized performance of mortgage assets, the dynamic analogue of the increase in δ_1 in the two-period model. The uncertainty

shock instead enters through the perceived risk loading on domestic mortgage exposure. This second channel is important because the model's mechanism is not simply that intermediary net worth falls. It is that domestic mortgage exposure becomes more costly to carry exactly when balance-sheet capacity is scarce.

The model is calibrated so that the steady state is close to the constraint but not initially binding. This is the relevant region for the chapter's empirical prediction. In normal times, local collateral shocks need not move dollar funding premia. Near the constraint, however, the same shock can change the shadow value of intermediary risk capacity and create a macro-financial reconnect.

3.6.2 Calibration and Stochastic Simulation Design

Table 3.2 reports the central calibration choices. The household side follows standard borrower-saver housing models: patient households are more patient than borrowers, borrowers face a collateral constraint, and housing adjustment costs slow the movement of housing demand. The intermediary block follows the logic of financial intermediary DSGE models, in which net worth accumulates through retained earnings and the value of balance sheet capacity affects asset pricing (Gertler and Karadi, 2011). The occasionally binding implementation follows Guerrieri and Iacoviello (2015). The VaR logic follows the procyclical leverage and risk-management interpretation in Adrian and Shin (2014).

Table 3.2: Key Calibration Values

Households and Preferences					
β_S	Patient discount factor	0.990	κ_H	Housing adjustment cost	5.000
β_B	Borrower discount factor	0.960	m	Loan-to-value limit	0.800
Financial Intermediary and Risk					
σ_{bank}	Intermediary retention rate	0.950	σ_H^2	Domestic mortgage risk	0.040
ω_{FX}	Dollar asset share	0.200	σ_ℓ^2	Dollar liquidity risk	0.020
κ_{CIP}	Hedging demand elasticity	2.000	$\sigma_{H\ell}$	Credit-liquidity covariance	-0.015
χ_{NPL}	Risk sensitivity to NPL	10.00			
Macroeconomy and Nominal Frictions					
τ	Trade elasticity	1.500	ϕ_{adj}	Rotemberg adjustment cost	100.0
ρ_R	Taylor rule smoothing	0.950	ϕ_π	Taylor rule inflation resp.	1.200

The calibration of $\sigma_{H\ell} < 0$ is the dynamic counterpart of Assumption 1 in the two-period model. It captures the idea that dollar liquidity is especially valuable in states in which domestic mortgage returns are low. The maintained calibration implies a negative local VaR loading on hedged dollar assets. In the final simulation, $\theta_{X,t}$ remains negative throughout the binding region.

3.6.3 Stochastic Simulation Design

To evaluate the non-linear transmission of macro-financial stress, I simulate the model for $T = 1,000$ periods using the piecewise linear solution method of the OccBin algorithm. This stochastic simulation draws random structural innovations in every period, allowing the economy to transition endogenously between the slack and binding intermediary regimes

while preserving the long-run distribution of the state variables.

I structure the simulation to separate the effects of realized credit losses from forward-looking uncertainty. Both disturbances follow persistent AR(1) processes:

$$Shock_t^{NPL} = \rho^{NPL} Shock_{t-1}^{NPL} + \varepsilon_t^{NPL}, \quad (3.85)$$

$$Shock_t^{RISK} = \rho^{RISK} Shock_{t-1}^{RISK} + \varepsilon_t^{RISK}, \quad (3.86)$$

where the persistence parameters are calibrated symmetrically at $\rho^{NPL} = \rho^{RISK} = 0.90$.

To thoroughly explore the state space, I examine each shock at three distinct magnitudes. For the pure realized credit distress channel (the NPL shock), I scale the sequence of structural innovations so that the maximum realized default shock matches calibrated baselines of one, two, and three percentage points over the 1,000-period horizon. For example, in the 2pp scenario:

$$\max_{t \in [1, T]} Shock_t^{NPL} = 0.02. \quad (3.87)$$

This shock enters directly into the actual loan default equation, capturing an episodic, ex-post erosion of intermediary net worth without mechanically altering the perceived riskiness of the remaining assets.

The second set of scenarios encapsulates a pure repricing of macro-financial uncertainty. Crucially, this shock does not alter the ex-post realized default volume. Instead, it shifts the perceived risk loading on domestic mortgage exposure. The latent uncertainty process dynamically inflates the intermediary's subjective risk loading, $risk_{H,t}$, in (3.81):

$$risk_{H,t} = 1 + \chi_{NPL} Shock_t^{RISK}, \quad (3.88)$$

where $\chi_{NPL} = 10$ governs the sensitivity of intermediary risk models to heightened economic anxiety. Similar to the NPL shock, I scale the risk-state process to peak at one, two, and

three percentage points. Under the 2pp calibration:

$$\max_{t \in [1, T]} Shock_t^{RISK} = 0.02. \quad (3.89)$$

Equation (3.88) implies that at the peak of this specific uncertainty cycle, the intermediary’s perceived risk weighting on domestic mortgage exposure expands by exactly 20 percent above its non-crisis steady state ($risk_{H,t} = 1.20$). This dynamic contraction of risk-bearing capacity forces a systemic tightening of the local VaR constraint, generating the scarcity of dollar liquidity required to reproduce the empirical spikes in the global CIP basis.

3.6.4 Quantitative Analysis: Disconnect and Reconnect in General Equilibrium

Having calibrated the model, I now evaluate its quantitative ability to resolve the exchange rate disconnect puzzle. The core empirical fact documented in Section 3 is state-dependent: domestic housing-market stress does not spill over to global dollar funding markets in normal times, but triggers a massive widening of the CIP basis during periods of elevated macro-financial stress.

To formally validate whether the model’s structural mechanism can quantitatively account for these motivating facts, I conduct a simulated regression exercise. In the real-world data, the true shadow cost of intermediary capital is unobservable, which necessitated the use of a discrete stress proxy such as the VIX. To ensure an exact apples-to-apples comparison with the empirical findings, I replicate this discrete identification strategy within the simulated economy.

Specifically, I simulate the full non-linear DSGE model for 1,000 periods. To strictly mirror the empirical exercise, I define “stress periods” as the top 20 percent of the model-implied stress proxy, $VaR_Boundary_t - \nu_t$, which is directly analogous to the top-quintile VIX classification used in the data. Using this simulated time-series, I estimate the exact

analogue of the empirical interaction specification:

$$b_t^{CIP} = \beta_0 + \beta_1 \Delta q_t + \beta_2 Stress_t + \beta_3 (\Delta q_t \times Stress_t) + \varepsilon_t. \quad (3.90)$$

I isolate the transmission channels by running the simulation under the six distinct scenarios described above: pure realized default shocks (NPL) peaking at 1pp, 2pp, and 3pp, and pure domestic mortgage-risk uncertainty shocks peaking at 1pp, 2pp, and 3pp. Table 3.3 reports the regression diagnostics for each scenario.

Table 3.3: Simulated Regression Diagnostics: The Disconnect and Reconnect

Shock Scenario	Peak CIP	Housing-CIP Gradient (<i>bps per 1% Δq</i>)		
	Basis (bps)	Normal Slope (β_1)	Interaction (β_3)	Stress Slope
Realized Default (NPL 1pp)	20.1	−45.01***	4.18	−40.83***
Realized Default (NPL 2pp)	40.1	−44.98***	4.85	−40.13***
Realized Default (NPL 3pp)	60.2	−44.92***	4.84	−40.09***
Credit Uncertainty (1pp)	28.4	1.15	−58.39***	−57.24***
Credit Uncertainty (2pp)	56.7	1.16	−58.17***	−57.01***
Credit Uncertainty (3pp)	85.1	1.16	−58.14***	−56.98***

Note: Simulated over 1,000 periods using the OccBin algorithm. Stress periods are defined as the top 20% of the model-implied stress proxy $VaR_Boundary_t - \nu_t$, mirroring the empirical VIX top-quintile classification.

*** denotes statistical significance at the 1% level. For all uncertainty shocks, the normal slope (β_1) is statistically indistinguishable from zero ($p > 0.600$), matching the empirical disconnect.

The simulation results reveal a clear dichotomy. As shown in the first three rows of Table 3.3, realized default shocks fail to replicate the empirical facts, regardless of their magnitude. A realized default (NPL) shock mechanically erodes bank capital and widens the basis, but it generates a strictly negative gradient in *all* periods ($\beta_1 \approx -45$). It fails to

produce the normal-time disconnect required for the state-dependent reconnect.

In contrast, the pure uncertainty shocks (bottom three rows) reproduce the empirical pattern closely. In normal times, when the intermediary constraint is slack, the relationship between housing prices and the CIP basis is flat and statistically insignificant ($\beta_1 \approx 1.16$, $p > 0.600$). This is a direct manifestation of the exchange rate disconnect. However, when the uncertainty shock forces a sudden repricing of domestic mortgage risk, the VaR constraint binds aggressively. Intermediaries deleverage and hoard dollar liquidity. This sudden dash for dollars yields a large and highly significant interaction term ($\beta_3 \approx -58$), a dynamic that holds robustly across different shock magnitudes.

These quantitative results confirm the central theoretical claim of the paper: the structural scarcity of intermediary risk-bearing capacity, driven by forward-looking mortgage-risk uncertainty rather than pure realized losses, is the fundamental engine linking local housing markets to the global dollar funding premium.

3.6.5 Sensitivity Analysis

To isolate the economic channels driving the state-dependent reconnect, this section conducts layer-by-layer counterfactual exercises by deactivating the endogenous risk-repricing channel ($\chi_{npl} = 0$) and the macro-natural hedge channel ($cov_{H,l} = 0$). Table 3.4 reports the peak diagnostics under both a pure credit uncertainty shock and a realized non-performing loan (NPL) shock. Figures 3.2 and 3.3 overlay the corresponding impulse responses.

Under a pure second-moment uncertainty shock (Panel A), deactivating the risk-repricing channel ($\chi_{npl} = 0$) completely flattens the macroeconomic trajectories. Because uncertainty shocks carry no direct, concurrent cash-flow drain, shutting down the risk-assessment block mechanically prevents the shock from impacting the intermediary's balance sheet. The critical channel transmitting this shock to the foreign exchange market is the macro-natural hedge. When this negative covariance is muted ($cov_{H,l} = 0$), the peak CIP deviation drops

Table 3.4: Stochastic Simulation Peak Diagnostics across Structural Shocks

Scenario	Max CIP (bps)	Min q (%)	Min N_{FI} (%)	Peak μ
Panel A: Macroeconomic Uncertainty Shock				
Baseline	204.0	-1.224	-2.023	0.1318
No Risk Repricing ($\chi_{npl} = 0$)	0.0	0.000	0.000	0.0000
No Macro-Natural Hedge ($cov_{H,l} = 0$)	2.1	-0.297	0.000	0.0924
Panel B: Realized NPL Shock				
Baseline	138.0	-1.438	-26.335	0.0874
No Risk Repricing ($\chi_{npl} = 0$)	138.0	-1.438	-26.335	0.0874
No Macro-Natural Hedge ($cov_{H,l} = 0$)	2.1	-1.334	-26.046	0.0768

from 204.0 basis points to a negligible 2.1 basis points. Without the negative covariance between local mortgage performance and the funding value of dollar resources, the intermediary treats dollar liquidity as a standard risky asset rather than a regulatory hedging buffer. The critical “dash for dollars” disappears, and the cross-currency basis remains decoupled despite a binding VaR constraint ($\mu_t = 0.0924$).

