ABSTRACT

Title of dissertation: ESSAYS ON INTERNATIONAL FINANCE

Giuliano Ariel Simoncelli
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Dissertation directed by: Professor Pierre De Leo
Department of Economics

Professor Pablo Ottonello
Department of Economics

This dissertation studies how monetary policy, international capital flows, and fiscal financing constraints shape bond yields, exchange rates, and inflation in open economies. Across three chapters, it shows that financial prices in these economies are influenced not only by conventional policy and market forces, but also by the information conveyed by central bank actions, the balance-sheet exposures created by cross-border investors, and the fiscal incentives behind inflationary finance.

Chapter 1 examines how monetary policy surprises affect the term structure of interest rates in advanced and emerging economies. Using a novel high-frequency dataset of sovereign nominal yield curves for more than 40 countries, together with information from inflation-indexed bonds, I show that a surprise monetary tightening lowers long-term nominal yields, term premia, and inflation expectations in advanced economies, but raises them in emerging markets. I also show that these contrasting responses are driven by differences in the informational content of policy surprises: in emerging markets, private agents extract relatively more information from unexpected policy decisions about future inflation. To ra-

tionalize these findings, I develop a preferred-habitat model with asymmetric information in which policy actions transmit not only through interest rates, but also through the beliefs they reveal about the economy.

Chapter 2, joint with Pierre De Leo, Lorena Keller, Mauricio Villamizar-Villegas and Tomás Williams, studies how global investors in local-currency bond markets shape the joint dynamics of government bond yields and exchange rates. Using transaction-level data from Colombia's bond and foreign exchange markets, we show that global investors' purchases of local-currency government bonds are systematically accompanied by simultaneous transactions in the spot foreign exchange market, whereas this is not the case for local investors. We develop a model in which these cross-market flows alter intermediaries' exposure to bond and currency risk, thereby affecting the joint behavior of bond and currency risk premia. The model accounts for the observed effects of foreign inflows on intermediaries' positions, asset returns, and the comovement between bond yields and exchange rates, and it helps explain cross-country differences in these joint dynamics.

Chapter 3 studies the optimal use of seigniorage in economies where agents can substitute away from domestic currency and into foreign money. Motivated by the fact that persistent fiscal deficits remain a central feature of high-inflation economies, I solve a multi-currency optimal taxation problem with and without foreign exchange controls. The analysis shows that restricting the use of foreign currency increases the government's control over the monetary base and allows it to sustain larger fiscal deficits through inflationary finance. The chapter also implies that the Friedman rule is never optimal in this environment, that countries with weaker tax systems rely more heavily on the inflation tax, and that foreign exchange controls may become more desirable when foreign nominal interest rates are high.

ESSAYS ON INTERNATIONAL FINANCE

by

Giuliano Ariel Simoncelli

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Advisory Committee:
Professor Pierre De Leo, Co-Chair
Professor Pablo Ottonello, Co-Chair
Professor William Witheridge
Professor John Shea
Professor Russell Wermers

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Dedication

A mis padres, Carolina y Enrique, a mi hermana, Gimena, y a mis abuelos: Ariel, Gloria, Alejandro y Raquel.

Gracias por acompañarme siempre, los amo.

Acknowledgments

It has been an amazing ride! Before moving to Maryland, I heard many mixed reviews about pursuing a Ph.D. in economics — often negative ones. I can honestly say, however, that I have enjoyed every single day of this experience, and I would gladly do it all over again. As I write these lines, I cannot help but reflect on how far I have come. Becoming the first Ph.D. in my family, and the first from my small city of Mercedes, fills me with immense pride. A dissertation is ultimately a collective achievement, and what follows is an extensive list of the people who have influenced me throughout my life.

This journey would not have been possible without the support of the many wonderful people I met along the way. First and foremost, I am deeply grateful to my advisors, Pierre De Leo, Pablo Ottonello, and William Witheridge, for their unwavering support, mentorship, and guidance throughout my Ph.D.

I owe an immense debt of gratitude to Pierre. He has been extraordinarily generous with his time and attention. Before starting my Ph.D., I often heard that the relationship between a student and an advisor can be complicated, and that academia can sometimes be a difficult environment to navigate. From the very first conversation we had, however, Pierre made me feel welcomed and supported. I am grateful to leave the program not only with an outstanding mentor, but also with a valued colleague and friend. He has helped me grow tremendously as a researcher and, more importantly, as a person. I wish him nothing but the best in the future.

Pablo provided invaluable feedback and advice throughout my time in the program. His high standards and professionalism constantly pushed me to become a more careful and effective researcher, and this dissertation is much stronger because of his engagement and guidance. His contributions to our research group have been exceptional, and he has become a role model for the profession. Will has been a wonderful addition to the department.

He consistently offered insightful comments, encouragement, and great companionship. I have no doubt that he will leave a significant mark on the field, and I am very grateful for his support, generosity, and positive energy.

I am also grateful to other faculty members at the University of Maryland who contributed to my development as an economist through their support, guidance, and encouragement, including Borağan Aruoba, Luminita Stevens, and Thomas Drechsel. To Martina Copelman, thank you very much for always being so nice and supportive. I loved being your TA.

Second, one of my biggest influences over the past decade has been my mentor, friend, and former boss, Guillermo Tolosa. Since we began working together in 2017, Guillermo has always been there for me. He is one of the main reasons I decided to pursue a doctoral degree. I have long been inspired by his knowledge, his curiosity, and the work he does. I hope that one day I can become even half the economist he is.

I am also deeply grateful to Gonzalo Varela for believing in me and giving me the opportunity to work and grow by his side. I have learned a great deal from him over the years, and I hope that one day we become neighbors in San Francisco (Uruguay).

Third, there are many other professors who have left a lasting mark on me along the way. I would not be here without the immeasurable support of the faculty at Universidad de Montevideo. I want to express my deepest admiration and gratitude to Ana Balsa, Marcelo Caffera, Alejandro Cid, and Juan Dubra. Their help, encouragement, and enthusiasm since my undergraduate years have meant a great deal to me. To Juan, and I speak on behalf of many young economists who have been guided by him, I can only express my deepest admiration. What he has done for the development of our profession in Uruguay is truly remarkable. In many ways, he has become one of the main international references for Uruguayan economics. Much like Diego Forlán or Luis Suárez have put Uruguay on the map in world football, Juan has done something similar for economics: it is not uncommon

for economists abroad to hear that someone is from Uruguay and immediately say, “Oh, say hi to Juan Dubra for me!”

Fourth, I feel incredibly fortunate to have shared my Ph.D. experience with such wonderful colleagues, all of whom I hope will remain lifelong friends. I could not have gone through this journey without them. In particular, I would like to thank Matilde Angarita, Tom Boesche, José Ignacio Cristi, Siqin Ding, Rolando Hernández, Kanat Isakov, Luis F. Jaramillo, Martim Leitão, Camilo Marchesini, Aditya Pande, Mariana Sans, Daniel Schwindt, Trí Phú Vũ, and Lixi Wang.

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Now, as I move to more personal acknowledgments, I will continue in Spanish. I hope the reader will bear with me.

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cuál no los haya sentido al lado mío. Desde chico siempre se han asegurado de que no me falte nada, y voy a necesitar cuatro vidas para poder devolverles todo lo que han hecho por mí. Esto también es de ustedes.

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List of Abbreviations

AE	Advanced Economy
AR	Autoregressive Process
ATSM	Affine Term Structure Model
BEI	Breakeven Inflation
BIS	Bank for International Settlements
CBC	Central Bank of Colombia
CIP	Covered Interest Parity
COP	Colombian Peso
CS	Currency Substitution
DCV	Central Securities Depository
EM	Emerging Market
FX	Foreign Exchange
FXI	Foreign Exchange Intervention
GDP	Gross Domestic Product
IFS	International Financial Statistics
ILB	Index-Linked Bond
IMF	International Monetary Fund
IMC	Intermediarios del Mercado Cambiario
LC	Local Currency
MEC	Colombian Electronic Market
MP	Monetary Policy
OLS	Ordinary Least Squares
PD	Primary Dealer
PP	Posición Propia

PPC	Posición Propia Contado
QE	Quantitative Easing
RDD	Regression Discontinuity Design
RMSE	Root Mean Squared Error
SEN	Electronic Trading System
SET-ICAP FX	SET-ICAP Foreign Exchange Platform
TES	Títulos de Tesorería
UIP	Uncovered Interest Parity
U.K.	United Kingdom
U.S.	United States
USD	United States Dollar
VAR	Vector Autoregression
WEO	World Economic Outlook

Chapter 1: Information Effects of Monetary Policy and Bond Yields in Emerging Markets

1.1 Introduction

Understanding the transmission of monetary policy lies at the core of monetary economics. Yet, despite decades of theoretical and empirical developments, substantial uncertainty remains about the effects of monetary policy.[1] This paper examines how domestic monetary policy surprises affect the nominal term structure of interest rates in advanced economies (AEs) and emerging markets (EMs). Long-term interest rates anchor key economic decisions (i.e., sovereign borrowing), making their response to short-rate changes a central aspect of policy transmission. The strength and nature of this pass-through ultimately determine the broader effectiveness of monetary policy.

I construct a novel, comprehensive high-frequency dataset of sovereign *nominal* yield curves using local-currency bond-level data for over 40 countries. From these yields, I obtain daily estimates of nominal *term premia* (the extra compensation investors require to hold longer-term bonds instead of planning to roll over short-term bonds).[2] I also compile data on inflation-indexed bonds to estimate *real* yield curves, which allows me to compute *breakeven inflation* rates, a market-based proxy for inflation expectations. Finally, I mea-

[1] See, for example, Bernanke (1990); Leeper et al. (1996); Christiano et al. (1999); Romer and Romer (2004); Gertler and Karadi (2015); Miranda-Agrippino and Ricco (2021).

[2] Across this literature, the estimation of term premia is typically based on affine term structure models, pioneered by Duffie and Kan (1996), which remain the workhorse framework for measuring term premia across countries. For a recent discussion, see Piazzesi (2010).

sure monetary policy surprises as the difference between the policy rate announced by the central bank and the average forecast of *Bloomberg* market analysts prior to each meeting.

To address this question, the paper combines an empirical analysis based on this novel dataset with a theoretical model that provides a mechanism to rationalize the findings. The empirical analysis shows that monetary policy surprises affect the term structure of interest rates in starkly different ways across advanced and emerging economies. It further documents that these surprises convey relatively more information about the macroeconomy in EMs than in AEs, consistent with the view that macroeconomic uncertainty is greater in emerging markets, while central bank communication is relatively less clear.[3] To account for these findings, the paper develops a portfolio-balance model with information asymmetries between a central bank and bond investors that can explain the divergent response of long-term yields to policy surprises.