A parallel analysis under a 3 percentage point realized NPL shock (Panel B) further clarifies this mapping. Because the NPL shock directly erodes bank capital via historical defaults, turning off the risk-repricing channel ($\chi_{npl} = 0$) yields identical paths to the baseline framework. By contrast, deactivating the macro-natural hedge channel ($cov_{H,l} = 0$) under the realized loss shock demonstrates a similar insulating effect in the FX swap market. Even though the intermediary suffers severe balance-sheet damage—with net worth plunging by over 26% and forcing the risk constraint to bind ($\mu_t = 0.0768$)—the annualized CIP deviation is effectively eliminated, collapsing to 2.1 basis points.

These counterfactuals deliver a concise structural insight: while financial stress can be triggered by either forward-looking risk repricing (χ_{npl}) or realized balance-sheet losses, its transmission into an international arbitrage wedge strictly requires the macro-natural hedge channel ($cov_{H,l} < 0$). The systematic co-movement between local mortgage defaults and the dynamic margin strains on the derivative book remains the indispensable link connecting local housing markets to global dollar funding premia.

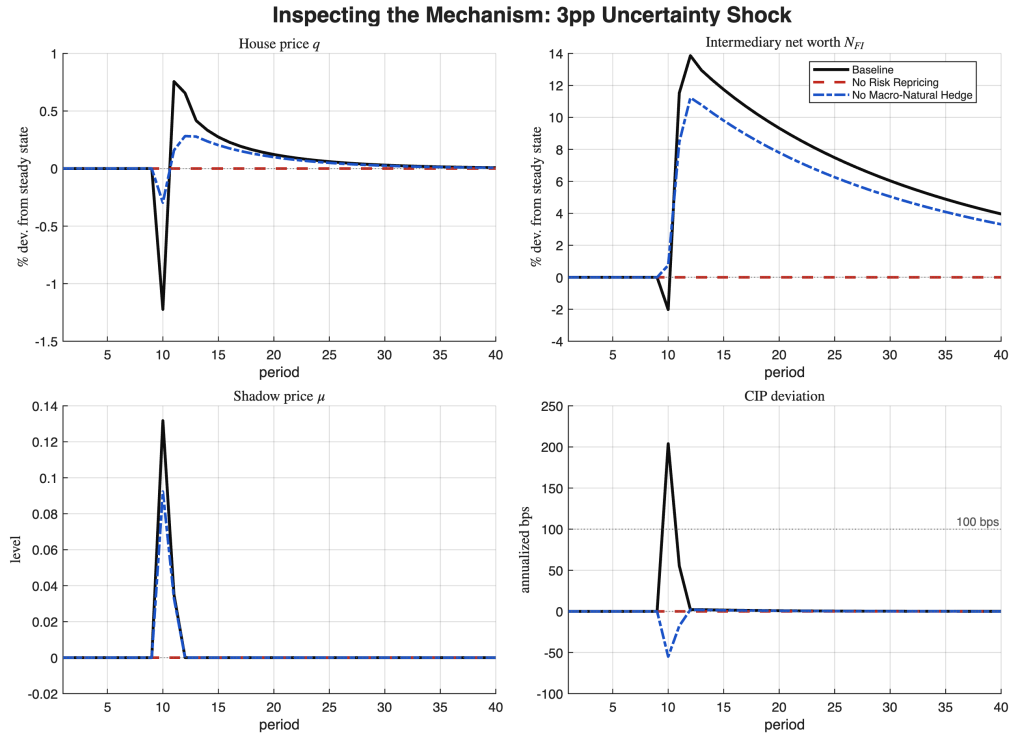


Figure 3.2: Three-Line Mechanism Overlay under an Uncertainty Shock

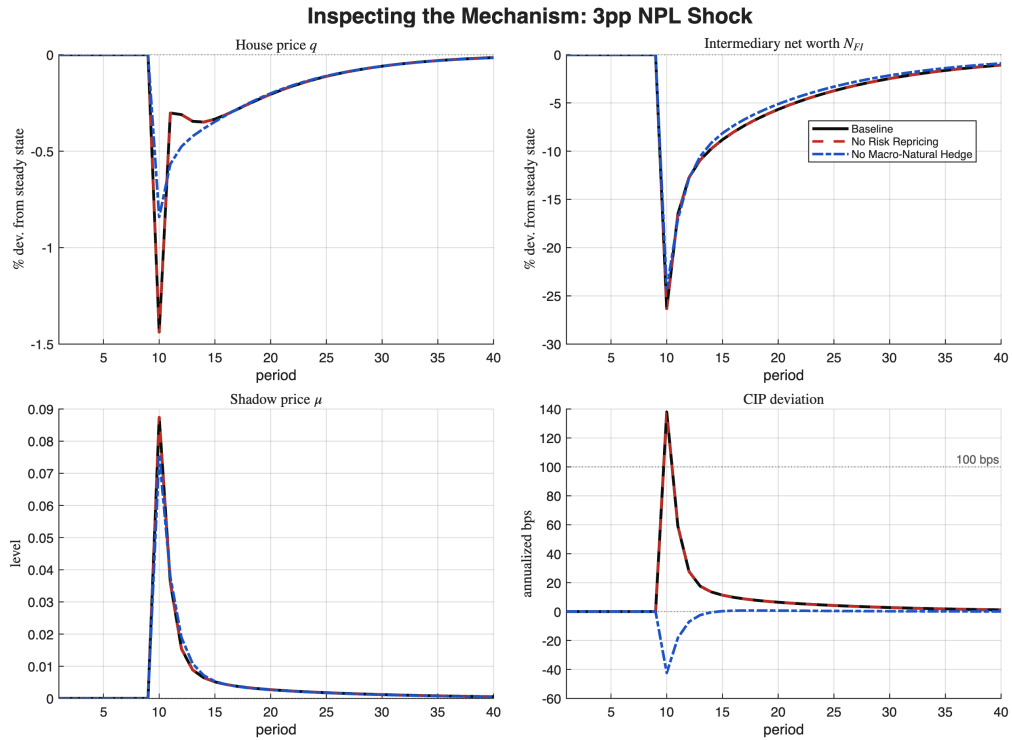


Figure 3.3: Three-Line Mechanism Overlay under a 3pp Realized NPL Shock

3.7 Conclusion

The chapter links local housing risk to global dollar funding premia through a constrained intermediary. The two-period model shows the mechanism in closed form. When the intermediary constraint binds and hedged dollar liquidity has a negative local VaR loading, the shadow value of balance sheet capacity widens the CIP basis. The full DSGE model embeds the same idea in a dynamic small open economy with housing collateral, default risk, nominal rigidities, endogenous forward premia, and external clearing. The quantitative exercise shows that forward-looking uncertainty about domestic mortgage exposure, rather than realized NPL losses alone, generates the normal-time disconnect and stress-time reconnect observed in the data. The empirical evidence supports the key state-dependence: housing price declines predict wider CIP deviations mainly when global financial conditions are tight.

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Appendix A

Chapter 1

A.1 Data Description

This appendix details the sources and construction methods for all variables used in the empirical analysis. Table A.1.1 describes the monthly data used for the Fama and Meese-Rogoff puzzles, including the construction of the key constructed variables. Table A.1.2 describes the quarterly data used for the Backus-Smith puzzle, focusing on the measurement of non-durable consumption and relative consumption.

Table A.1.1: Data for Fama and Meese-Rogoff Puzzles (Monthly, 1994m1–2019m12)

Variable	Description	Source
<i>Core Market Variables</i>		
Spot Exchange Rate ($S_{i,t}$)	Bilateral vs. USD (LCU per USD)	FRED
Interest Rate ($i_{i,t}$, $i_{us,t}$)	3-month Interbank Rate	FRED (OECD)
<i>Macroeconomic Fundamentals</i>		
Consumer Price Index (CPI)	All Items (2015=100)	FRED (OECD)
Financial Conditions (NFCI)	National Financial Conditions Index	FRED (Chicago Fed)
Trade Balance (TB)	U.S. Goods & Services Trade Balance	FRED
Gross Domestic Product (GDP)	U.S. Nominal GDP, Quarterly	FRED
VIX	CBOE Volatility Index	FRED (CBOE)
<i>Key Constructed Variables (Monthly)</i>		
Exchange Rate Change ($\Delta s_{i,t}$)	Log-difference: $s_{i,t} - s_{i,t-1}$	Authors' calculation
3M Exchange Rate Change ($\Delta s_{i,t+3}$)	Log-difference: $s_{i,t+3} - s_{i,t}$	Authors' calculation
Interest Rate Differential (i_t^d)	$(i_{us,t} - i_{i,t})/4$	Authors' calculation
Inflation ($\pi_{i,t}$)	12-month log-difference: $cpi_{i,t} - cpi_{i,t-12}$	Authors' calculation
Real Interest Rate ($r_{i,t}$)	$i_{i,t} - \pi_{i,t}$	Authors' calculation
Real Exchange Rate ($q_{i,t}$)	$s_{i,t} + cpi_{i,t} - cpi_{us,t}$	Authors' calculation
Trade Balance / GDP	U.S. TB (interpolated) / U.S. GDP (interpolated)	Authors' calculation

Notes: All quarterly series (U.S. GDP, U.S. TB, CPI for Australia and New Zealand) were converted to monthly frequency via linear interpolation (using 'ipolate') before constructing variables. For EU and Mexico, data prior to 1999m1 and 1997m1, respectively, are excluded due to availability. Exchange rates are defined as LCU per USD, so an increase is a depreciation of the local currency.

Table A.1.2: Data for Backus-Smith Puzzle (Quarterly, 1994q1–2019q4)

Variable	Description	Source
<i>Core Variables</i>		
Real Consumption ($C_{i,t}$)	Non-durable consumption, seasonally adjusted Deflated by CPI (2015=100)	OECD / Eurostat
Population ($Pop_{i,t}$)	Total Population	OECD
Real Exchange Rate ($q_{i,t}$)	Bilateral vs. USD (See Table A.1.1)	Authors' calculation
<i>Key Constructed Variables (Quarterly)</i>		
Consumption per Capita	$C_{i,t}/Pop_{i,t}$	Authors' calculation
Relative Consumption ($c_t - c_t^*$)	$\log(C_{us,t}/Pop_{us,t}) - \log(C_{i,t}/Pop_{i,t})$	Authors' calculation
Change in RER ($\Delta q_{i,t}$)	$q_{i,t} - q_{i,t-1}$	Authors' calculation
Change in Rel. Consumption ($\Delta(c_t - c_t^*)$)	$(c_t - c_t^*) - (c_{t-1} - c_{t-1}^*)$	Authors' calculation

Notes: Data for consumption by durability (non-durable) are sourced from OECD National Accounts and Eurostat for European countries. All variables are sampled at a quarterly frequency.

A.2 Full Model

This appendix outlines the full structure of the two-country New Keynesian model used in the simulation analysis. The model follows the framework of Itskhoki and Mukhin (2021), with some notational and structural adaptations for consistency with this paper's context. We focus on the home economy, assuming symmetry with the foreign country.

A.2.1 Households

A representative household maximizes lifetime utility:

$$E_0 \sum_{t=0}^{\infty} \beta^t \left(\frac{1}{1-\sigma} C_t^{1-\sigma} - \frac{1}{1+1/\nu} L_t^{1+1/\nu} \right), \quad (\text{A.2.1})$$

subject to the intertemporal budget constraint:

$$P_t C_t + P_t Z_t + \frac{B_{t+1}}{R_t} \leq W_t L_t + R_K^t K_t + B_t + \Pi_t, \quad (\text{A.2.2})$$

where K_t accumulates as:

$$K_{t+1} = (1-\delta)K_t + \left[Z_t - \frac{\kappa}{2} \left(\frac{\Delta K_{t+1}}{K_t} \right)^2 \right]. \quad (\text{A.2.3})$$

Households consume both domestic and foreign goods according to a Kimball aggregator, with within-period allocation:

$$P_t C_t = P_{Ht} C_{Ht} + P_{Ft} C_{Ft}, \quad (\text{A.2.4})$$

$$\int_0^1 \left[(1-\gamma) g \left(\frac{C_{Ht}(i)}{(1-\gamma)C_t} \right) + \gamma g \left(\frac{C_{Ft}(i)}{\gamma C_t} \right) \right] di = 1. \quad (\text{A.2.5})$$

A.2.2 Production and Pricing

Firms use capital, labor, and intermediates to produce output:

$$Y_t = \left(e^{a_t} K_t^\vartheta L_t^{1-\vartheta} \right)^{1-\phi} X_t^\phi, \quad (\text{A.2.6})$$

where productivity evolves as:

$$a_t = \rho_a a_{t-1} + \sigma_a \epsilon_t^a, \quad \epsilon_t^a \sim iid(0, 1). \quad (\text{A.2.7})$$

Nominal marginal costs are given by:

$$MC_t = \frac{1}{\bar{\omega}} \left[e^{-a_t} (R_K^t)^\vartheta W_t^{1-\vartheta} \right]^{1-\phi} P_t^\phi, \quad (\text{A.2.8})$$

where $\bar{\omega}$ collects technology parameters.

Firms price with markups based on Kimball demand elasticity. With local currency pricing (LCP):

$$P_{Ht} = \mu \left(\frac{P_{Ht}}{P_t} \right) \cdot MC_t, \quad (\text{A.2.9})$$

$$P_{Ht}^* = \mu \left(\frac{P_{Ht}^*}{P_t^*} \right) \cdot \frac{MC_t}{E_t}. \quad (\text{A.2.10})$$

A.2.3 Nominal Rigidities

Calvo frictions imply forward-looking Phillips curves. The domestic price NKPC is:

$$\pi_t^{Ht} = k_p [(1 - \alpha) m c_t + \alpha p_t - p_t^{Ht}] + \beta E_t \pi_{t+1}^{Ht}, \quad (\text{A.2.11})$$

and similarly for wages and export prices (under LCP or PCP invoicing).

A.2.4 Market Clearing and External Balance

The goods market clears both domestically and externally:

$$Y_{Ht} = (1 - \gamma)h \left(\frac{P_{Ht}}{P_t} \right) (C_t + X_t + Z_t), \quad (\text{A.2.12})$$

$$Y_{Ht}^* = \gamma h \left(\frac{P_{Ht}^*}{P_t^*} \right) (C_t^* + X_t^* + Z_t^*). \quad (\text{A.2.13})$$

The country budget constraint is:

$$\frac{B_{t+1}}{R_t} - B_t = NX_t, \quad NX_t = E_t P_{Ht}^* Y_{Ht}^* - P_{Ft} Y_{Ft}. \quad (\text{A.2.14})$$

A.2.5 Calibration

For the full derivation and estimation framework, see Itskhoki and Mukhin (2021) Appendix A.

Table A.2.1: Model Calibration (Following IM21)

Parameter	Description	Value
<i>Preferences and Demand</i>		
σ	Relative risk aversion	2.0
β	Quarterly discount factor	0.99
$1/\phi$	Frisch elasticity of labor supply	1.0
θ	Elasticity of substitution	1.5
γ	Imports-in-expenditure ratio	0.07
<i>Production and Pricing</i>		
ϕ	Intermediate input share	0.5
α	Elasticity of strategic complementarities	0.4
ϑ	Capital share in value added	0.3
δ	Quarterly capital depreciation	0.025
λ_p	Calvo price stickiness (Prob. non-adjustment)	0.75
λ_w	Calvo wage stickiness (Prob. non-adjustment)	0.85
κ	Capital adjustment cost parameter	6.8
<i>Monetary Policy</i>		
ϕ_π	Taylor rule response to inflation	2.15
ρ_π	Taylor rule smoothing	0.95
<i>Shock Processes</i>		
ρ	AR(1) persistence (all shocks)	0.97
$\text{corr}(\Delta q, \Delta c_t^*)$	Prod. shock correlation target	-0.4
$\text{corr}(\Delta gdp_t, \Delta gdp_t^*)$	Prod. shock correlation target	0.35

A.3 A Structural Break in Fama Puzzle

Table A.3.1: Pre-GFC: 1994m1 to 2007m12

	(1) UK	(2) JA	(3) CA	(4) SD	(5) NO	(6) EU	(7) SW	(8) AU	(9) NZ	(10) KO	(11) ME	(12) Panel
INT_DIFF	-1.463 (0.889)	-2.643* (1.105)	-2.071* (0.899)	-2.975*** (0.677)	-1.923** (0.607)	-4.689*** (1.170)	-4.875*** (1.060)	-4.790*** (0.931)	-3.647*** (1.002)	0.331 (0.641)	-0.143 (0.187)	-0.904*** (0.178)
_cons	0.0725 (0.355)	2.549* (1.144)	0.629* (0.259)	0.591 (0.349)	0.306 (0.339)	1.540** (0.468)	3.695*** (0.774)	-1.214* (0.472)	-1.782* (0.762)	0.0234 (0.879)	-1.044 (0.562)	0.113 (0.127)
N	168	168	168	168	168	108	168	168	168	168	132	1752
R^2	0.016	0.033	0.031	0.104	0.057	0.132	0.113	0.137	0.074	0.002	0.004	0.015
adj. R^2	0.01010	0.02749	0.02510	0.09882	0.05135	0.12336	0.10770	0.13229	0.06837	-0.00441	-0.00316	0.00841
F	2.704	5.720	5.300	19.31	10.04	16.06	21.16	26.46	13.25	0.267	0.588	25.86

Standard errors in parentheses

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table A.3.2: Post-GFC: 2008m1 to 2019m12

	(1) UK	(2) JA	(3) CA	(4) SD	(5) NO	(6) EU	(7) SW	(8) AU	(9) NZ	(10) KO	(11) ME	(12) Panel
INT_DIFF	4.936** (1.503)	3.602* (1.780)	2.229 (2.419)	1.235 (1.414)	1.902 (1.464)	2.069 (1.231)	-0.154 (1.409)	0.192 (1.348)	3.683* (1.419)	0.968 (1.405)	-3.384 (2.049)	1.521*** (0.452)
_cons	-0.701 (0.399)	-0.675 (0.513)	-0.606 (0.381)	-1.079* (0.521)	-0.634 (0.655)	-0.822* (0.411)	0.197 (0.532)	-0.624 (0.967)	1.449 (0.932)	-0.0658 (0.730)	-5.378* (2.491)	-0.332 (0.173)
N	144	144	144	144	144	144	144	144	144	144	144	1584
R^2	0.071	0.028	0.006	0.005	0.012	0.020	0.000	0.000	0.045	0.003	0.019	0.007
adj. R^2	0.06401	0.02119	-0.00106	-0.00166	0.00479	0.01262	-0.00696	-0.00690	0.03860	-0.00369	0.01193	0.00019
F	10.78	4.096	0.849	0.763	1.688	2.827	0.0120	0.0203	6.742	0.474	2.727	11.31

Standard errors in parentheses

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table A.3.3: Fama Regression Panel Estimation (1994m1-2019m12)

	(1) Full	(2) Pre-GFC	(3) Post-GFC	(4) Intercept Break	(5) Slope Break	(6) Full Break
INT_DIFF	-0.453** (0.157)	-0.904*** (0.178)	1.521*** (0.452)	-0.423** (0.157)	-0.479** (0.157)	-0.446** (0.157)
dummy_post				-0.978*** (0.185)		-0.790*** (0.198)
dummy_INT_DIFF					1.295*** (0.295)	0.856** (0.314)
_cons	-0.249* (0.0997)	0.113 (0.127)	-0.332 (0.173)	0.223 (0.134)	-0.118 (0.104)	0.219 (0.133)
N	3336	1752	1584	3336	3336	3336
R^2	0.002	0.015	0.007	0.011	0.008	0.013
adj. R^2	-0.00082	0.00841	0.00019	0.00720	0.00464	0.00911
F	8.264	25.86	11.31	18.09	13.77	14.56

Standard errors in parentheses

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

A.4 A Structural Break in Meese-Rogoff Puzzle

Table A.4.1: Out-of-Sample Forecast Performance vs. Random Walk (Clark-West Test)

Country	Pre-GFC (1995–2007)		Post-GFC (2008–2019)	
	CW-statistic	p-value	CW-statistic	p-value
GBR	1.778**	(0.038)	0.516	(0.303)
JPN	1.458*	(0.072)	0.011	(0.495)
CAN	1.245	(0.107)	3.602***	(0.000)
SWE	−0.052	(0.521)	0.940	(0.174)
NOR	−0.076	(0.530)	1.509*	(0.066)
CHE	0.253	(0.400)	0.369	(0.356)
AUS	1.806**	(0.035)	2.615***	(0.004)
NZL	0.946	(0.172)	2.686***	(0.004)
KOR	−0.133	(0.553)	1.326*	(0.092)
MEX	1.801**	(0.036)	2.469***	(0.007)
Panel (FE)	1.185	(0.118)	6.284***	(0.000)

Notes: The table reports the Clark and West (2006, 2007) test statistic and corresponding one-sided p-values for the null hypothesis that the model’s predictive performance is equal to that of a random walk. A positive and significant statistic indicates that the model outperforms the random walk. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table A.4.2: Meese and Rogoff Regression : Pre-GFC (1995m1 - 2007m12)