In emerging markets, I find that a surprise policy rate hike increases the compensation investors demand for holding long-term nominal bonds, leading both long-term yields and term premia to rise. In contrast, the same tightening reduces long rates and term premia in advanced economies, consistent with evidence for the United States documented by Gürkaynak et al. (2005a) and Kekre et al. (2024a).[4] Moreover, higher-than-expected policy rates raise long-run inflation expectations in EMs, while reducing inflation expectations in AEs.

A well-established literature on U.S. monetary policy shows that unexpected changes in the policy rate reflect not only monetary policy *shocks*, which refer to movements in interest rates unrelated to macroeconomic conditions (i.e., inflation), but also an *information* component, reflecting discrepancies in beliefs about the state of the economy between the

[3] Appendix A.2 documents that IMF Article IV Staff Reports for emerging markets frequently recommend that central banks strengthen their communication practices.

[4] Previous research has shown, for instance for the U.S., that unexpected monetary policy tightenings typically lower nominal long-term interest rates, consistent with the view that such actions reduce long-run inflation expectations.

central bank and the private sector (Nakamura and Steinsson, 2018; Miranda-Agrippino and Ricco, 2021).[5]

Empirically, I recover monetary policy shocks -9by disentangling the information effects of monetary policy surprises, and obtain two results. First, the observed increase in long-term risk compensation following surprise tightenings in EMs is largely explained by information effects. For example, when EM central banks raise nominal rates unexpectedly, investors may interpret these actions as signals of higher or more volatile inflation ahead and prompting them to demand additional compensation for holding long-term nominal bonds.[6] In line with this result, García-Schmidt (2024) documents sizable central-bank information effects in Brazil and Chile.[7]

Second, monetary policy surprises convey relatively more information about the macroeconomy in emerging markets than in advanced economies. This result is also consistent with Andritzky et al. (2007), who argue that the less transparent and more uncertain environment in emerging markets amplifies the information value of macroeconomic data releases.

To account for the empirical findings, I develop a model that embeds information frictions into a canonical preferred-habitat framework (Vayanos and Vila, 2021a). I deviate from their model in two ways. First, policy announcements convey information about the path of inflation. The central bank observes inflation directly and sets the short rate via a Taylor-type rule, while private agents receive noisy signals about inflation and therefore partly rely on policy actions when forming beliefs, as in Tang (2015). While higher interest rates exert the standard disinflationary force by reducing aggregate demand, in this setting

[5] Blinder et al. (2008) note that imperfect and asymmetric information between central banks and the public is the rule, not the exception, in monetary policy.

[6] Evidence consistent with this example can be found in Tang (2015), Falck et al. (2021), and Witheridge (2024).

[7] Work by Aruoba et al. (2021) also presents evidence of an information channel in monetary policy announcements for Chile, where expected inflation exhibits a hump-shaped response following unexpected tightenings.

they also serve as signals that may lead agents to revise their inflation outlook upward.

Second, preferred-habitat investors adjust their bond holdings in line with their perceived expectations of inflation (Campbell and Viceira, 2001).[8] When inflation expectations are high, habitat investors reduce their demand for long-term nominal bonds, leaving more of the residual supply to be absorbed by risk-averse arbitrageurs. Arbitrageurs, in turn, demand higher compensation for bearing risk across maturities, resulting in a higher nominal term premium.

In the model, a surprise policy tightening may increase nominal term premia when private agents have less precise information about the inflation rate. The mechanism rests on the assumption that private agents in emerging markets possess noisier information than those in advanced economies and therefore extract relatively more information from monetary policy decisions. As a result, policy surprises in EMs may trigger adverse market reactions if investors revise inflation expectations upward. I test this assumption empirically by examining the dispersion of inflation forecasts from market analysts, finding that disagreement is substantially higher in EMs than in AEs. This greater dispersion reflects noisier private information and supports the idea that monetary policy announcements in emerging markets convey valuable information about future inflation.

These findings carry important implications for monetary policy design in emerging markets. First, they highlight that monetary policy surprises can have unwanted effects on sovereign borrowing costs. The effectiveness and costs of monetary policy thus depend on the information environment in which it is conducted. Second, strengthening communication frameworks (i.e., publishing clearer forecasts, articulating the central bank's assessment of risks, and improving the transparency of policy deliberations) can help narrow the informational gap between the central bank and the private sector. Third, the results

[8] This formulation also captures the empirical observation that investors respond to higher perceived expectations of inflation by reducing bond holdings (Ceballos and Ng, 2024).

challenge the prevailing view that emerging-market borrowing conditions are driven predominantly by global financial forces, showing instead that domestic monetary policy has substantial effects on the entire yield curve.

Related Literature This paper contributes to several strands of literature.

First, I contribute to the literature on the transmission of domestic monetary policy in emerging markets (Obstfeld, 2015).[9] While early evidence by Kaminsky et al. (2005) documented that monetary policy in EMs tends to be procyclical, more recent research shows that many EMs have increasingly "tamed the cycle". McGettigan et al. (2013) argue that the maturation of inflation-targeting regimes, together with deeper financial markets and more flexible exchange rates, has enabled EM central banks to conduct more countercyclical policy. Similarly, Vegh and Vuletin (2014) document that EMs have largely overcome the fear of large exchange-rate movements during crises. Consistent with these developments, De Leo et al. (2024a) find that EM central banks follow Taylor-type rules, adjusting policy rates countercyclically and responding systematically to inflation fluctuations.

Second, this study is related to the literature on interest rate surprises.[10] Existing work for advanced economies has emphasized that these surprises contain *information effects*; in which policy actions reveal the central bank's private assessment of the economy.[11] More recent contributions have sought to disentangle different forces within high-frequency sur-

[9]The literature studying how monetary policy transmits to financial markets and the broader economy in AEs is vast. An incomplete list of papers that examine this transmission in the U.S. includes Cook and Hahn (1989), Kuttner (2001), Cochrane and Piazzesi (2002), Gürkaynak et al. (2005b), Hanson and Stein (2015), Nakamura and Steinsson (2018), Swanson (2021a), and Kekre et al. (2024a). For the euro area, see for example Altavilla et al. (2019) and Andrade and Ferroni (2021).

[10]Early high-frequency strategies that used narrow event windows around U.S. policy announcements to capture unexpected short-rate movements (Cook and Hahn, 1989; Kuttner, 2001)

[11]Romer and Romer (2000) argue that shifts in the federal funds rate signal the Fed's private information about inflation, and Nakamura and Steinsson (2018) show that contractionary surprises raise long-term yields and output forecasts, consistent with the Fed signaling stronger fundamentals.

prises.[12] While much of this debate has focused on advanced economies, identifying monetary policy shocks and studying their impact on the macroeconomy in emerging markets presents additional challenges. Thin financial markets and the limited availability of high-frequency data make it difficult to separate pure policy shocks from information effects in EMs. Several recent studies attempt to construct monetary policy shocks for emerging markets.[13]

Third, this is the first attempt to bridge two strands of research that have so far developed largely in parallel: the theoretical literature on monetary policy with imperfect information and the preferred-habitat framework of the term structure. By embedding information frictions into a model with segmented bond markets, I highlight how central bank actions can transmit to bond yields not only through interest rate changes but also through the beliefs they signal about the macroeconomy.

On the one hand, a handful of empirical studies have shown that surprise monetary tightenings may, under certain conditions, lead private agents to raise rather than lower their inflation forecasts (Romer and Romer, 2000; Tang, 2015; Falck et al., 2021; Aruoba et al., 2021). These findings highlight the role of information effects and the limits of the

[12] Miranda-Agrippino and Ricco (2021) recover monetary policy shocks using the Fed's internal forecasts, while Cieslak and Schrimpf (2019) exploit joint stock–bond dynamics to separate monetary from non-monetary news. Meanwhile, the literature has debated whether these patterns reflect genuine information effects or instead predictable responses of the central bank to public news. Bauer and Swanson (2023a,b) argue that much of what appears as a "central bank information effect" is better interpreted as a "response-to-news" channel, whereby the central bank and private agents react simultaneously to the same economic developments. In contrast, Jarociński and Karadi (2025) propose a three-way decomposition of monetary policy surprises into exogenous policy shocks, central bank information shocks, and Fed responses to news, and find robust evidence supporting the presence of information effects.

[13] Checo et al. (2024) construct monetary policy shocks from survey expectations of policy rate decisions. Brandao-Marques et al. (2020) and Deb et al. (2023) use Taylor-rule residuals or quarterly forecast errors of analysts' expectations measured months in advance. Meanwhile, Bolhuis et al. (2024) build a cross-country dataset leaning on daily changes in interest rate swap rates around monetary policy announcements.

full-information benchmark.[14] Forecast revisions are also often sluggish in response to policy shocks (Andolfatto et al. 2002; Coibion and Gorodnichenko 2012, 2015; Angeletos et al. 2021; Cole et al. 2024), consistent with models of imperfect or dispersed information. The foundational work of Morris and Shin (2002) formalizes how public signals from the monetary authority are processed in economies with heterogeneous information. Subsequent theoretical contributions emphasize the signaling role of monetary policy: Melosi (2017) shows how policy actions transmit information about fundamentals, while Falck et al. (2021) and Tanaka (2025) examine how uncertainty and dispersed beliefs affect the expectations channel of monetary policy.[15] Overall, this line of research supports the view that central bank actions and communications serve not only as policy instruments but also as signals that shape beliefs about future economic conditions.[16] The literature studying information effects in emerging markets, however, remains scarce. Aruoba et al. (2021) and García-Schmidt (2024) examine the response of output and inflation to unexpected policy changes in Chile and Brazil and conclude that their findings are consistent with central bank information effects. Mine is the first study to analyze the role of information effects in the transmission of monetary policy to the entire term structure of interest rates, and to do so across a broad set of emerging economies.

On the other hand, this study is related to the theoretical literature pioneered by Vayanos and Vila (2021a) that models the effects of segmented bond markets and preferred-habitat investors on the term structure of interest rates. This research demonstrates how shifts in

[14] Falck et al. (2021) develops a regime-switching dispersed information New Keynesian model and shows that a contractionary monetary policy shock increases inflation expectations when (1) information is very noisy or (2) the Taylor rule coefficient on inflation is low. In the former case, the public signal is less informative and the signaling channel of monetary policy dominates (Melosi, 2017), while in the latter, monetary policy is perceived as more accommodative (Clarida et al., 1999).

[15] Related work on imperfect information further shows that policy conduct should be more hawkish under imperfect knowledge to anchor expectations and avoid episodes of stagflation (Orphanides and Williams, 2004).

[16] In related work, Candian et al. (2023) show that information effects also arise in exchange-rate policy: foreign-exchange interventions influence expectations by revealing or concealing the central bank's information about fundamentals.

the demand for specific maturities, driven by institutional investors or market segmentation, can generate variation in risk premia and yield curve dynamics.[17] Specifically, Vayanos and Vila (2021a) show that an unexpected increase in the short rate lowers the term premium through a revaluation of risk, while Kekre et al. (2024a) break this comovement by endogenizing arbitrageur wealth. Within their framework, when arbitrageurs have positive duration, an unexpected increase in the short rate may increase the term premium through a revaluation of their wealth. Most models in this tradition focus on index-linked bonds, leaving little scope to analyze inflation dynamics.[18] Work that explicitly incorporates nominal bonds and studies inflation risk remains scarce.[19]

Outline The rest of the paper is organized as follows. Section 1.2 describes the financial data used, the construction of yield curves, and the estimation of term premia. It then analyzes the response of the term structure of interest rates to monetary policy surprises. Section 1.3 describes the methodology used to recover monetary policy shocks and studies their effect on term premia. Section 1.4 develops the preferred-habitat model with asymmetric information, discusses its main implications, and links it to the empirical findings of the previous sections. Section 1.5 tests empirically the main mechanism of the model. Lastly, section 1.6 discusses policy implications and concludes.

[17] See Campbell and Viceira (2002) for earlier efforts to model portfolio choice with multiple assets. This framework is extended in many directions. For example, Greenwood and Vayanos (2014a) allow for the supply risk factor to affect short rates, while Greenwood et al. (2018) introduce a defaultable long-term bond. Alternatively, Ray et al. (2019) embeds this framework into a New Keynesian model to study the effectiveness of unconventional monetary policy during financial crises.

[18] Further, studies with multiple currencies include Gourinchas et al. (2022a), Greenwood et al. (2023b), and De Leo et al. (2025).

[19] Lengyel (2022) includes both nominal and index-linked bonds to study the impact of supply shocks on inflation, while Diez de los Rios (2025) embeds a Taylor-type rule into a portfolio-balance model to investigate the relationship between bond supply and inflation. See also Ang et al. (2007) for an earlier effort to embed a Taylor-type rule into a no-arbitrage term structure framework.

1.2 Bond Yields, Term Premia, and Monetary Policy Surprises

This section details the financial data sources and empirical methodologies employed in the study. I then present one of the main contributions of this paper: bond yields in emerging markets respond differently to domestic policy surprises than those in advanced economies.

1.2.1 Constructing Nominal Zero-Coupon Yield Curves and Term Premia

To obtain daily information on the term structure of nominal interest rates, I proceed in three steps. First, I compile daily sovereign local-currency bond data for 42 countries from *Bloomberg* and *Refinitiv* over 2010–22, restricting the sample to bonds with up to 10 years to maturity to enhance comparability and mitigate liquidity concerns (Kekre et al., 2024a).[20] Second, I estimate daily zero-coupon yield curves using standard parametric methods (Nelson and Siegel, 1987; Svensson, 1994), which provide a flexible fit across maturities and ensure that yield curves are measured on a comparable basis across countries. Third, I construct daily nominal term premia estimates using the three-step no-arbitrage framework of Adrian et al. (2013), an approach that mitigates small-sample bias concerns (Bauer et al., 2014). This yields a novel, high-frequency dataset of term premia for a large cross-section of countries, covering both advanced and emerging economies. Such coverage makes it possible to directly compare how monetary policy transmits through sovereign bond markets across different environments.

Specifically, I follow Adrian et al. (2013) to decompose each fitted zero-coupon yield into an expectations component and a term-premium component. For each country, I summarize the cross section of yields with a small number of principal components and estimate a Gaussian affine term structure model in which these factors evolve over time and bond prices are affine in the state vector. The model then implies, for each maturity, both

[20] The sample extends back to 2000 for advanced economies.

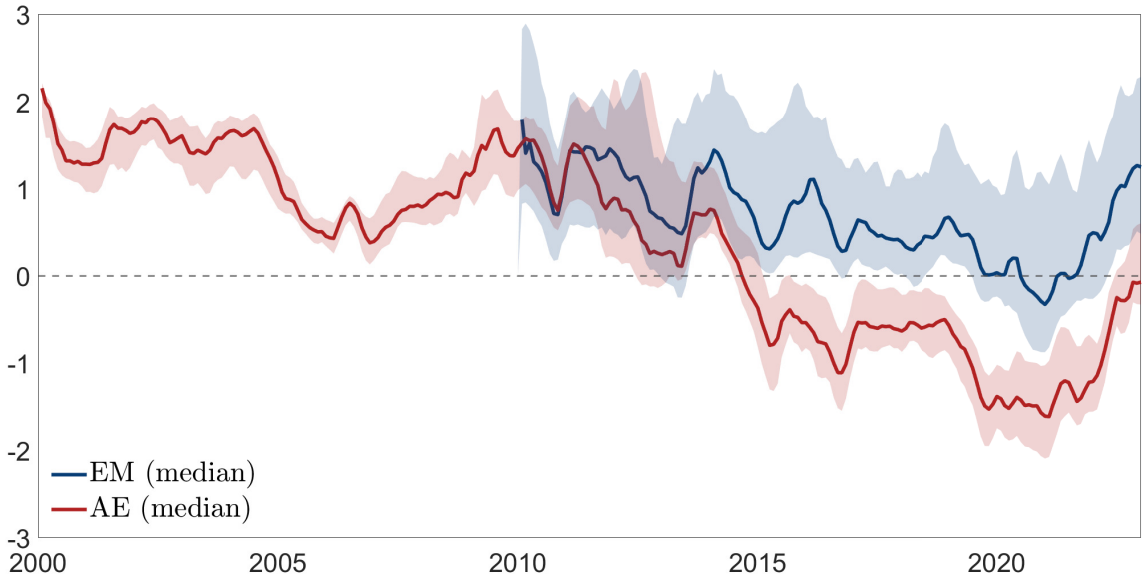
the yield on a zero-coupon bond and the average expected future path of short-term interest rates over the same horizon. The nominal term premium is then measured as the difference between the two. This approach provides a disciplined no-arbitrage decomposition of long-term yields into expectations and risk-compensation components, while remaining sufficiently parsimonious for a broad cross-country application. Appendix A.1 provides the full estimation details.

Figure 1.1 plots the median nominal term premium on a 10-year sovereign bond over time for emerging markets and advanced economies. Several important patterns emerge. In advanced economies, term premia have steadily declined since the early 2000s and have been negative for much of the past decade, consistent with the decline in global inflation risk compensation documented by Wright (2011) and Turner (2014). In contrast, emerging markets exhibit persistently higher and generally positive premia, reflecting greater macroeconomic uncertainty and weaker anchoring of expectations.[21] The shaded interquartile bands highlight that this difference is not driven by a few outliers but reflects a systematic difference between EMs and AEs.

The descriptive evidence presented here highlights a persistent divergence in long-term risk compensation between advanced and emerging economies. A natural question is what role unexpected changes in the monetary policy rate play in driving these differences. Because policy rate decisions are among the most visible signals that central banks send to markets, they provide a useful lens for studying how beliefs and risk compensation are shaped. To quantify the effect of policy actions on bond markets, I next construct measures of monetary policy surprises based on high-frequency data around central bank announcements.

[21] I also estimate nominal term premia on 5- and 7-year bonds, as shown in Figures A.3 and A.4. The same patterns arise.

Figure 1.1: Median Nominal Term Premium on a 10-Year Sovereign Bond, by Country Group



Note: This figure presents the median nominal term premium on a 10-year sovereign bond for emerging markets (blue line) from 2010 to 2022, and advanced economies (red line) from 2000 to 2022. The shaded bands indicate the interquartile range for each group. Term premia are estimated using the three-step no-arbitrage framework of Adrian et al. (2013) based on daily zero-coupon yield curves constructed from sovereign bond prices. See Appendix A.1 for further details on the sample, estimation methods, and country classifications.

1.2.2 Measuring Monetary Policy Surprises

A common strategy to identify exogenous monetary policy changes exploits high-frequency movements in interest rates within narrow windows (often as short as 30 minutes) around central bank announcements.[22] This approach isolates market reactions to policy news and has become the benchmark in the literature. Yet its application in the existing literature is confined to a handful of advanced economies with deep and liquid markets. For most emerging markets, where such instruments are unavailable, high-frequency measures of policy surprises are not available in existing literature. This gap severely limits the scope for cross-country analysis of how monetary policy transmits to bond markets.

To enable systematic cross-country analysis, I construct a dataset of monetary policy

[22] See Bauer and Swanson (2023b) for a review of the literature and methodologies.

rate surprises that can be applied across both advanced and emerging economies, while remaining in line with the related literature.[23] The sample includes 15 emerging markets, 11 advanced economies with independent monetary policy, and 16 advanced economies that are pegged to or have adopted the euro. In this framework, I define a domestic monetary policy surprise as the difference between the policy rate announced by the central bank and the average forecast submitted by market analysts to Bloomberg in the two weeks prior to the meeting. Because Bloomberg allows analysts to incorporate information up to the time of the policy meeting, their forecast errors provide a natural measure of the surprise component of the policy rate decision.[24] Formally,

$$mps_{c,t} = i_{c,t}^{realized} - \frac{1}{N_{c,t}} \sum_{j=1}^{N_{c,t}} f_{j,c,t}, \tag{1.1}$$

where $mps_{c,t}$ denotes the monetary policy surprise for country c at time t, $i_{c,t}^{realized}$ is the policy rate announced on day t for country c, $f_{j,c,t}$ is the forecast of analyst j, and $N_{c,t}$ is the number of forecasts submitted for that policy meeting. In total, the dataset comprises 2,795 monetary policy announcements between 2010 and 2022, with about a quarter of announcements classified as *surprise tightenings* (announced rates above expectations) and one-fifth as *surprise easings* (announced rates below expectations). Table 1.1 summarizes the coverage and distribution of surprises across countries. Monetary policy rate surprises are more frequent in emerging markets than in advanced economies, and their dispersion

[23] There is a large body of literature that uses the difference of announced data and market survey expectations as measure of macroeconomic surprises. See, for example, Cook and Korn (1991); Gürkaynak et al. (2005a); Taylor (2010); Ehrmann et al. (2011).