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
	UK	JA	CA	SD	NO	EU	SW	AU	NZ	KO	ME	Panel
D.real_INT_US	-0.0193* (0.00911)	-0.00569 (0.0163)	-0.00662 (0.0101)	0.0195 (0.0149)	-0.0159 (0.0115)	0.0136 (0.0189)	0.0154 (0.0156)	-0.0294 (0.0158)	-0.0201 (0.0161)	0.0105 (0.0159)	-0.00882 (0.0113)	-0.000324 (0.00408)
D.real_INT	0.0334*** (0.00964)	-0.00353 (0.0303)	0.00990 (0.00744)	0.00262 (0.0128)	-0.00104 (0.00712)	-0.0133 (0.0249)	-0.0164 (0.0133)	0.0194 (0.0162)	0.00367 (0.00875)	-0.0294*** (0.00332)	-0.00475*** (0.00112)	-0.00838*** (0.00121)
PI_US	-1.225** (0.442)	0.549 (0.734)	-0.504 (0.410)	-0.775 (0.626)	0.203 (0.528)	-0.875 (0.705)	-0.0876 (0.664)	-0.952 (0.726)	-0.841 (0.863)	0.338 (0.756)	0.858 (0.549)	-0.0467 (0.180)
PI	0.960** (0.354)	0.0713 (0.373)	0.539* (0.226)	0.713** (0.254)	-0.450* (0.198)	0.426 (0.802)	0.172 (0.444)	0.0976 (0.207)	0.426 (0.336)	-0.240 (0.181)	-0.0260 (0.107)	0.0140 (0.0380)
D.log_VIXCLS	0.0102 (0.00840)	-0.00936 (0.0165)	-0.0348*** (0.00916)	0.00721 (0.0137)	0.0155 (0.0113)	0.0219 (0.0162)	0.0437** (0.0138)	-0.0402** (0.0139)	-0.0237 (0.0151)	0.0156 (0.0158)	-0.0470*** (0.0110)	-0.00268 (0.00418)
L.log_RER	-0.0521* (0.0211)	-0.0745* (0.0333)	0.0195 (0.0158)	-0.00623 (0.0170)	-0.0140 (0.0164)	-0.0457 (0.0286)	-0.0262 (0.0182)	0.00839 (0.0152)	-0.00276 (0.0136)	-0.0128 (0.0176)	-0.105* (0.0452)	-0.0106* (0.00500)
TBGDP	-0.152 (0.105)	0.382 (0.274)	0.00285 (0.126)	0.0524 (0.172)	-0.202 (0.129)	-0.823 (0.503)	-0.133 (0.165)	-0.0812 (0.186)	-0.0937 (0.219)	-0.336 (0.198)	-0.500 (0.338)	-0.197*** (0.0484)
_cons	-0.0152 (0.0147)	0.306* (0.148)	0.0267 (0.0248)	0.00157 (0.0337)	-0.0415 (0.0507)	-0.0581 (0.0330)	-0.0295 (0.0255)	0.0281 (0.0295)	0.00366 (0.0277)	-0.116 (0.149)	-0.408* (0.158)	-0.0230* (0.00998)
N	156	156	156	156	156	107	156	156	156	156	131	1642
R ²	0.148	0.053	0.151	0.070	0.079	0.101	0.095	0.106	0.061	0.383	0.292	0.042
adj. R ²	0.10762	0.00840	0.11045	0.02644	0.03496	0.03705	0.05204	0.06332	0.01615	0.35409	0.25193	0.03237
F	3.670	1.188	3.749	1.601	1.802	1.583	2.216	2.497	1.363	13.14	7.254	10.27

Standard errors in parentheses

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table A.4.3: Meese and Rogoff Regression : Post-GFC (2008m1 - 2019m12)

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
	UK	JA	CA	SD	NO	EU	SW	AU	NZ	KO	ME	Panel
D.real_INT_US	-0.0285* (0.0121)	0.0106 (0.00961)	-0.00799 (0.00876)	-0.0108 (0.0110)	-0.0205* (0.0102)	-0.0222 (0.0124)	-0.00358 (0.0107)	-0.0355** (0.0112)	-0.0160 (0.0130)	-0.0185* (0.00931)	0.000652 (0.0113)	-0.0126*** (0.00323)
D.real_INT	0.0581*** (0.0141)	0.153 (0.135)	0.0380* (0.0151)	0.0239 (0.0154)	0.0235 (0.0130)	0.00831 (0.0200)	-0.00609 (0.0224)	0.0687*** (0.0137)	0.0840*** (0.0199)	0.0135 (0.0128)	-0.0110 (0.0164)	0.0282*** (0.00461)
PLUS	-0.883 (0.460)	0.319 (0.599)	-0.812 (0.636)	-1.358 (0.727)	-0.881 (0.583)	-1.190 (0.638)	-1.865** (0.663)	-0.972 (0.723)	-0.878 (0.753)	-0.869 (0.518)	-1.043 (0.771)	-0.946*** (0.189)
PI	0.255 (0.231)	0.459* (0.219)	0.0252 (0.328)	-0.310 (0.291)	-0.568* (0.240)	0.441 (0.328)	-0.470 (0.380)	-0.0882 (0.510)	0.434 (0.328)	0.118 (0.253)	0.0332 (0.274)	0.0438 (0.0789)
D.log_VIXCLS	-0.0114 (0.00825)	-0.0270* (0.0107)	-0.0517*** (0.00985)	-0.0532*** (0.0122)	-0.0240* (0.00980)	-0.0418*** (0.0108)	-0.0292* (0.0117)	-0.0703*** (0.0121)	-0.0756*** (0.0134)	-0.0241* (0.00929)	-0.0692*** (0.0125)	-0.0442*** (0.00341)
L.log_RER	-0.0149 (0.0202)	-0.0543** (0.0202)	-0.0295 (0.0212)	-0.0367 (0.0222)	-0.0195 (0.0180)	-0.0442 (0.0251)	-0.105** (0.0355)	-0.0350 (0.0201)	-0.101*** (0.0295)	-0.0180 (0.0310)	-0.0475 (0.0287)	-0.0259*** (0.00635)
TBGDP	0.717 (0.425)	0.762* (0.337)	-0.215 (0.423)	-0.678 (0.691)	0.0598 (0.449)	0.365 (0.585)	-1.002 (0.587)	0.0603 (0.631)	0.465 (0.590)	1.230* (0.520)	-0.496 (0.557)	0.255 (0.140)
_cons	0.0240 (0.0183)	0.218** (0.0756)	-0.0238 (0.0274)	-0.0578 (0.0515)	-0.0272 (0.0556)	-0.00395 (0.0264)	-0.0811* (0.0399)	-0.0126 (0.0272)	-0.0985* (0.0428)	-0.0935 (0.251)	-0.173 (0.107)	-0.0183 (0.0120)
N	144	144	144	144	144	144	144	144	144	144	144	1584
R ²	0.205	0.157	0.252	0.194	0.153	0.191	0.137	0.398	0.329	0.210	0.222	0.166
adj. R ²	0.16449	0.11381	0.21379	0.15298	0.10901	0.14937	0.09274	0.36654	0.29404	0.16982	0.18212	0.15657
F	5.022	3.624	6.555	4.690	3.499	4.587	3.088	12.82	9.509	5.179	5.549	44.41

Standard errors in parentheses

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table A.4.4: Structural Break Test for Predictability Model Regression (Panel FE)

Variable	Coefficient	Std. Err.	t-stat	$P > t $
<i>GFC-Pre Period Effects</i>				
Δ Real Interest Rate (US)	-0.00537***	(0.00098)	-5.47	0.000
Δ Real Interest Rate (Foreign)	-0.00185**	(0.00067)	-2.76	0.006
Inflation (US)	-0.00612	(0.02934)	-0.21	0.835
Inflation (Foreign)	0.00334	(0.01655)	0.20	0.840
Lagged log Real Exchange Rate	-0.01277***	(0.00252)	-5.06	0.000
Δ Financial Conditions (NFCI)	0.00452*	(0.00236)	1.91	0.056
Trade Balance / GDP (US)	-0.13925***	(0.03508)	-3.97	0.000
<i>GFC-Post Period Interaction ($\times$ Post-GFC)</i>				
Δ Real Interest Rate (US) $\times$ Post-GFC	0.00167	(0.00316)	0.53	0.597
Δ Real Interest Rate (Foreign) $\times$ Post-GFC	0.02709***	(0.00405)	6.68	0.000
Inflation (US) $\times$ Post-GFC	-0.21368***	(0.07347)	-2.91	0.004
Inflation (Foreign) $\times$ Post-GFC	-0.02109	(0.04516)	-0.47	0.640
Lagged log RER $\times$ Post-GFC	0.00003	(0.00037)	0.08	0.937
Δ Financial Conditions (NFCI) $\times$ Post-GFC	-0.07429***	(0.00483)	-15.38	0.000
Trade Balance / GDP (US) $\times$ Post-GFC	-0.03862	(0.11370)	-0.34	0.734
Post-GFC Dummy (Intercept Shift)	0.00180	(0.00382)	0.47	0.638
Constant	-0.03196***	(0.00590)	-5.41	0.000
<i>Diagnostics</i>				
Observations	5,666			
Number of Groups (Countries)	11			
R-squared (Within)	0.0793			
Country Fixed Effects	Yes			
F-test for Structural Break	41.32			
Prob > F	0.0000			

Notes: Dependent Variable: Log Change in Spot Exchange Rate ($\Delta s_{i,t}$). Panel fixed-effects (within) regression. Standard errors clustered by country are in parentheses. The Post-GFC dummy equals 1 for $t \geq 2008m1$.

*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

A.5 A Structural Break in the Backus-Smith Puzzle

To empirically investigate this structural break, my analysis begins by examining the correlation between the growth of relative consumption and changes in the real exchange rate. I conduct this analysis in two distinct ways. First, for a bilateral analysis, I use quarterly data on non-durable consumption for each country relative to the United States, paired with the bilateral real exchange rate against the U.S. dollar. Second, to assess risk-sharing in a broader multilateral context, I use the real Effective Exchange Rate (EER) provided by the Bank for International Settlements (BIS). The EER measures a currency's value against a trade-weighted basket of other currencies. To maintain consistency with this measure, I construct a corresponding "Rest of World" (ROW) consumption aggregate for each country, calculated as the trade-weighted average of its partners' consumption growth.

Table A.5.1: Backus-Smith Correlations: Pre- vs. Post-Global Financial Crisis

Country	Bilateral vs. USD			EER vs. Rest of World (ROW)		
	Pre-GFC	Post-GFC	Change	Pre-GFC	Post-GFC	Change
GBR	-0.057	0.370	0.427	-0.057	-0.143	-0.086
EUR	-0.053	0.497	0.550	0.138	0.313	0.175
JPN	-0.046	0.169	0.215	-0.089	0.129	0.218
CAN	0.129	0.514	0.385	-0.289	0.350	0.639
NZL	0.054	0.342	0.288	-0.171	0.257	0.428
AUS	0.154	0.331	0.177	-0.123	0.030	0.153
CHE	0.270	0.124	-0.146	-0.091	0.070	0.161
SWE	0.292	0.264	-0.028	-0.112	0.146	0.258
NOR	0.048	0.218	0.170	0.242	-0.178	-0.420
KOR	0.235	0.380	0.145	-0.112	0.273	0.385
MEX	-0.036	0.359	0.395	-0.019	0.175	0.194
USA	—	—	—	-0.074	-0.001	0.073
Average	0.090	0.324	0.234	-0.063	0.118	0.182

Note: The table reports the correlation between the quarterly growth of relative non-durable consumption and the quarterly change in the real exchange rate. The "Bilateral vs. USD" columns use consumption relative to the U.S. and the bilateral real exchange rate against the USD. The "EER vs. ROW" columns use the BIS real effective exchange rate and trade-weighted consumption relative to the Rest of World. The Pre-GFC period is 1995:Q1–2007:Q4; the Post-GFC period is 2008:Q1–2019:Q4. "Change" is the Post-GFC correlation minus the Pre-GFC correlation.

Additionally, to formally test for the structural break presented in the correlation analysis, I use a fixed-effects panel regression model. This approach allows for a more rigorous examination of the structural changes by controlling for country-specific, time-invariant characteristics. The model is specifically designed to identify a change in the Backus-Smith coefficient following the Global Financial Crisis (GFC) and is specified as follows:

$$\Delta \log q_{i,t} = \alpha_i + \beta_1 \Delta \log(c_{i,t}/c_{i,t}^*) + \beta_2 (D_{\text{postGFC}} \times \Delta \log(c_{i,t}/c_{i,t}^*)) + \gamma D_{\text{postGFC}} + \epsilon_{i,t} \quad (\text{A.5.1})$$

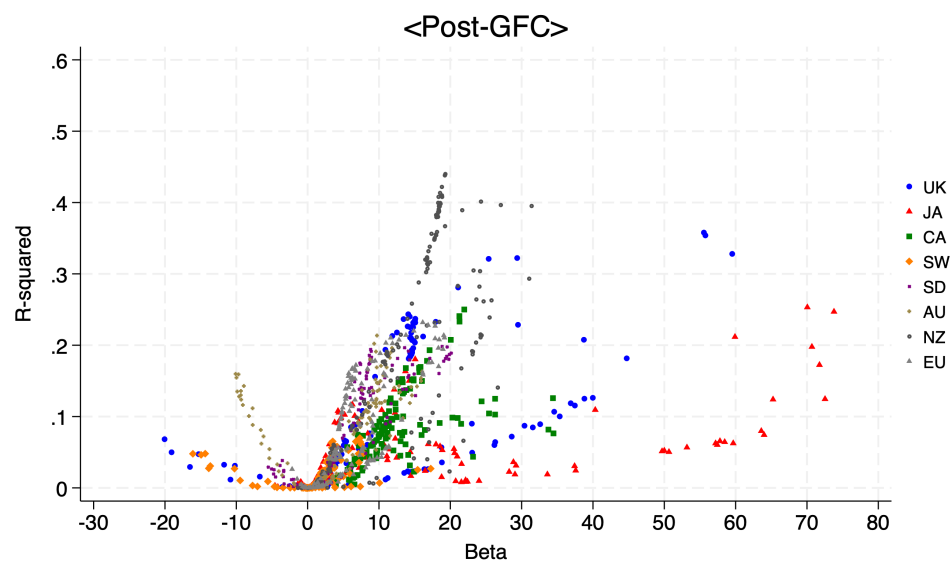
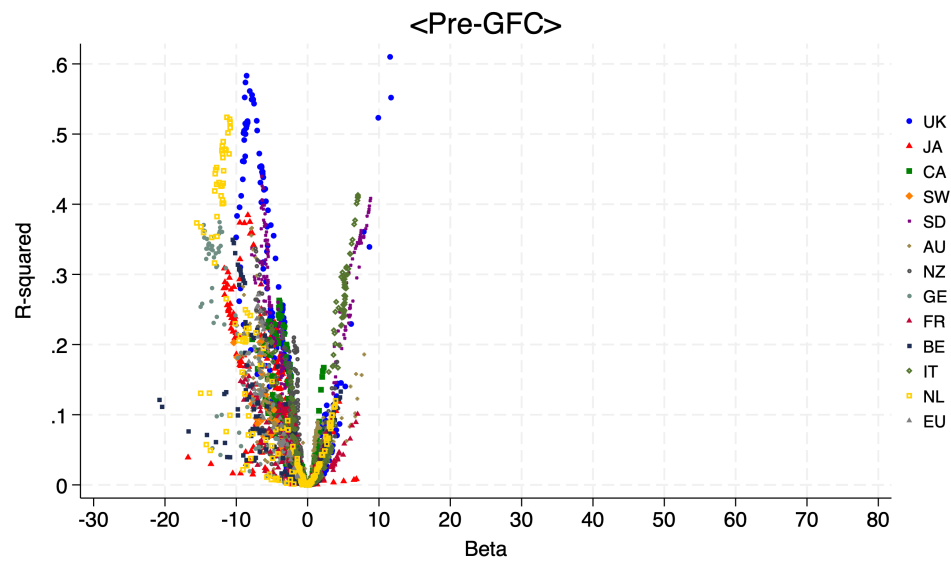
where $\Delta \log q_{i,t}$ is the change in the log real exchange rate for country i at time t , $\Delta \log(c_{i,t}/c_{i,t}^*)$ is the growth of relative consumption for country i against the U.S.A., D_{postGFC} is a dummy variable that equals 1 for the period from 2008:Q1 to 2019:Q4, and 0 otherwise. Lastly, α_i represents country-specific fixed effects, controlling for time-invariant characteristics unique to each country. β_1 is the coefficient measuring the Backus-Smith relationship in the pre-GFC period and β_2 measures the change in the Backus-Smith relationship in the post-GFC period.

Table A.5.2: Panel Regression Analysis of the Backus-Smith Puzzle

VARIABLES	(1)	(2)	(3)	(4)	(5)	(6)
	Full Sample (Pooled)	Pre-GFC (Subsample)	Post-GFC (Subsample)	Intercept Break Model	Slope Break Model	Full Interaction
	Dependent Variable: $\Delta \log q$					
$\Delta \log(c/c^*)$	0.333*** (0.067)	0.101 (0.086)	0.692*** (0.104)	0.331*** (0.066)	0.096 (0.085)	0.098 (0.085)
Post-GFC Dummy				-0.007* (0.003)		-0.007* (0.003)
Post-GFC $\times \Delta \log(c/c^*)$					0.602*** (0.135)	0.592*** (0.135)
Constant	-0.002 (0.002)	0.001 (0.002)	-0.005* (0.002)	0.001 (0.002)	-0.002 (0.002)	0.001 (0.002)
Observations	1,079	551	528	1,079	1,079	1,079
Within R-squared	0.023	0.003	0.079	0.027	0.041	0.045
Country Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes

Notes: The table presents results from panel fixed-effects regressions for the period 1995:Q1–2019:Q4. The Pre-GFC period ends in 2007:Q4, and the Post-GFC period begins in 2008:Q1. Standard errors are in parentheses. *** p<0.001, ** p<0.01, * p<0.05.

A.6 Rolling Window Scatterplot for Individual Country



A.7 Rolling Window R2-beta

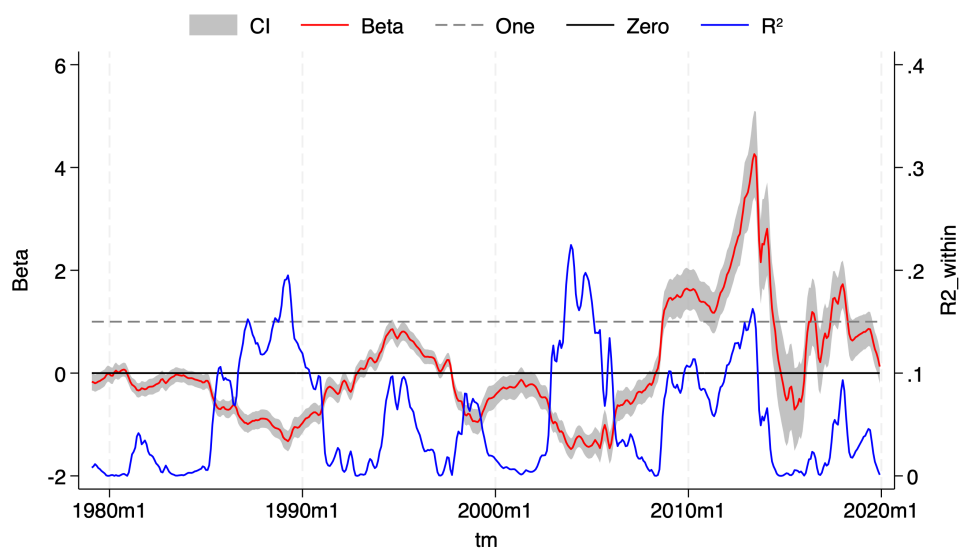


Figure A.7.1: The Asymmetric β - R^2 Relationship: This figure plots the 60-month rolling-window Fama coefficient ($\beta_{(\tau)}^f$, navy line, left axis) and the corresponding within- R^2 (red line, right axis) from the panel regression (Eq. 1.1).

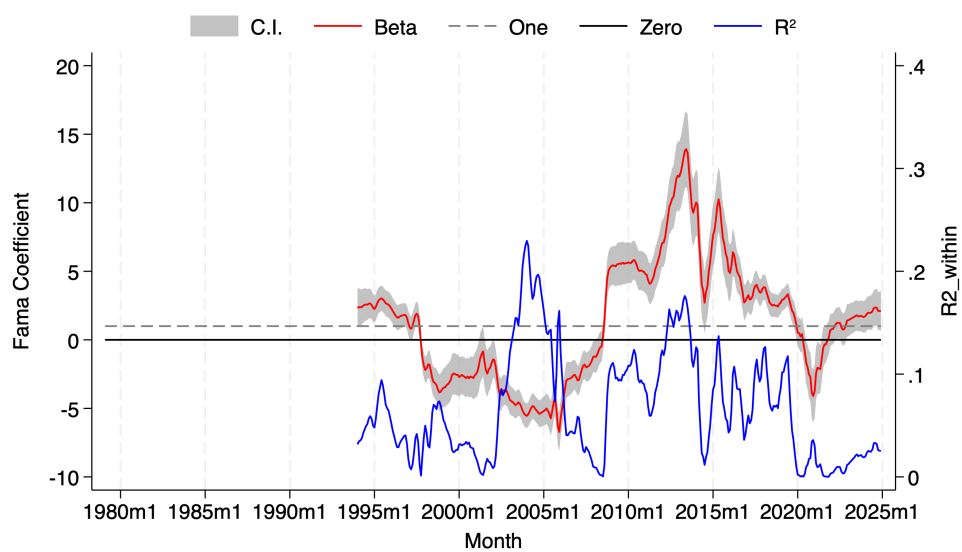


Figure A.7.2: Shadow Rate Fama estimation

A.8 beta-R2 regression

Variable	Dependent Variable: R_t^2			
	Pre-GFC		Post-GFC	
	(1)	(2)	(3)	(4)
β_t^2	0.00544*** (0.00018)	0.00551*** (0.00012)	0.00071*** (0.00004)	0.00066*** (0.00003)
$\text{Var}(i_t^d)$		0.05210*** (0.00268)		0.00083*** (0.00014)
Constant	0.01608*** (0.00220)	-0.01234*** (0.00205)	0.02089*** (0.00243)	-0.00085 (0.00268)
Observations	342	342	204	204
R^2	0.739	0.876	0.623	0.805

Standard errors in parentheses, *** p<0.01, ** p<0.05, * p<0.1

A.9 Augmented Fama Regression

As established in the theoretical discussion, the standard Fama regression suffers from an endogeneity bias arising from the central bank’s systematic policy responses, which can distort the estimated coefficient. To mitigate this bias, this section augments the baseline Fama regression by introducing control variables—specifically inflation and proxies for financial shocks—that aim to capture the key drivers of the endogenous policy response. Building on the previous insights, I estimate this augmented Fama regression using both simulated data under the two different monetary policy rules and real-world panel data, divided into the pre- and post-GFC subsamples, to assess how the inclusion of these state variables affects the Fama coefficient across policy regimes.

First, I estimate the augmented Fama regression in the simulated model using the following specification:

$$s_{t+1} - s_t = \alpha + \beta i_t^d + \gamma \pi_t^{\text{Home}} + \theta \pi_t^{\text{Foreign}} + \phi \Delta\psi_{t+1} + u_t \quad (\text{A.9.1})$$

where the change in financial shocks, $\Delta\psi_{t+1}$, is aligned with the horizon of exchange rate changes. This term captures the cumulative effect of financial shocks due to the persistence of the underlying $AR(1)$ process.

Using the data-generating process (DGP) with the conventional Taylor rule in equation (1.9) and the augmented Taylor rule in equation (1.10), I simulate 1,000 repetitions of 300-month samples and estimate the Fama regression in each simulation. This approach allows me to evaluate the bias and variability of the OLS estimator under endogeneity, depending on the specification and monetary policy rule.