[24] Unlike national survey sources, the Bloomberg survey allows analysts to submit and update their policy rate forecasts at any time prior to the monetary policy meeting. Submissions can be revised in light of new economic or financial information up to the day of the decision, ensuring that forecasts incorporate the latest available data. Analysts also have strong incentives to be accurate, as their names are visible to all Bloomberg users and the platform regularly ranks forecasters based on performance. Approximately 97% of all forecasts are submitted within the two weeks preceding policy meetings, with submission activity peaking in the final week, indicating that analysts update their expectations very close to the announcement.

is also notably larger. While the average surprise is somewhat lower in emerging markets than in advanced economies, a two-sample test of equality of means fails to reject that the two group averages are the same ($p-$value = 0.398). This partly reflects the effective lower bound, which constrained policy rates in many advanced economies during the 2010s and mechanically reduced the scope for unexpected moves. Furthermore, greater macroeconomic and financial uncertainty in emerging markets makes it harder for analysts to anticipate central bank decisions, leading to larger forecast errors. Finally, weaker institutional credibility and less transparent communication may further limit the ability of market participants to form accurate expectations (Woodford, 2005; Candia et al., 2020).

As with any survey-based identification, this measure has limitations. It captures unexpected movements in current policy rates but abstracts from shocks operating through forward guidance or asset purchases, which have played a central role in advanced economies where policy rates have often been constrained by the effective lower bound. This concern is less severe in emerging markets, where policy rates have generally remained well above zero and conventional rate changes have continued to be the main policy instrument. Checo et al. (2024) show that forecast-based surprises track high-frequency, market-based surprises for the United States quite closely, lending credibility to this approach. While I do not attempt a similar comparison here, their findings suggest that forecast-based surprises provide a reliable proxy in contexts where high-frequency data are unavailable.

1.2.3 Bond Yield Responses to Monetary Policy Surprises

I begin by examining how monetary policy surprises affect the nominal term structure of interest rates in advanced and emerging economies. To do so, I estimate panel regressions separately by country group and for each bond maturity using monthly data from 2010 onward. The dependent variable is the change (in basis points) in zero-coupon sovereign

Table 1.1: Summary Statistics of Monetary Policy Announcements, 2010-22

A. Emerging Markets

	Monetary policy announcements	Surprise tightenings	Surprise easings	Mean (bps)	St. dev. (bps)	Start date
Brazil	104	25	35	-1.3	9.9	Jan-10
Chile	137	38	27	0.8	9.6	Jan-10
Colombia	139	49	38	-0.6	10.9	Jan-10
Hungary	157	27	28	0.7	11.1	Jan-10
India	87	43	20	2.0	10.1	Jan-10
Indonesia	81	27	24	0.4	8.9	May-16
Malaysia	78	24	15	0.2	6.5	Jan-10
Mexico	109	30	19	0.5	8.8	Jan-10
Nigeria	71	24	19	3.9	49.7	May-11
Peru	156	34	29	0.0	7.3	Jan-10
Philippines	54	13	10	-0.1	8.8	Jun-16
Romania	101	18	33	-1.7	8.6	Jan-10
Russia	80	35	29	-5.1	36.4	Sep-13
South Africa	79	27	18	0.0	13.4	Jan-10
Thailand	102	42	16	-0.2	6.6	Jan-10
TOTAL	**1535**	**456**	**360**	**-0.34**	**14.95**	

B. Advanced Economies

	Monetary policy announcements	Surprise tightenings	Surprise easings	Mean (bps)	St. dev. (bps)	Start date
Australia	144	27	26	-0.5	6.6	Jan-10
Canada	104	16	9	0.2	5.3	Jan-10
Czech Republic	104	17	16	0.7	9.8	Jan-10
Euro Area	124	25	8	0.2	4.2	Jan-10
Japan	103	12	0	0.2	0.9	Jan-10
New Zealand	96	16	10	-0.5	6.3	Jan-10
Norway	91	24	11	-0.2	7.4	Jan-10
South Korea	132	42	28	-0.1	7.2	Jan-10
Sweden	75	20	12	0.2	6.8	Jan-10
Taiwan	52	11	13	0.3	4.7	Mar-10
United Kingdom	131	14	7	0.3	2.2	Jan-10
United States	104	10	10	-0.1	2.5	Jan-10
TOTAL	**1260**	**234**	**150**	**0.04**	**5.86**	

Note: Surprise tightenings refer to positive differences between the monetary policy rate announced by the central bank and survey-based expectations. Surprise easings are negative differences. Start date indicates the beginning of the sample.

yields at maturities from 12 to 120 months (1 to 10 years), constructed from end-of-month observations.

To identify the causal effect of unanticipated monetary policy changes, I instrument the monthly change in the policy rate with the high-frequency measure of policy surprises, which isolates the unexpected component of monetary policy decisions.[25] The baseline specification includes country fixed effects and three lags of both the dependent variable

[25] In the rare instances when multiple policy announcements occur within a single month, I average the policy rate surprises to obtain a single monthly measure.

and the policy surprise. Further details on the data and sample construction are provided in Appendix A.1.1. Formally, I estimate the following panel regression specification:

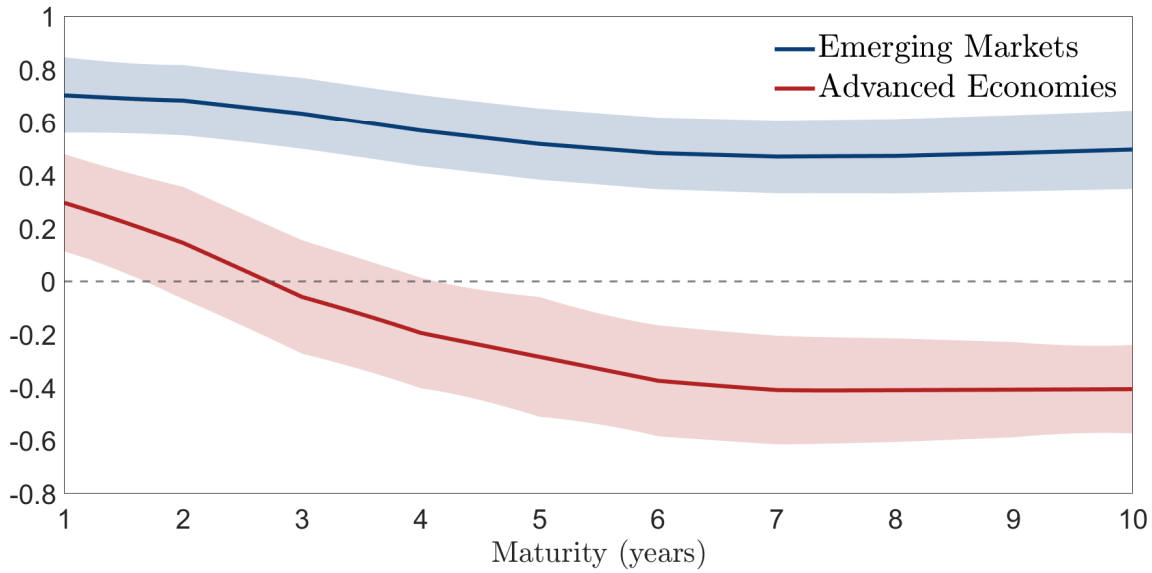
$$y_{c,t}^{N,(\tau)} - y_{c,t-1}^{N,(\tau)} = \alpha_c + \beta \widehat{\Delta i_t^p} + \sum_{k=1}^{3} \rho_k Z_{c,t-k} + \varepsilon_{c,t}, \tag{1.2}$$

where $\tau \in \{12, 13, \ldots, 120\}$ denotes bond maturity. Here, α_c are country fixed effects, $\widehat{\Delta i_t^p}$ is the instrumented unanticipated change in the policy rate, and $Z_{c,t-k}$ includes three lags of the dependent variable and the policy surprise.

The results in Figure 1.2 reveal that the pass-through of monetary policy surprises to long-term yields differs across country groups. For the sample of 27 AEs, a 1 p.p. unanticipated policy rate increase raises the 1-year yield by 0.32 p.p., but *lowers* the 10-year yield by 0.40 p.p., producing a pronounced flattening of the curve. By contrast, the response in emerging markets is fundamentally different. For the sample of 15 EMs, a 1 p.p. surprise shifts the entire yield curve upward: the 1-year yield increases by 0.72 p.p. and the 10-year yield by 0.51 p.p. This indicates that investors in EMs demand substantially greater compensation to hold long-term nominal bonds after an unexpected tightening. To my knowledge, this is the first evidence documenting such systematically divergent yield-curve responses between advanced and emerging economies.

This raises an important question: are the results driven primarily by changes in expectations of future short rates, or by shifts in the compensation investors require to hold long-term bonds? The former would imply that monetary policy surprises mainly alter beliefs about the future path of policy, while the latter would suggest that they reshape perceptions of risk, affecting term premia. To disentangle these channels, I decompose yields into expectations and term premium components using the no-arbitrage framework in Adrian et al. (2013), and estimate the response of term premia to monetary policy surprises in emerging

Figure 1.2: Response of the Yield Curve to a 1 p.p. Surprise Tightening



Note: This figure presents the estimated response of zero-coupon yields at different maturities to a 1 percentage point unanticipated increase in the domestic policy rate, separately for advanced economies (red line) and emerging markets (blue line). Shaded areas denote 90% confidence intervals. The sample is restricted to bonds with up to 10 years to maturity to minimize liquidity premia effects. The sample period begins in 2010. See Appendix A.1 for details on sample construction, methodology, and the list of countries included.

and advanced economies:

$$TP_{c,t}^{(\tau)} - TP_{c,t-1}^{(\tau)} = \alpha_c + \beta \widehat{\Delta i_t^p} + \sum_{k=1}^{3} \rho_k Z_{c,t-k} + \varepsilon_{c,t}, \tag{1.3}$$

for $\tau \in \{60, 84, 120\}$ months (i.e., 5-, 7-, and 10-year maturities). The dependent variable is the change (in basis points) in the estimated term premium at each maturity. The specification mirrors that of the yield regressions in equation (1.2).

The results in Table 1.2 reveal a striking divergence in the drivers of yield-curve responses. In advanced economies, a 1 p.p. surprise tightening reduces nominal term premia by about 40 to 60 basis points across the 5- to 10-year maturities. Because long-term yields can be decomposed into expected future short rates and term premia, this implies that the decline in long-term yields after a policy surprise is largely explained by falling

term premia, consistent with stronger credibility and lower perceived inflation risk. In emerging markets, however, the same 1 p.p. policy surprise appears to raise both expected future short rates and term premia, with the latter increasing by roughly 30 basis points across the 5- to 10-year maturities.[26] In other words, the compensation investors require to hold long-term local-currency bonds rises after a surprise tightening, suggesting that unanticipated short-rate hikes heighten rather than reduce perceptions of risk. As shown in Appendix A.1.7, the results are robust to a variety of alternative specifications, including controlling for global shocks and measures of risk, different sample periods, and different lag lengths. Furthermore, Appendix A.1.4 complements the monthly panel analysis with a high-frequency event-study approach.[27]

Table 1.2: Response of Term Premia to Surprise Tightenings Across Country Groups

	Emerging Markets			Advanced Economies		
	5y	*7y*	*10y*	*5y*	*7y*	*10y*
Surprise tightening	0.316^{***}	0.294^{***}	0.274^{***}	-0.437^{***}	-0.519^{***}	-0.590^{***}
(1 p.p.)	(0.056)	(0.066)	(0.104)	(0.053)	(0.056)	(0.059)
Observations	1930	1930	1930	4023	4023	4023
Countries	15	15	15	27	27	27
R^2	0.72	0.72	0.60	0.92	0.94	0.95

Note: This table presents the estimated response of nominal term premia at 5-, 7-, and 10-year maturities to a 1 percentage point surprise monetary policy tightening, for emerging markets and advanced economies. The nominal term premium is defined as the excess yield investors require to hold sovereign long-term nominal bonds relative to short-term ones. All specifications include country fixed effects and three lags of the dependent variable and the policy surprise. The sample period begins in 2010. Significance stars indicate: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. See Appendix A.1 for further details on data and methodology.

[26] The effect of monetary policy surprises on expected future short rates can be approximated by looking at the difference in results in Figure 1.2 and Table 1.2.

[27] Specifically, I study the two-day change in the 10-year nominal term premium around monetary policy announcements. Analyzing the change in financial variables in narrow windows around policy decisions is a standard approach to isolate the immediate effects of monetary policy surprises from other macroeconomic news. For example, Witheridge (2024) examines the two-day change in market-based inflation expectations (the breakeven inflation rate) around monetary policy announcements.

1.2.4 Index-Linked Bonds and Breakeven Inflation Rates

While the analysis so far focuses on nominal yields, an equally important dimension of monetary policy transmission operates through real interest rates. Index-linked bonds, whose principal and coupon payments are tied to realized inflation, provide a direct window into how investors price future inflation.

To study this margin, I compile data on sovereign local-currency index-linked bonds from *Bloomberg* and *Refinitiv*, starting in 2010 where available. As in the nominal bond sample, the dataset is restricted to securities with a remaining maturity of at most ten years. Due to the limited availability of inflation-linked instruments, zero-coupon yield curves are constructed for a smaller set of countries: six emerging markets and eleven advanced economies.[28] Additional details on the sample and data construction are provided in Appendix A.1.5.

With this data, I construct the breakeven inflation rate (BEI) — the yield spread between a nominal and an inflation-indexed zero-coupon bond of the same maturity — as

$$BEI_t^{(\tau)} = y_t^{N,(\tau)} - y_t^{R,(\tau)}, \tag{1.4}$$

where $y_t^{N,(\tau)}$ and $y_t^{R,(\tau)}$ denote the zero-coupon yields on nominal and index-linked bonds with maturity τ, respectively.[29] These BEI rates provide a market-based measure of expected inflation over the relevant horizon.

I study the response of the breakeven inflation rate on 10-year bonds over time to monetary policy surprises. First, I extract end-of-month real yields, which I use to calculate BEI rates for each country. Second, I estimate dynamic responses of BEI rates using panel

[28] Table A.3 reports descriptive statistics on government debt composition for these countries.

[29] The breakeven inflation rate may also be influenced by liquidity differentials between nominal and inflation-indexed bonds; see d'Amico et al. (2018). I abstract from these considerations in this analysis.

local projections with instrumental variables (Jordà, 2005; Stock and Watson, 2018). The regression specification is:

$$BEI^{10y}_{c,t+h} - BEI^{10y}_{c,t-1} = \alpha^h_c + \beta^h \widehat{\Delta i^p_t} + \sum_{k=1}^{3} \rho^h_k Z_{c,t-k} + \varepsilon^h_{c,t+h}, \tag{1.5}$$

for horizons $h = 0, 1, ..., 8$ months. Here, α^h_c denotes country fixed effects at each horizon, and $\widehat{\Delta i^p_t}$ is the instrumented unanticipated policy rate change. The specification also includes three lags of both the dependent variable and the instrumented policy surprise.

The empirical analysis is conducted for five advanced economies (Australia, France, Italy, the United Kingdom, and the United States) and five emerging markets (Brazil, Chile, Colombia, Mexico, and South Africa). Among advanced economies, I restrict the sample to countries where index-linked bonds account for at least 5% of total government debt securities. For emerging markets, I focus on countries where index-linked bond data are both available and reliable.[30]

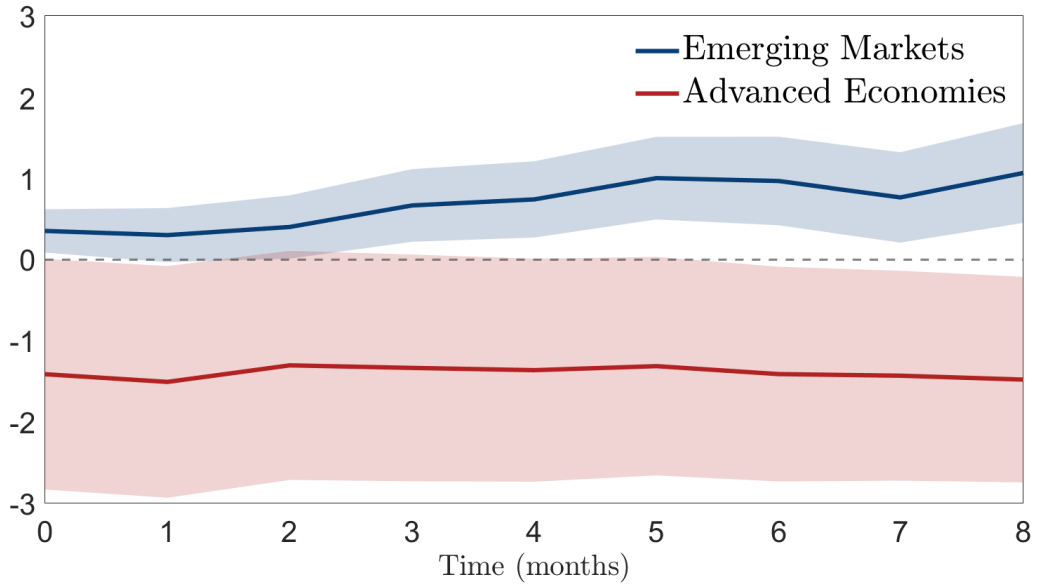
The evidence in Figure 1.3 suggests that 10-year breakeven inflation rates in EMs tend to rise following surprise tightenings, and the increase seems to build gradually over subsequent months.[31] On the contrary, market-based inflation expectations decrease following the same surprise tightening in AEs.

Taking stock The evidence in this section shows that monetary policy surprises have very different implications in advanced and emerging economies. In advanced economies, surprise policy tightenings flatten the yield curve by lowering long-term yields, compressing nominal term premia, and reducing market-based long-run inflation expectations. Investors appear more willing to absorb long-term nominal bonds after observing a higher-than-

[30] Accordingly, six advanced economies are dropped from the sample due to low index-linked bond issuance, whereas Türkiye is excluded due to lack of data.

[31] This pattern is consistent with Witheridge (2024), who documents similar responses on impact at shorter bond maturities.

Figure 1.3: Response of the 10-Y BEI to a 1 p.p. Surprise Tightening



Note: This figure presents the estimated response of the 10-year breakeven inflation rate to a 1 percentage point monetary policy surprise, separately for advanced economies (red line) and emerging markets (blue line). Shaded areas denote 90% confidence intervals. The sample is restricted to bonds with up to 10 years to maturity to minimize liquidity premia effects. The sample period begins in 2010. The sample consists of five advanced economies (Australia, France, Italy, the United Kingdom, and the United States) and five emerging markets (Brazil, Chile, Colombia, Mexico, and South Africa).

expected policy rate set by the monetary authority.

In emerging markets, by contrast, a surprise tightening shifts the entire yield curve upward, significantly increasing nominal term premia and inflation expectations. In this environment, an unanticipated policy rate increase makes long-term nominal bonds less attractive.

In the next section, I take a step further and show that monetary policy surprises can be decomposed into two components: a monetary policy shock — the change in the policy rate orthogonal to macroeconomic fundamentals — and an information component that reflects differences between the central bank's and the public's assessment of the economic outlook. Once I disentangle the information about the macroeconomy embedded in policy surprises, unexpected rate hikes no longer raise term premia in EMs. By contrast, using

monetary policy surprises or shocks yield virtually identical results in advanced economies. These patterns are consistent with policy surprises in emerging markets carrying relatively greater informational content — a dimension of monetary transmission that remains largely understudied.

1.3 The Informational Content of Monetary Policy Surprises

Monetary policy surprises, while unanticipated relative to market forecasts at the time of the policy meeting, are often systematically related ex-post to macroeconomic and financial variables observed by the econometrician. The literature analyzing the anatomy of policy surprises has expanded rapidly, especially for the United States, offering multiple interpretations of what this systematic component represents (Bauer and Swanson, 2023a; Jarociński and Karadi, 2025). While differing in emphasis, these studies share the premise that monetary policy announcements are informative events that lead markets to update their beliefs about the economy.

Much of the progress that has been made in studying this informational component in advanced economies has generally not been possible in emerging markets. This is mostly due to the fact that standard identification strategies rely on high-frequency changes in interest rates within narrow windows around policy announcements. Next, I describe the orthogonalization procedure used to separate monetary policy shocks from information effects in emerging markets and then examine how these shocks affect term premia across country groups.

1.3.1 Monetary Policy Shocks and the Information Channel

The approach presented in this paper builds on Checo et al. (2024) and involves orthogonalizing high-frequency monetary policy rate surprises with respect to a broad set of macroe-

conomic and financial variables.[32] This procedure allows one to obtain a series of monetary policy *shocks* — movements in interest rates unrelated to macroeconomic conditions. I apply this methodology consistently across both emerging and advanced economies to ensure comparability.

Specifically, I run country-by-country regressions of monetary policy surprises on a broad set of both backward-looking and forward-looking macroeconomic and financial variables:

$$mps_t^c = \alpha^c + \beta^c X_t^c + \varepsilon_t^c, \tag{1.6}$$

where mps_t^c is the policy surprise for country c at time t (as defined in equation (1.1)), α^c is a constant, and X_t^c is a vector of macroeconomic and financial variables. The vector includes four price indicators (headline inflation, one-year-ahead inflation expectations, the inflation rate of imported commodities (Gruss and Kebhaj, 2019), and nominal wage growth); three activity indicators (the growth rate and one-year-ahead expected growth rate of industrial production, and the unemployment rate); and three financial market variables (the change and one-year-ahead change in the exchange rate, and the change in stock-market prices).[33] Despite the inclusion of multiple regressors, only those that prove statistically significant are retained.