Table A.9.1 reports the median coefficients from the simulated Fama regressions. Specifications (1) and (5) correspond to the standard Fama regression under the conventional and financial-stability Taylor rules, respectively. The remaining specifications include inflation

Table A.9.1: Augmented Fama regression (Simulated data)

	Conventional Taylor Rule				Financial-Stability Taylor Rule			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	$\Delta s_{i,t+1}$	$\Delta s_{i,t+1}$	$\Delta s_{i,t+1}$	$\Delta s_{i,t+1}$	$\Delta s_{i,t+1}$	$\Delta s_{i,t+1}$	$\Delta s_{i,t+1}$	$\Delta s_{i,t+1}$
Interest Diff.	-1.467 (0.727)	-1.289 (0.968)	-1.516 (1.079)	-1.332 (1.016)	1.620 (0.398)	1.553 (0.610)	1.921 (0.466)	2.009 (0.755)
π^{Home}		-0.072 (0.160)		-0.075 (0.167)		-0.006 (0.070)		0.010 (0.076)
$\pi^{Foreign}$		0.067 (0.166)		0.071 (0.174)		0.287 (0.300)		0.206 (0.308)
$\Delta \psi_{t+1}$			-0.826 (1.007)	-1.043 (1.004)			3.856 (2.330)	3.519 (2.621)
Constant	0.037 (1.042)	0.144 (11.247)	0.038 (1.080)	0.153 (11.806)	0.006 (2.219)	0.702 (26.362)	-0.023 (2.747)	-0.403 (28.435)
R^2	0.012 (0.008)	0.024 (0.013)	0.017 (0.011)	0.031 (0.014)	0.041 (0.012)	0.059 (0.015)	0.052 (0.014)	0.068 (0.016)

Median of 1,000 simulations; standard deviations in parentheses.

measures and financial shock proxies as control variables.

Two key results stand out. First, the sign of the interest rate differential coefficient (i_t^d) reverses across the two regimes: it is consistently negative under the conventional Taylor rule (columns 1–4) and positive under the financial-stability rule (columns 5–8). This pattern replicates the empirical shift observed in real-world data between the pre- and post-GFC periods. However, controlling for inflation and financial shocks fails to mitigate the endogeneity problem in the estimation of the Fama coefficients.

Second, the coefficient on the change in financial shocks, $\Delta \psi_{t+1}$, also flips sign. Under the conventional rule, $\Delta \psi_{t+1}$ enters with a negative coefficient (columns 3 and 4), whereas under the financial-stability rule, the coefficient is strongly positive (columns 7 and 8). This suggests that increases in financial shocks are associated with a contemporaneous appreciation of the home currency in the pre-GFC regime. This result is mechanically consistent with the model’s structure: the country budget constraint induces a sharp depreciation at the moment of the shock, followed by a gradual appreciation afterward—resembling the dynamics of Dornbusch’s overshooting mechanism (Itskhoki and Mukhin, 2021).

However, when monetary policy behavior is adjusted to respond directly to financial shocks, the sign of the coefficient reverses. This highlights the importance of policy regime interactions in determining how financial shocks propagate to exchange rate movements.

Appendix B

Chapter 2

B.1 Full Model

This appendix outlines the full structure of the two-country New Keynesian model used in the simulation analysis. The model builds upon the framework of Itskhoki and Mukhin (2021), with specific structural adaptations to capture the endogenous relationship between monetary policy and financial frictions. I focus on the home economy, assuming symmetry with the foreign country.

B.1.1 Households

A representative household maximizes lifetime utility:

$$E_0 \sum_{t=0}^{\infty} \beta^t \left(\frac{1}{1-\sigma} C_t^{1-\sigma} - \frac{1}{1+1/\nu} L_t^{1+1/\nu} \right), \quad (\text{B.1.1})$$

subject to the intertemporal budget constraint:

$$P_t C_t + P_t Z_t + \frac{B_{t+1}}{R_t} + \frac{E_t B_{t+1}^*}{R_t^* \Theta_t} \leq W_t L_t + R_K^t K_t + B_t + E_t B_t^* + \Pi_t, \quad (\text{B.1.2})$$

where K_t accumulates with standard quadratic adjustment costs, and Θ_t represents the financial friction (intermediation premium) in the international bond market. Households consume both domestic and foreign goods according to a Kimball aggregator.

B.1.2 Production, Pricing, and Nominal Rigidities

Firms use capital and labor to produce output subject to standard Cobb-Douglas technology and AR(1) productivity shocks. Following Itskhoki and Mukhin (2021), firms price with markups based on Kimball demand elasticity. Nominal rigidities are introduced via Calvo frictions, yielding standard forward-looking New Keynesian Phillips Curves (NKPC) for domestic prices, wages, and export prices under Local Currency Pricing (LCP).

B.1.3 Monetary Policy

The central bank sets the short-term nominal interest rate $i_t = \log(R_t)$ according to a standard Taylor rule:

$$i_t = \rho_i i_{t-1} + (1 - \rho_i)(\phi_\pi \pi_t + \phi_y y_t) + \epsilon_t^m, \quad (\text{B.1.3})$$

where $\epsilon_t^m \sim iid(0, \sigma_m^2)$ is the exogenous structural monetary policy shock.

B.1.4 International Asset Markets and the Endogenous UIP Wedge

The core mechanism of delayed overshooting in this framework operates through the international asset market. Log-linearizing the first-order conditions for domestic and foreign bond holdings yields the Uncovered Interest Rate Parity (UIP) condition:

$$i_t - i_t^* = E_t \Delta e_{t+1} + \psi_t, \quad (\text{B.1.4})$$

where e_t is the log nominal exchange rate (home currency per unit of foreign currency), and $\psi_t = \log(\Theta_t)$ is the time-varying UIP wedge, traditionally interpreted as a financial shock or

risk premium.

Departing from the standard assumption of a strictly exogenous ψ_t , I model the UIP wedge as a reduced-form function of the monetary policy shock to capture the balance sheet constraints of global financial intermediaries:

$$\psi_t = \rho_\psi \psi_{t-1} + u_t^\psi + \Gamma(S_{t-1}) \epsilon_t^m. \quad (\text{B.1.5})$$

Here, u_t^ψ is an exogenous noise component, and $\Gamma(S_{t-1}) \geq 0$ is a state-dependent elasticity parameter governing the transmission of monetary policy to financial frictions.

This formulation mechanically links the monetary policy stance to the limits to arbitrage. A contractionary monetary policy shock ($\epsilon_t^m > 0$) raises funding costs and constraints the risk-bearing capacity of dealer banks. This structural friction is captured by an increase in the required premium ψ_t . Consequently, the term $(i_t - i_t^* - \psi_t)$ rises by less than the nominal interest rate differential itself. This muted adjustment in the required expected depreciation ($E_t \Delta e_{t+1}$) prevents the exchange rate from exhibiting the instantaneous, frictionless Dornbusch appreciation on impact, thereby generating the delayed overshooting trajectory observed in the data.

B.1.5 Calibration

Table B.1.1: Model Calibration

Parameter	Description	Value
<i>Preferences and Technology</i>		
σ	Relative risk aversion	2.0
β	Quarterly discount factor	0.99
γ	Imports-in-expenditure ratio	0.07
λ_p	Calvo price stickiness	0.75
<i>Monetary Policy and Frictions</i>		
ϕ_π	Taylor rule response to inflation	2.15
ρ_i	Taylor rule smoothing	0.85
ρ_ψ	Persistence of financial shock	0.90
Γ	Sensitivity of UIP wedge to MP shock	1.50 (High Friction)
		0.00 (Low Friction)

Note: Standard macroeconomic parameters follow Itskhoki and Mukhin (2021)

B.2 Diagnostic: Shock-Induced State Transitions

The static Kitagawa-Blinder-Oaxaca (KBO) decomposition assumes that the macroeconomic state is determined at time $t - 1$ and remains invariant to the monetary policy shock over the projection horizon. If the state variable endogenously transitions in response to the shock, the static decomposition becomes invalid. This appendix verifies the validity of the chosen 5-year moving average state variable. Figure B.2.1 plots the dynamic response of 3-year, 5-year, and 7-year moving average state variables to a one-standard-deviation contractionary monetary policy shock. The results demonstrate the econometric necessity of the 5-year window. For the 3-year moving average, the point estimate drifts upwards, and the 90 percent confidence interval strictly excludes zero after 12 months. This indicates that a shorter window is endogenously contaminated by the policy shock. In contrast, the baseline 5-year and alternative 7-year moving averages exhibit no statistically significant response; their 90 percent confidence intervals comfortably encompass zero across the entire 25-month horizon. Along with extreme unconditional persistence—the AR(1) coefficients for the 3-year, 5-year, and 7-year measures are 0.9989, 0.9995, and 0.9997, respectively—these diagnostics confirm that the 5-year state variable remains strictly “frozen” relative to the policy shock. This

validates the methodology used in the main text to isolate the limits-to-arbitrage wedge without requiring complex transition-mediated modeling.

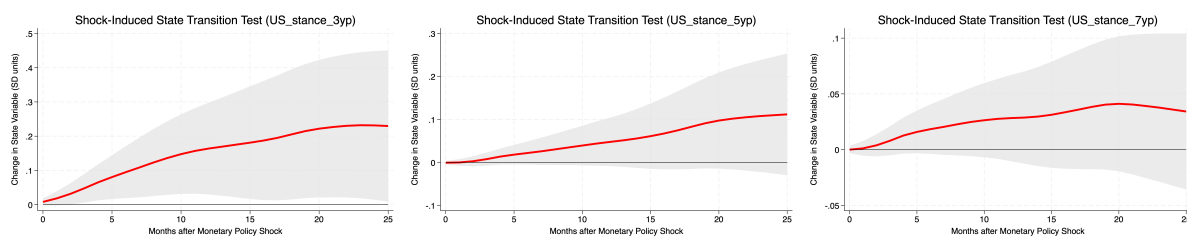


Figure B.2.1: **Shock-Induced State Transition Diagnostics.** The figure plots the cumulative change in the 3-year, 5-year, and 7-year moving average state variables following a one-standard-deviation contractionary monetary policy shock. The solid red lines indicate the point estimates, and shaded areas denote 90% confidence intervals. While the 3-year measure exhibits significant endogenous transitions, the 5-year and 7-year measures remain invariant to the shock over the entire projection horizon.