The information channel The construction of monetary policy shocks using (1.6) rests on the assumption that the econometrician observes a broader set of macroeconomic and financial variables than those on which market participants fully condition when forming expectations.[34] In other words, the variables included in the orthogonalization pro-

[32] I am very grateful to the authors for sharing their dataset of monetary policy shocks for emerging markets.

[33] Following Checo et al. (2024), I omit exchange-rate crises by removing 24 months of data before and 12 months after depreciations that exceed 50 percent within a year—Russia in January 2015 and Egypt in December 2016. See Appendix A.1.6 for a description of data sources.

[34] This is in line with Bianchi et al. (2022). The authors show that survey respondents systematically overweight the idiosyncratic, judgment-based component of their forecasts and under-utilize publicly available macroeconomic information, leading to inefficient expectation formation.

cedure proxy for information that is observable ex post but not necessarily incorporated into agents' forecasts at the time of the policy meeting. This assumption is consistent with the presence of informational frictions that limit agents' ability to observe or process macroeconomic conditions in real time. Such frictions may stem from uncertainty about the economic environment, limited transparency in central bank communication, or sizeable revisions to macroeconomic data releases.[35] Under this view, the orthogonalization removes the portion of policy surprises correlated with macroeconomic fundamentals, thereby removing the information component and leaving a residual that captures the purely unexpected component of monetary policy actions.

Importantly, this empirical procedure provides a tractable proxy for the information channel: it captures the idea that monetary policy announcements convey information about the central bank's private assessment of economic conditions. In this sense, the difference between the original surprises and the residuals in (1.6) reflects the *informational content of policy decisions* — that is, the discrepancies between the central bank's and the private sector's beliefs about the state of the economy.

Under the assumption outlined above, ε_t^c represents the monetary policy surprise cleansed of information effects for country c, and I refer to these residuals as monetary policy *shocks*.

1.3.2 Effect of Monetary Policy Shocks on Term Premia

I now return to the specifications in (1.3), but instrument the change in the policy rate with the monetary policy shocks.

The estimates in Table 1.3 provide two central insights into the nature of monetary

[35] All three phenomena have been extensively documented in the literature. For example, Fraga et al. (2003); Ha et al. (2019) show that inflation volatility is pervasive in emerging markets and significantly higher than in advanced economies. Similarly, Blinder et al. (2008); Dincer et al. (2022) find that central banks in emerging markets communicate less frequently and less transparently. Finally, the literature on data revisions is extensive, emphasizing that revisions are economically meaningful (Gilbert, 2011) and tend to be larger and more consequential in emerging markets (Ley and Misch, 2014).

policy surprises and their effect on term premia across country groups. First, panel (a) shows that while policy surprises have a significant positive effect on nominal term premia in emerging markets, monetary policy shocks do not. This pattern suggests that the rise in long-term risk compensation following surprise tightenings largely reflects the information that markets infer from the policy decision rather than a pure policy shock.[36]

Second, comparing panels (a) and (b) reveals that the informational content of policy surprises is substantially greater in emerging markets than in advanced economies. In advanced economies, the responses to monetary policy surprises and monetary policy shocks are nearly identical, implying that policy announcements contain little additional information beyond the purely unexpected change in the policy rate. In contrast, in emerging markets, the gap between the two estimates reflects the higher informational value of policy decisions, consistent with a less transparent and more uncertain macroeconomic environment. This pattern aligns with the identification strategy discussed in Section 1.3.1, where controlling for a broad set of macroeconomic and financial variables removes the component of policy surprises that is systematically related to observable macroeconomic fundamentals and therefore likely reflects information effects.

In the next section I incorporate these information frictions into a segmented-markets framework that organizes these empirical findings into a single mechanism. In the model, policy surprises influence what investors believe about the path of inflation, and those belief revisions translate into changes in risk compensation through portfolio rebalancing between habitat investors and arbitrageurs. This framework naturally generates the cross-country pattern in Table 1.3: when the information environment is less transparent (as in many EMs) surprise tightenings can raise term premia as markets infer the economic outlook partly from central bank actions. Conversely, when the baseline information that markets

[36]Results for bond yields at different maturities are similar in emerging markets when using either surprises or shocks (see Figure A.8). These findings are consistent with policy surprises conveying differential information that affects risk compensation rather than expected short rates.

possess is more precise, surprise tightenings have little additional informational content, and term premia fall.

Table 1.3: Response of Nominal Term Premia to Interest Rate Changes

Panel (a): Emerging Markets

	Nominal term premium			Nominal term premium		
	5y	*7y*	*10y*	*5y*	*7y*	*10y*
Int. rate surprises	0.316^{***}	0.294^{***}	0.274^{***}			
(1 p.p.)	(0.056)	(0.066)	(0.104)			
MP shocks				0.087	0.080	-0.043
(1 p.p.)				(0.057)	(0.069)	(0.110)
Observations	1930	1930	1930	1930	1930	1930
Countries	15	15	15	15	15	15
R^2	0.72	0.72	0.60	0.74	0.73	0.61

Panel (b): Advanced Economies

	Nominal term premium			Nominal term premium		
	5y	*7y*	*10y*	*5y*	*7y*	*10y*
Int. rate surprises	-0.437^{***}	-0.519^{***}	-0.590^{***}			
(1 p.p.)	(0.053)	(0.056)	(0.059)			
MP shocks				-0.452^{***}	-0.537^{***}	-0.609^{***}
(1 p.p.)				(0.057)	(0.060)	(0.063)
Observations	4023	4023	4023	4023	4023	4023
Countries	27	27	27	27	27	27
R^2	0.92	0.94	0.95	0.92	0.94	0.95

Note: This table reports the estimated response of nominal term premia at 5-, 7-, and 10-year maturities in emerging markets (Panel (a)) and advanced economies (Panel (b)) to a 1 percentage point increase in the domestic policy rate. Results are shown separately when using monetary policy surprises (left panel) and monetary policy shocks (right panel) as instruments. All specifications include country fixed effects and three lags of the dependent variable and the instrument. The sample period begins in 2010. Significance stars indicate: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. See Appendix A.1 for details on sample construction and methodology.

1.4 A Preferred-Habitat Model With Asymmetric Information

The model features two interconnected building blocks. First, I introduce a signal-extraction framework in which the central bank sets the policy rate (the short-term interest rate) according to a Taylor-type rule aimed at controlling inflation. Without loss of generality, I

assume that the monetary authority can observe current inflation perfectly. Private agents, in contrast, have inferior information about current inflation: they can observe the policy rate perfectly but receive possibly noisy signals about inflation itself. Within this framework, it can be shown that when these private signals are noisy, an interest rate increase can actually *raise* private agents' expectations of future inflation.

I then embed this asymmetric-information block in the canonical preferred-framework (Greenwood and Vayanos, 2014a; Vayanos and Vila, 2021a), which I modify to include nominal sovereign bonds across maturities. Preferred-habitat investors adjust their nominal bond holdings according to their inferred outlook on inflation, while risk-averse arbitrageurs absorb the residual supply across maturities and maximize over expected real wealth. Consequently, when monetary policy rate surprises increase private agents' inflation expectations, arbitrageurs may demand higher yields to hold nominal bonds.

The integrated model demonstrates that when information frictions are present a surprise tightening can raise expected inflation and, in turn, nominal term premia. This result mirrors the empirical patterns documented for EMs in Sections 1.2 and 1.3. Conversely, with more accurate information about the macroeconomy, a surprise increase in the policy rate reduces expected inflation and lowers the term premium, consistent with the patterns exhibited by AEs.

1.4.1 Model Setup

The model is set in continuous time and features three types of agents: the central bank and two sets of private agents: preferred-habitat investors and risk-averse arbitrageurs.

Policy-rate Setting and the Signal-Extraction Problem I consider a simple signal-extraction environment in which the central bank sets the nominal monetary policy rate

according to a contemporaneous Taylor-type rule:[37]

$$di_t = \phi\,(\pi_t - \bar{\pi})\,dt + \sigma_i\,dB_{i,t}, \tag{1.7}$$

where $\phi > 1$ is the policy reaction coefficient, $\pi_t - \bar{\pi}$ is the deviation of inflation from its unconditional mean, $B_{i,t}$ is a standard Brownian motion, and $\sigma_i > 0$ is a constant.

Inflation is perfectly observed by the central bank and evolves according to the following Ornstein-Uhlenbeck process:

$$d\pi_t = -\kappa_\pi\,(\pi_t - \bar{\pi})\,dt - \theta\,di_t + \sigma_\pi\,dB_{\pi,t}, \tag{1.8}$$

where $\kappa_\pi \in [0,1]$ is the mean reversion rate of inflation, $B_{\pi,t}$ is a standard Brownian motion, and $\sigma_\pi > 0$. I allow for a contemporaneous disinflationary effect caused by a policy rate change, governed by the parameter $\theta \geq 0$. Consequently, the state dynamics become (using equations (1.7) and (1.8))

$$d\pi_t = -(\kappa_\pi + \phi\,\theta)\,(\pi_t - \bar{\pi})\,dt + \sigma_\pi\,dB_{\pi,t} - \theta\,\sigma_i\,dB_{i,t}. \tag{1.9}$$

Private agents observe the interest rate set by the central bank perfectly at each instant, but have imperfect information about the current inflation rate. This specification captures the idea that monetary policy actions reflect contemporaneous information about inflation known to the central bank but not fully observable to the public. Specifically, following

[37] The closest related asymmetric information framework is Tang (2015), who also model a Taylor-type rule with the central bank reacting to current inflation.

Fajgelbaum et al. (2017) and Tanaka (2025), private agents receive signals about inflation:[38]

$$ds_t = (\pi_t - \bar{\pi})\,dt + \sigma_s\,dB_{s,t}, \tag{1.10}$$

where $B_{s,t}$ is a standard Brownian motion, independent of $B_{i,t}$ and $B_{\pi,t}$. Here, σ_s captures the noisiness of the private signals.

Timing within each instant t is as follows. First, the central bank observes the current inflation rate and is subject to shocks given by the increments $dB_{i,t}$ and $dB_{\pi,t}$. It then sets the policy rate according to equation (1.7). Private agents subsequently observe the newly set interest rate and receive their own signals about current inflation. Based on these two sources of information, private agents update their expectations about inflation.

Expectations Formation Since private agents do not observe the current inflation rate directly, they infer it using both the observed monetary policy rate and the noisy private signals described above. This setup gives rise to a classic signal-extraction problem. In continuous time, the optimal estimate of the hidden state (inflation) given observed signals is provided by the Kalman-Bucy filter.

Let $\hat{\pi}_t$ denote private agents' expectation of the current inflation rate based on their information up to instant t, which includes the history of observed monetary policy rates and private signals. The evolution of their inflation expectations is given by:[39]

$$d\hat{\pi}_t = -(\kappa_\pi + \phi\,\theta)\,(\hat{\pi}_t - \bar{\pi})\,dt + K_i\,dI_t + K_s\,dS_t, \tag{1.11}$$

[38] In the original framework of Tanaka (2025), private agents receive a large number J/Δ_j of noisy signals about the hidden state, indexed by $j \in [0, J]$. Aggregating these signals and assuming $J = 1$ yields (1.10).

[39] A full derivation can be found in Appendix A.3.1.

where the policy rate and private signal surprises, respectively, are

$$dI_t \equiv di_t - \phi(\hat{\pi}_t - \bar{\pi})\,dt; \tag{1.12}$$

$$dS_t \equiv ds_t - (\hat{\pi}_t - \bar{\pi})\,dt, \tag{1.13}$$

and satisfy $\mathbb{E}_t[dI_t] = \mathbb{E}_t[dS_t] = 0$, $\text{var}_t[dI_t] = \sigma_i^2\,dt$, and $\text{var}_t[dS_t] = \sigma_s^2\,dt$. The Kalman gains K_i and K_s weight each innovation according to relative informativeness and the volatility of the underlying disturbances. The first term of equation (1.11) captures the mean reversion of inflation, while the second and third terms represent Bayesian updating: private agents adjust expectations whenever the observed policy rate or private signal deviates from what agents would have predicted based on their priors. The optimal weights placed on policy-rate and private signal surprises are summarized in the following lemma.

Lemma 1. *In steady state, the Kalman gains associated with the policy rate and private signal surprises, respectively, are*

$$K_i = \frac{\phi}{\sigma_i^2}\,\frac{-\kappa_\pi + \sqrt{\kappa_\pi^2 + \Lambda\,\sigma_\pi^2}}{\Lambda} - \theta\,; \tag{1.14}$$

$$K_s = \frac{1}{\sigma_s^2}\,\frac{-\kappa_\pi + \sqrt{\kappa_\pi^2 + \Lambda\,\sigma_\pi^2}}{\Lambda}, \tag{1.15}$$

where

$$\Lambda = \frac{\phi^2}{\sigma_i^2} + \frac{1}{\sigma_s^2}.$$

Proof. See Appendix A.3.3. □

Lemma 1 shows that the weight private agents place on each source of information depends on their informativeness. In particular, when private signals are relatively more precise (i.e., when agents have better information about the inflation rate), the Kalman gain K_s is larger, so agents update inflation expectations primarily from these direct signals rather than from

short-rate changes. By contrast, when private signals are noisier, K_s falls and K_i becomes relatively more important, leading agents to rely more heavily on monetary policy rate movements as an indirect signal of inflation. In this case, a surprise tightening may raise agents' inflation expectations. The following proposition formalizes this result. The key question is whether agents interpret a policy tightening as disinflationary news or as a signal of higher inflation ahead.

Proposition 1. *When private signals are very noisy (high σ_s), a surprise policy rate tightening may lead to higher inflation expectations from the private sector. Specifically,*

$$\frac{\partial d\hat{\pi}_t}{\partial dI_t} = K_i, \tag{1.16}$$

where

$$K_i = \frac{\phi}{\sigma_i^2} \frac{-\kappa_\pi + \sqrt{\kappa_\pi^2 + \Lambda \sigma_\pi^2}}{\Lambda} - \theta \tag{1.17}$$

is the Kalman gain on the interest rate signal, and

$$\Lambda = \frac{\phi^2}{\sigma_i^2} + \frac{1}{\sigma_s^2}. \tag{1.18}$$

In particular, a surprise tightening raises private agents' inflation expectations if $K_i > 0$, or

$$\frac{\phi}{\sigma_i^2} \frac{-\kappa_\pi + \sqrt{\kappa_\pi^2 + \Lambda \sigma_\pi^2}}{\Lambda} > \theta. \tag{1.19}$$

Proof. See Appendix A.3.4. □

Proposition 1 highlights that the updating weight on the monetary policy surprise decomposes according to (1.17). The first term represents an informational component, capturing how strongly a monetary policy rate movement reveals news about the inflation outlook. The second term reflects the mechanical disinflation effect of a rate increase. When the

informational component dominates the contemporaneous feedback, the Kalman gain satisfies $K_i > 0$, implying that a surprise tightening leads private agents to revise their inflation expectations upward.

Having characterized how private agents form beliefs about inflation, I now turn to how these beliefs affect bond pricing in segmented markets. I embed this mechanism into a preferred-habitat framework in which bond investors react to unexpected policy rate changes by adjusting their portfolios.

The Sovereign Bond Market At each instant t, there is a continuum of nominal zero-coupon bonds with maturities $\tau \in (0, T]$ supplied by an unmodeled government and traded by preferred-habitat investors and arbitrageurs. Let $P_t^{(\tau)}$ denote the price at time t of a bond maturing at $(t+\tau)$. Each nominal bond pays one unit of currency at maturity. The continuously-compounded yield to maturity (i.e., the spot rate) is denoted by $y_t^{(\tau)}$, which is related to the price via

$$y_t{}^{(\tau)} = -\frac{\log P_t^{(\tau)}}{\tau}. \tag{1.20}$$

Arbitrageurs Risk-averse arbitrageurs allocate their real wealth across the entire spectrum of nominal bonds. They can also invest at the short rate, which gives the instantaneous yield i_t. Their objective is to balance the expected real return against the variance of real wealth changes. Let W_t denote their time-t real wealth, and let $x_t^{(\tau)}$ be the share of wealth invested in a nominal zero-coupon bond of maturity τ at time t. The arbitrageur's budget constraint is:

$$\frac{dW_t}{W_t} = \int_0^T x_t^{(\tau)} \left(\frac{dP_t^{(\tau)}}{P_t^{(\tau)}} - \hat{\pi}_t dt \right) d\tau + \left(1 - \int_0^T x_t^{(\tau)} d\tau \right) (i_t - \hat{\pi}_t) dt. \tag{1.21}$$

The first term captures the real return on nominal bonds (price change minus inflation), while the second term is the real return on wealth allocated to the instantaneous nominal

rate. Arbitrageurs choose their portfolio to maximize a mean-variance objective:

$$\max_{\left\{x_t^{(\tau)}\right\}_{\tau\in(0,T]}} \mathbb{E}_t\left(\frac{dW_t}{W_t}\right) - \frac{\gamma}{2}\operatorname{var}_t\left(\frac{dW_t}{W_t}\right), \tag{1.22}$$

where $\gamma > 0$ is a risk aversion parameter.

Preferred-Habitat Investors The economy features a continuum of preferred-habitat investors indexed by their investment horizon. Each investor is buy-and-hold, consumes only at the end of its own horizon, and invests exclusively in the nominal bond whose maturity matches that horizon.[40] Let $d_t^{(\tau)}$ denote the aggregate demand, relative to arbitrageurs' real wealth $D_t^{(\tau)}$, of the clientele with habitat τ. Following Hamilton and Wu (2012), I assume that this maturity-specific demand decreases when inflation expectations rise above their unconditional mean:

$$d_t^{(\tau)} = \overline{d}_t^{(\tau)} - \alpha(\tau)\left(\hat{\pi}_t - \bar{\pi}\right), \tag{1.23}$$

where $\overline{d}_t^{(\tau)}$ measures the average demand for maturity τ, and $\alpha(\tau)$ is a deterministic function of maturity τ. I assume that $\alpha(\tau)$ has the following properties:

Assumption 1. *The function $\alpha(\tau)$ satisfies*

(i). $\int_0^T \alpha(\tau)d\tau \geq 0$.

(ii). There exists $\tau^\star \in [0,T)$ such that $\alpha(\tau) < 0$ for $\tau < \tau^\star$ and $\alpha(\tau) > 0$ for $\tau > \tau^\star$.

The first part of Assumption 1 rules out declines in total (wealth-scaled) supply net of preferred-habitat investor demand when $\hat{\pi}_t$ increases. The second part allows short-end net supply to decrease witn $\hat{\pi}_t$ even as total net supply weakly increases. Assumption 1 ensures that an increase in $\hat{\pi}_t$ makes the portfolio that arbitrageurs hold in equilibrium more sensitive to interest-rate and inflation risk.

[40] See, for example, Brennan and Xia (2002) and Wachter (2003).

This formulation captures the empirical observation that investors respond to higher inflation expectations by reducing bond holdings (Ceballos and Ng, 2024). It is also consistent with the portfolio-choice framework in Campbell and Viceira (2001), who show that long-term investors prefer indexed bonds and shorten their nominal portfolios when inflation expectations are higher. In this setup, where only nominal bonds are available, Assumption 1 captures this idea by allowing preferred-habitat demand to fall more sharply at longer maturities, so that arbitrageurs are left holding a riskier residual supply. Importantly, this does not imply that any individual preferred-habitat investor substitutes across maturities. Rather, it means that the aggregate demand schedule across habitats shifts in a maturity-dependent way when inflation expectations increase.

Sources of risk A key novelty of the model is that interest-rate and inflation risks arise from the signal-extraction problem faced by private agents. In particular, the two equations that govern arbitrageurs' sources of risk are the Taylor-type rule (equation (1.7)) and the evolution of their inflation expectations (equation (1.11)). They are shown next for completeness:

$$di_t = \phi(\hat{\pi}_t - \bar{\pi})dt + dI_t; \tag{1.24}$$

$$d\hat{\pi}_t = -(\kappa_\pi + \phi\,\theta)(\hat{\pi}_t - \bar{\pi})\,dt + K_i\,dI_t + K_s\,dS_t, \tag{1.25}$$

where the first equation follows from (1.12).

Market clearing The market-clearing condition requires that the supply of nominal bonds at each maturity, expressed relative to arbitrageurs' real wealth, is matched by demand from preferred-habitat investors and arbitrageurs:

$$q_t^{(\tau)} = d_t^{(\tau)} + x_t^{(\tau)}. \tag{1.26}$$

1.4.2 Equilibrium Term Structure

To pin down equilibrium bond prices, $P_t^{(\tau)}$, I follow Greenwood and Vayanos's (2014a) solution strategy. I conjecture that prices are affine functions of the state variables:

$$P_t^{(\tau)} = e^{-[A_i(\tau) i_t + A_\pi(\tau)\hat{\pi}_t + C(\tau)]}, \tag{1.27}$$

where $A_i(\tau)$, $A_\pi(\tau)$, and $C(\tau)$ are functions of the time to maturity τ. Here, $A_i(\tau)$ and $A_\pi(\tau)$ capture the sensitivity of bond prices to the instantaneous nominal rate i_t, and to private agents' expected inflation $\hat{\pi}_t$, respectively. The equilibrium can be obtained in two steps: first, solve the arbitrageurs' optimization problem, imposing the conjectured form of bond prices; second, use the market-clearing condition to pin down the coefficients and verify that the conjecture is consistent with equilibrium.