Appendix C

Chapter 3

C.1 Endogenous macro-natural hedge

Proposition 4 (Endogenous macro-natural hedge). *The negative covariance assumed in the intermediary's partial equilibrium problem (Assumption 1) is endogenously verified in the general equilibrium. A realization of the aggregate default shock at $t = 1$ simultaneously lowers domestic credit returns and increases the state-contingent liquidity value of the hedged dollar position.*

Proof. Let δ_1 be the realization of the domestic credit shock at $t = 1$. The ex-post return on domestic credit is strictly decreasing in the default rate:

$$\frac{\partial R_{H,1}}{\partial \delta_1} = -R_{B,0} < 0. \quad (\text{C.1.1})$$

Concurrently, the realization of a severe default shock tightens the intermediary's balance sheet capacity, compressing domestic expenditure and forcing the period-1 exchange rate S_1 to depreciate (i.e., S_1 rises) to clear the external balance. Since the liquidity service of the

dollar position is defined as $\ell_1 = \varphi R_0^* \frac{S_1}{S_0}$, it follows that:

$$\frac{\partial \ell_1}{\partial \delta_1} = \frac{\varphi R_0^*}{S_0} \frac{\partial S_1}{\partial \delta_1} > 0. \quad (\text{C.1.2})$$

Because $R_{H,1}$ and ℓ_1 move in strictly opposite directions in response to the fundamental macroeconomic state δ_1 , their covariance over the state space is strictly negative:

$$\text{Cov}_0(R_{H,1}, \ell_1) < 0. \quad (\text{C.1.3})$$

This verifies Assumption 1 and confirms that the hedged dollar position acts as an endogenous macro-natural hedge. $\square$

C.2 Full DSGE Model

I develop a Small Open Economy (SOE) New Keynesian model characterized by segmented financial markets. The model features heterogeneous households (Savers and Borrowers) and a global financial intermediary sector that acts as the marginal pricer of risk. This setup combines the housing collateral constraints of Iacoviello (2005) with the intermediary friction of Gertler and Karadi (2011), augmented by a CIP-based dollar funding wedge and a VaR-style balance sheet constraint whose foreign-asset loading can be negative.

The world consists of a Home country and a Foreign country (denoted by asterisk *). Time is discrete $t = 0, 1, 2, \dots$

C.2.1 Households

There is a continuum of households indexed by $i \in [0, 1]$. A fraction γ are patient Savers (S) and $1 - \gamma$ are impatient Borrowers (B), distinguished by their discount factors $\beta_S > \beta_B$.

Patient Households (Savers)

Patient households provide deposits to the banking sector and supply labor. They maximize the following lifetime utility:

$$U_S = \mathbb{E}_0 \sum_{t=0}^{\infty} \beta_S^t \left[\ln C_{S,t} + \xi_{H,t} \ln H_{S,t} - \frac{N_{S,t}^{1+\eta}}{1+\eta} \right] \quad (\text{C.2.1})$$

where $C_{S,t}$ is consumption, $H_{S,t}$ is housing services, and $N_{S,t}$ is labor hours. The housing preference $\xi_{H,t}$ follows an AR(1) process: $\ln(\xi_{H,t}/\xi_H) = \rho_{\xi_H} \ln(\xi_{H,t-1}/\xi_H) + e_{\xi_{H,t}}$.

The saver's budget constraint in real terms, incorporating quadratic housing adjustment costs, is given by:

$$C_{S,t} + d_t + q_t(H_{S,t} - H_{S,t-1}) + \frac{\kappa_H}{2}(H_{S,t} - H_{S,t-1})^2 = w_t N_{S,t} + R_{D,t-1} \frac{d_{t-1}}{\pi_t} + \frac{Div_t}{\gamma} + \Pi_t^F \quad (\text{C.2.2})$$

where d_t denotes real deposits, q_t is the real housing price, w_t is the real wage, $R_{D,t}$ is the gross nominal deposit rate, π_t is gross inflation, κ_H represents the housing transaction rigidity, and Div_t represents net dividends from financial intermediaries.

Let $\lambda_{S,t}$ be the real Lagrange multiplier on the budget constraint. The First Order Conditions (FOCs) are:

$$[C_{S,t}] : \quad \lambda_{S,t} = \frac{1}{C_{S,t}} \quad (\text{C.2.3})$$

$$[N_{S,t}] : \quad w_t = \frac{N_{S,t}^\eta}{\lambda_{S,t}} \quad (\text{C.2.4})$$

$$[d_t] : \quad 1 = \beta_S \mathbb{E}_t \left[\frac{\lambda_{S,t+1}}{\lambda_{S,t}} \frac{R_{D,t}}{\pi_{t+1}} \right] \quad (\text{C.2.5})$$

$$[H_{S,t}] : \quad q_t + \kappa_H(H_{S,t} - H_{S,t-1}) = \frac{\xi_{H,t}}{H_{S,t} \lambda_{S,t}} + \beta_S \mathbb{E}_t \left[\frac{\lambda_{S,t+1}}{\lambda_{S,t}} (q_{t+1} + \kappa_H(H_{S,t+1} - H_{S,t})) \right] \quad (\text{C.2.6})$$

Impatient Households (Borrowers)

Impatient households derive utility from similar sources but are distinguished by a lower discount factor ($\beta_B < \beta_S$) and are constrained in their borrowing capacity. They maximize:

$$U_B = \mathbb{E}_0 \sum_{t=0}^{\infty} \beta_B^t \left[\ln C_{B,t} + \xi_{H,t} \ln H_{B,t} - \frac{N_{B,t}^{1+\eta}}{1+\eta} \right] \quad (\text{C.2.7})$$

The budget constraint in real terms incorporates an endogenous default rate δ_t on past debt obligations:

$$C_{B,t} + q_t(H_{B,t} - H_{B,t-1}) + \frac{\kappa_H}{2}(H_{B,t} - H_{B,t-1})^2 + (1 - \delta_t)R_{B,t-1} \frac{b_{t-1}}{\pi_t} = w_t N_{B,t} + b_t \quad (\text{C.2.8})$$

where b_t denotes the real stock of debt at the end of period t , and $R_{B,t}$ is the gross nominal borrowing rate set at time t .

Crucially, borrowing is subject to a collateral constraint based on the expected future value of housing holdings. The constraint limits the principal such that the repayment does not exceed a fraction m of the expected liquidation value of the house:

$$b_t \leq m \mathbb{E}_t \left[\frac{q_{t+1} \pi_{t+1}}{R_{B,t}} \right] H_{B,t} \quad (\text{C.2.9})$$

Let $\lambda_{B,t}$ be the real Lagrange multiplier on the budget constraint and Γ_t be the multiplier on the borrowing constraint. The First Order Conditions (FOCs) are:

$$[C_{B,t}] : \quad \lambda_{B,t} = \frac{1}{C_{B,t}} \quad (\text{C.2.10})$$

$$[N_{B,t}] : \quad w_t = \frac{N_{B,t}^\eta}{\lambda_{B,t}} \quad (\text{C.2.11})$$

$$[b_t] : \quad \lambda_{B,t} - \Gamma_t = \beta_B \mathbb{E}_t \left[\lambda_{B,t+1} (1 - \delta_{t+1}) \frac{R_{B,t}}{\pi_{t+1}} \right] \quad (\text{C.2.12})$$

$$[H_{B,t}] : \quad (q_t + \kappa_H(H_{B,t} - H_{B,t-1}))\lambda_{B,t} = \frac{\xi_{H,t}}{H_{B,t}} + \beta_B \mathbb{E}_t [\lambda_{B,t+1} (q_{t+1} + \kappa_H(H_{B,t+1} - H_{B,t}))] \\ + \Gamma_t m \mathbb{E}_t \left[\frac{q_{t+1} \pi_{t+1}}{R_{B,t}} \right] \quad (\text{C.2.13})$$

Equation (C.2.12) shows that the borrower's Euler equation internalizes the expected future default rate $(1 - \delta_{t+1})$. Furthermore, when the collateral constraint binds ($\Gamma_t > 0$), the marginal utility of consumption today exceeds the discounted marginal cost of repayment.

C.2.2 Consumption Aggregation and Prices

The aggregate consumption C_t is a CES composite of Home ($C_{H,t}$) and Foreign ($C_{F,t}$) goods:

$$C_t = \left[(1 - \alpha)^{\frac{1}{\tau}} C_{H,t}^{\frac{\tau-1}{\tau}} + \alpha^{\frac{1}{\tau}} C_{F,t}^{\frac{\tau-1}{\tau}} \right]^{\frac{\tau}{\tau-1}} \quad (\text{C.2.14})$$

where α is the degree of openness and τ is the elasticity of substitution between Home and Foreign goods.

The Consumer Price Index (CPI) follows:

$$P_t = \left[(1 - \alpha) P_{H,t}^{1-\tau} + \alpha P_{F,t}^{1-\tau} \right]^{\frac{1}{1-\tau}} \quad (\text{C.2.15})$$

where $P_{F,t} = S_t P_t^*$ assuming the Law of One Price holds for traded goods, with S_t being the

nominal exchange rate. The demand functions are given by standard CES algebra:

$$C_{H,t} = (1 - \alpha) \left(\frac{P_{H,t}}{P_t} \right)^{-\tau} C_t, \quad C_{F,t} = \alpha \left(\frac{P_{F,t}}{P_t} \right)^{-\tau} C_t \quad (\text{C.2.16})$$

C.2.3 Firms and Price Setting

To capture nominal rigidities, I assume that firms operate in a monopolistically competitive environment subject to Rotemberg pricing adjustment costs. Firms produce differentiated domestic goods using a linear technology in labor: $Y_t = A_t N_t$, where A_t is an exogenous productivity shock.

The real marginal cost is given by $mc_t = \frac{w_t/A_t}{p_{H,t}}$, where $p_{H,t} = P_{H,t}/P_t$ is the relative price of domestic goods. Profit maximization yields the New Keynesian Phillips Curve (NKPC) for domestic inflation $\pi_{H,t}$:

$$(\pi_{H,t} - \bar{\pi})\pi_{H,t} = \beta_S \mathbb{E}_t \left[\frac{\lambda_{S,t+1}}{\lambda_{S,t}} \frac{Y_{t+1}}{Y_t} (\pi_{H,t+1} - \bar{\pi})\pi_{H,t+1} \right] + \frac{\varepsilon}{\phi_{adj}} \left(\frac{w_t/A_t}{p_{H,t}} - \frac{\varepsilon - 1}{\varepsilon} \right) \quad (\text{C.2.17})$$

where ε is the elasticity of substitution among domestic varieties, and ϕ_{adj} is the Rotemberg parameter. The relative price of domestic goods evolves according to $p_{H,t} = p_{H,t-1} \frac{\pi_{H,t}}{\pi_t}$.

Firms's profit defines :

$$\Pi_t^F = p_{H,t} * Y_t - w_t N_t - \frac{\phi_{adj}}{2} (\pi_{H,t} - \bar{\pi})^2 Y_t \quad (\text{C.2.18})$$

C.2.4 The Representative Financial Intermediary

I consider a representative financial intermediary that manages the consolidated balance sheet of the banking sector. The intermediary channels funds from patient households (depositors) to impatient households (mortgage borrowers) and acts as the sole counterparty to exogenous hedgers and liquidity traders in the foreign exchange market.

Asset Returns and Net Worth Accumulation

The intermediary's domestic asset $A_{H,t}$ represents aggregate real mortgage loans, b_t . Because mortgage contracts are nominal with a predetermined gross rate $R_{B,t-1}$, the realized real return on domestic assets is exposed to inflation and an endogenous default rate δ_t :

$$R_{H,t} = (1 - \delta_t) \frac{R_{B,t-1}}{\pi_t} \quad (\text{C.2.19})$$

A key feature of this framework is the endogenous formulation of the default rate δ_t on domestic mortgage loans. Rather than assuming a constant survival rate of loans, I model aggregate credit risk as a strictly decreasing function of the collateral asset value, specifically the real housing price q_t . The default rate evolves according to a logistic function:

$$\delta_t = \frac{\bar{\delta}}{\bar{\delta} + (1 - \bar{\delta}) \exp\left(\kappa_d \left(\frac{q_t}{q_{ss}} - 1\right)\right)} \quad (\text{C.2.20})$$

where $\bar{\delta}$ is the steady-state default rate, q_{ss} is the steady-state housing price, and $\kappa_d > 0$ dictates the sensitivity of defaults to asset price fluctuations.

The foreign asset $A_{FX,t}$ is a covered dollar position. The intermediary buys the foreign safe asset in the spot market and simultaneously sells the foreign currency forward, so its contractual payoff is governed by the FX swap rate rather than by the realized spot depreciation. Let

$$f_t \equiv \frac{F_t}{S_t} \quad (\text{C.2.21})$$

denote the gross nominal forward premium. The covered real return paid at $t + 1$ on a position chosen at t is

$$R_{FX,t+1} = \frac{R_t^* f_t}{\pi_{t+1}}. \quad (\text{C.2.22})$$

This is the standard macro-finance FX swap payoff: the foreign nominal risk-free payoff R_t^* is converted back into domestic currency at the pre-arranged forward rate F_t . Hence

the contractual return is fully hedged and does not contain the realized change in the spot exchange rate. The uncovered payoff is kept only as a state-value object for dollar liquidity,

$$R_{U,t} = \frac{RER_t}{RER_{t-1}} R_{t-1}^*, \quad (\text{C.2.23})$$

with

$$\ell_t = \phi_\ell \frac{R_{U,t}}{R^*}. \quad (\text{C.2.24})$$

Thus, spot depreciation can raise the state value of dollar liquidity, but it does not mechanically recapitalize the intermediary through the contractual return on its hedged FX position. The forward premium f_t , not the spot depreciation, is the price that adjusts in the covered funding market.

The gross real return on the intermediary's consolidated portfolio is the weighted average of these returns:

$$R_{port,t} = \frac{A_{H,t-1} R_{H,t} + RER_{t-1} A_{FX,t-1} R_{FX,t}}{A_{Total,t-1}} \quad (\text{C.2.25})$$

where $A_{Total,t} = A_{H,t} + RER_t A_{FX,t}$.

Net Worth Accumulation. The representative intermediary accumulates net worth based on the difference between the return on its portfolio and the cost of deposits, net of operational management costs. I assume these costs increase quadratically with the absolute scale of total assets, taking the form $\frac{\psi}{2} A_{Total,t-1}^2$.

Using the aggregate balance sheet identity $A_{Total,t-1} = N_{FI,t-1} + d_{t-1}$, the pure internal accumulation of net worth realized at period t is derived as follows:

$$\begin{aligned} N_{FI,t}^{pure} &= R_{port,t} A_{Total,t-1} - \frac{R_{D,t-1}}{\pi_t} d_{t-1} - \frac{\psi}{2} A_{Total,t-1}^2 \\ &= \left[\left(R_{port,t} - \frac{R_{D,t-1}}{\pi_t} - \frac{\psi}{2} A_{Total,t-1} \right) \phi_{t-1} + \frac{R_{D,t-1}}{\pi_t} \right] N_{FI,t-1} \end{aligned} \quad (\text{C.2.26})$$

$$\equiv ROE_t \times N_{FI,t-1} \quad (\text{C.2.27})$$

where $\phi_{t-1} = A_{Total,t-1}/N_{FI,t-1}$ represents the aggregate leverage ratio chosen in the previous period. Equation (C.2.26) defines the pure internal return on equity, ROE_t .

I assume the infinitely-lived representative intermediary retains a constant fraction σ_{bank} of its internal earnings to build net worth, and pays out the remaining fraction $1 - \sigma_{bank}$ as dividends. Furthermore, to prevent the banking sector from shrinking to a trivial zero-net-worth state following severe crisis episodes, I assume the patient household provides an exogenous, lump-sum equity injection W_{inj} every period. Thus, the aggregate net worth evolves as:

$$N_{FI,t} = \sigma_{bank} N_{FI,t}^{pure} + W_{inj} = \sigma_{bank} ROE_t N_{FI,t-1} + W_{inj} \quad (C.2.28)$$

The dividend payout to the patient household, Div_t , is the distributed portion of the realized earnings net of this exogenous capital injection:

$$Div_t = (1 - \sigma_{bank}) ROE_t N_{FI,t-1} - W_{inj} \quad (C.2.29)$$

The Agency Problem and Objective Function

The banker manages the financial intermediary to maximize the expected present value of the stream of future dividends, discounted by the household's stochastic discount factor $\Lambda_{t,t+j}^S$. The total franchise value of the intermediary, V_t , is written recursively as:

$$V_t = \mathbb{E}_t [\Lambda_{t,t+1}^S (Div_{t+1} + V_{t+1})] \quad (C.2.30)$$

Because the intermediary receives an exogenous lump-sum injection W_{inj} every period, the total value function V_t is not strictly proportional to current net worth. Instead, I conjecture an affine value function:

$$V_t(N_{FI,t}) = \nu_t N_{FI,t} + \Upsilon_t \quad (C.2.31)$$

where ν_t is the time-varying marginal value of an additional unit of endogenous net worth, and Υ_t represents the present discounted value of all future exogenous equity injections (W_{inj}).

When the banker makes portfolio allocation decisions ($A_{H,t}, A_{FX,t}$), they evaluate the objective function at the margin. Because the future sequence of W_{inj} (and thus Υ_t) is strictly independent of the bank's internal portfolio choices, the derivative of the franchise value with respect to internally generated equity drops the Υ_t term entirely. Therefore, the relevant metric for the banker's incentive constraint is strictly the marginal continuation value ν_t .

Substituting Div_{t+1} and the affine guess into Equation (C.2.30), and matching coefficients for the terms multiplying $N_{FI,t}$, we can isolate the recursive equation for the marginal value of net worth:

$$\nu_t = \mathbb{E}_t [\Lambda_{t,t+1}^S (1 - \sigma_{bank} + \sigma_{bank}\nu_{t+1}) ROE_{t+1}] \equiv \mathbb{E}_t [\Lambda_{t,t+1}^S \Omega_{t+1} ROE_{t+1}] \quad (\text{C.2.32})$$

where $\Omega_{t+1} \equiv (1 - \sigma_{bank}) + \sigma_{bank}\nu_{t+1}$ represents the shadow marginal value of a unit of internal funds. This ν_t cleanly isolates the endogenous financial accelerator mechanism from the exogenous recapitalization dynamics.

C.3 Derivation of Optimal Portfolio Choice

The representative banker maximizes the franchise value V_t , defined as the expected present value of future dividends. The optimization is subject to the incentive constraint $V_t \geq \theta A_{Total,t}$.

1. Evolution of Net Worth. First, we express the evolution of net worth in terms of individual asset choices. The balance sheet identity implies that deposits are the residual of

total assets minus net worth:

$$d_t = A_{Total,t} - N_{FI,t} = (A_{H,t} + S_t A_{FX,t}) - N_{FI,t} \quad (\text{C.3.1})$$

Substituting this into the law of motion for net worth, and incorporating the quadratic management cost $\frac{\psi}{2} A_{Total,t}^2$, we obtain:

$$\begin{aligned} N_{FI,t+1} &= R_{H,t+1} A_{H,t} + R_{FX,t+1} S_t A_{FX,t} - \frac{R_{D,t}}{\pi_{t+1}} d_t - \frac{\psi}{2} A_{Total,t}^2 \\ &= \left(R_{H,t+1} - \frac{R_{D,t}}{\pi_{t+1}} \right) A_{H,t} + \left(R_{FX,t+1} - \frac{R_{D,t}}{\pi_{t+1}} \right) S_t A_{FX,t} \\ &\quad + \frac{R_{D,t}}{\pi_{t+1}} N_{FI,t} - \frac{\psi}{2} (A_{H,t} + S_t A_{FX,t})^2 \end{aligned} \quad (\text{C.3.2})$$

2. The Lagrangian Problem. The banker chooses $\{A_{H,t}, A_{FX,t}\}$ to maximize the objective function subject to the incentive constraint. The Lagrangian $\mathcal{L}_t$ is formulated as:

$$\mathcal{L}_t = (1 + \lambda_t) \mathbb{E}_t [\Lambda_{t,t+1}^S \Omega_{t+1} N_{FI,t+1}] - \lambda_t \theta (A_{H,t} + S_t A_{FX,t}) \quad (\text{C.3.3})$$

where λ_t is the Lagrange multiplier on the incentive constraint, and $\Omega_{t+1} \equiv (1 - \sigma) + \sigma \nu_{t+1}$ is the shadow value of internal funds.

3. First-Order Conditions. We differentiate the Lagrangian with respect to each asset class.

Step 3a: Domestic Assets ($A_{H,t}$). Differentiating with respect to $A_{H,t}$, noting that $\frac{\partial A_{Total,t}}{\partial A_{H,t}} = 1$:

$$\frac{\partial \mathcal{L}_t}{\partial A_{H,t}} = (1 + \lambda_t) \mathbb{E}_t \left[\Lambda_{t,t+1}^S \Omega_{t+1} \left(R_{H,t+1} - \frac{R_{D,t}}{\pi_{t+1}} - \psi A_{Total,t} \right) \right] - \lambda_t \theta = 0 \quad (\text{C.3.4})$$

Rearranging terms yields the FOC for domestic assets:

$$\mathbb{E}_t \left[\Lambda_{t,t+1}^S \Omega_{t+1} \left(R_{H,t+1} - \frac{R_{D,t}}{\pi_{t+1}} - \psi A_{Total,t} \right) \right] = \frac{\lambda_t}{1 + \lambda_t} \theta \quad (\text{C.3.5})$$

Step 3b: Foreign Assets ($A_{FX,t}$). Differentiating with respect to $A_{FX,t}$ requires applying the chain rule, as the foreign asset position enters the balance sheet multiplied by the nominal exchange rate S_t . Note that $\frac{\partial A_{Total,t}}{\partial A_{FX,t}} = S_t$.

$$\begin{aligned} \frac{\partial \mathcal{L}_t}{\partial A_{FX,t}} &= (1 + \lambda_t) \mathbb{E}_t \left[\Lambda_{t,t+1}^S \Omega_{t+1} \left(S_t R_{FX,t+1} - S_t \frac{R_{D,t}}{\pi_{t+1}} - \psi A_{Total,t} S_t \right) \right] - \lambda_t \theta S_t \\ &= S_t \left\{ (1 + \lambda_t) \mathbb{E}_t \left[\Lambda_{t,t+1}^S \Omega_{t+1} \left(R_{FX,t+1} - \frac{R_{D,t}}{\pi_{t+1}} - \psi A_{Total,t} \right) \right] - \lambda_t \theta \right\} = 0 \quad (\text{C.3.6}) \end{aligned}$$

Since $S_t > 0$, we can divide the entire equation by S_t . The exchange rate level cancels out, yielding the FOC for foreign assets:

$$\mathbb{E}_t \left[\Lambda_{t,t+1}^S \Omega_{t+1} \left(R_{FX,t+1} - \frac{R_{D,t}}{\pi_{t+1}} - \psi A_{Total,t} \right) \right] = \frac{\lambda_t}{1 + \lambda_t} \theta \quad (\text{C.3.7})$$

4. Implied Conditions. Comparing (C.3.5) and (C.3.7), we derive the no-arbitrage condition (UIP) by equating the left-hand sides:

$$\mathbb{E}_t \left[\Lambda_{t,t+1}^S \Omega_{t+1} R_{H,t+1} \right] = \mathbb{E}_t \left[\Lambda_{t,t+1}^S \Omega_{t+1} R_{FX,t+1} \right] \quad (\text{C.3.8})$$

Simultaneously, summing the conditions determines the optimal aggregate leverage of the intermediary.