Lemma 2. *The arbitrageurs' first-order condition is*

$$\mu_t^{(\tau)} - i_t = A_i(\tau)\lambda_{i,t} + A_\pi(\tau)\lambda_{\hat{\pi},t}, \tag{1.28}$$

where

$$\begin{aligned}\mu_t^{(\tau)} = {} & A_i'(\tau) i_t + A_\pi'(\tau)\hat{\pi}_t + C'(\tau) - A_i(\tau)\phi(\hat{\pi}_t - \bar{\pi}) + A_\pi(\tau)(\kappa_\pi + \phi\,\theta)(\hat{\pi}_t - \bar{\pi}) \\ & + A_i(\tau)A_\pi(\tau)K_i\sigma_i^2 + \frac{1}{2}\left[A_i(\tau)^2\sigma_i^2 + A_\pi(\tau)^2\left(K_i^2\sigma_i^2 + K_s^2\sigma_s^2\right)\right],\end{aligned} \tag{1.29}$$

and

$$\lambda_{i,t} = \gamma\sigma_i^2\left[\int_0^T x_t^{(\tau)}\left[A_i(\tau) + A_\pi(\tau)K_i\right]d\tau\right]; \tag{1.30}$$

$$\lambda_{\hat{\pi},t} = K_i\,\gamma\sigma_i^2\left[\int_0^T x_t^{(\tau)}\left[A_i(\tau) + A_\pi(\tau)K_i\right]d\tau\right] + \gamma K_s^2\sigma_s^2\left[\int_0^T x_t^{(\tau)} A_\pi(\tau)d\tau\right]. \tag{1.31}$$

Proof. See Appendix A.3.6. □

Equation (1.28) states that the instantaneous expected excess return on a nominal bond (relative to the nominal short rate) is a linear function of the bond's exposure to the various risk factors. The terms $\lambda_{i,t}$ and $\lambda_{\hat{\pi},t}$ represent the market prices of short-rate risk and inflation risk, each measuring the additional expected return required per unit of sensitivity to its respective risk factor. The market price of short-rate risk, $\lambda_{i,t}$, depends on the aggregate exposure of arbitrageurs' portfolios to the short rate, including the cross-effect that imperfect information has on their inflation expectations. Similarly, the market price of inflation risk, $\lambda_{\hat{\pi},t}$, reflects the risks arising from belief updating about inflation under imperfect information.

The conditions for an equilibrium solution to exist and the characterization of the functions $A_k(\cdot)$ for factor $k = i, \pi$ are presented in Theorem 1.

Theorem 1. *The functions $A_i(\tau)$ and $A_\pi(\tau)$ are given by*

$$A_i(\tau) = \tau; \tag{1.32}$$

$$A_\pi(\tau) = Z\left(\frac{\tau}{\kappa_\pi^\star} + \frac{e^{-\kappa_\pi^\star \tau} - 1}{\kappa_\pi^{\star 2}}\right), \tag{1.33}$$

respectively, where

$$Z \equiv \frac{\phi + \gamma\sigma_i^2 \int_0^T \tau\,\alpha(\tau)d\tau}{1 - \gamma\sigma_i^2 K_i \int_0^T \left(\frac{\tau}{\kappa_\pi^\star} + \frac{e^{-\kappa_\pi^\star}-1}{\kappa_\pi^{\star 2}}\right)\alpha(\tau)d\tau}. \tag{1.34}$$

and

$$\kappa_\pi^\star = \kappa_\pi + \phi\,\theta - \gamma\left[\sigma_i^2 K_i \int_0^T \tau\alpha(\tau)d\tau + Z\left(K_s^2\sigma_s^2 + \sigma_i^2 K_i^2\right)\int_0^T \left(\frac{\tau}{\kappa_\pi^\star} + \frac{e^{-\kappa_\pi^\star}-1}{\kappa_\pi^{\star 2}}\right)\alpha(\tau)d\tau\right]. \tag{1.35}$$

Additionally, a solution to (1.35) exists if γ is below a threshold $\bar{\gamma} > 0$.

Proof. See Appendix A.3.7. □

The existence argument follows Greenwood and Vayanos (2014a, Theorem 1). Equation (1.35) admits either no solution or an even number of solutions.[41] An equilibrium exists whenever the risk-aversion parameter satisfies $\gamma < \bar{\gamma}$, where $\bar{\gamma}$ is the smallest value at which $\kappa_\pi^\star$ touches zero. Throughout, I focus on that case and select the largest solution of equation (1.35), which delivers the stable, non-explosive equilibrium. Finally, note that equations (1.32) and (1.33) imply that both factors are increasing with maturity τ. This last observation will be useful when studying the impact of policy rate surprises on term premia.

1.4.3 Effects of a Monetary Policy Rate Surprise

I now examine how the model compares with the main empirical findings of this study. In particular, I trace out the response of the term structure of interest rates to surprise changes in the policy rate.

Sovereign Bond Yields Using equations (1.20) and (1.27), the affine bond-price solution implies nominal yields

$$y_t^{(\tau)} = \frac{1}{\tau}\left[A_i(\tau) i_t + A_\pi(\tau)\hat{\pi}_t + C(\tau)\right] \tag{1.36}$$

Consider now the effect of a surprise increase in the policy rate. A surprise one-unit increase in the policy rate, $dI_t = 1$, directly raises the short rate by one, but it also affects inflation expectations through the signal-extraction mechanism. Two forces are therefore at play. On the one hand, a higher-than-expected policy rate has a disinflationary effect,

[41] For a discussion on equilibrium multiplicity refer to Theorem 1 in Greenwood and Vayanos (2014a).

which dampens the response of long-term yields. On the other hand, under imperfect information, agents may update their inflation expectations upwards. This informational effect raises perceived inflation and can amplify the impact of the shock across the yield curve.

Which channel dominates depends on the precision of private signals about the inflation rate. When private signals are very clear ($K_i \to -\theta$), the informational channel is small.[42] In this case, the disinflationary force ensures that the effect of monetary policy rate surprises is concentrated at the short end, while long-term yields move little, or may even decline. By contrast, when signals are noisy ($K_i > 0$), the informational effect becomes more important. Since $A_\pi(\tau)$ increases with maturity, this channel generates a stronger and more persistent response at longer horizons. The model thus rationalizes the empirical findings of Figure 1.2. Specifically, it demonstrates why in some environments policy rate surprises have little impact on the long end, while in others they shift the entire curve upward. The next proposition formalizes this result.

Proposition 2. *Consider a surprise increase of one-unit in the policy rate, $dI_t = 1$. The response of bond yields is*

$$\frac{\partial y_t^{(\tau)}}{\partial dI_t} = 1 + \frac{A_\pi(\tau) K_i}{\tau} \tag{1.37}$$

Proof. See Appendix A.3.8. □

The decomposition in (1.37) shows that yield responses reflect two forces. The mechanical pass-through ensures that shorter-term yields always move in the same direction as the monetary policy rate. Beyond that, the slope of the yield curve hinges on the informational channel. When private signals are precise ($K_i < 0$), agents interpret a tightening as disinflationary news, so long-term yields increase less than short-term ones, producing a flattening of the yield curve. By contrast, when private signals are noisy ($K_i > 0$), private

[42] This can be readily seen in equation (1.14) because $\Lambda \to \infty$ as $\sigma_s \to 0$.

agents' inflation expectations increase, generating an upward shift of the entire nominal yield curve.

Market Prices of Risk In this environment, risk compensation is captured by the market prices of risk, $\lambda_{i,t}$ and $\lambda_{\hat{\pi},t}$, which are endogenous and summarize how much extra expected return arbitrageurs require per unit of exposure to each factor (see (1.30) and (1.31)). From (1.28), the expected excess return on a bond of maturity τ satisfies

$$\mu_t^{(\tau)} - i_t = A_i(\tau)\lambda_{i,t} + A_\pi(\tau)\lambda_{\hat{\pi},t}. \tag{1.38}$$

A monetary policy rate surprise, dI_t, affects these market prices in two steps. First, it shifts beliefs about inflation through the Kalman-Bucy update (1.25), with the strength of updating governed by K_i. Second, the resulting change in beliefs alters preferred-habitat demand (1.23), leaving arbitrageurs to absorb a different residual supply profile and thus changing their risk exposures. When private signals are noisier ($K_i > 0$), a surprise tightening raises $\hat{\pi}_t$, preferred-habitat investors shed longer-term nominal bonds, and arbitrageurs are forced to hold portfolios more sensitive to both interest rate and inflation risk. In this case, both $\lambda_{i,t}$ and $\lambda_{\hat{\pi},t}$ increase. By contrast, when private signals are more precise ($K_i < 0$), a tightening lowers expected inflation, bond demand from habitat investors rises, residual risk borne by arbitrageurs declines, and both market prices of risk fall. The following proposition formalizes these results.

Proposition 3. *Consider a surprise increase of one-unit in the policy rate dI_t. The response of the market prices of short-rate risk and inflation risk are*

$$sign\left(\frac{\partial \lambda_{i,t}}{\partial dI_t}\right) = sign\left(\frac{\partial \lambda_{\hat{\pi},t}}{\partial dI_t}\right) = sign(K_i) \tag{1.39}$$

Proof. See Appendix A.3.9. □

Nominal Term Premia A central object of interest, both empirically and in the model, is the nominal term premium: the compensation investors require for bearing interest rate and inflation risk when holding longer-term nominal bonds, beyond the expected path of policy rates. Empirically, in Section 1.2, I estimate term premia from yield curve data and examine how they respond to policy rate surprises.

In the model, a tightly analogous decomposition is possible. The nominal term premium at each maturity is defined as the difference between the forward rate and the expected future policy rate. To formalize this, I use the forward rate as a summary statistic for the marginal yield investors require to extend the maturity of their investment. The forward rate from $(t+\tau-\Delta\tau)$ to $(t+\tau)$ is

$$f_t^{(\tau-\Delta\tau,\tau)} = -\frac{\log\left(P_t^{(\tau)}/P_t^{(\tau-\Delta\tau)}\right)}{\Delta\tau}. \tag{1.40}$$

In the limit as $\Delta\tau \to 0$, the instantaneous forward rate is given by

$$f_t^{(\tau)} = -\frac{\partial \log\left(P_t^{(\tau)}\right)}{\partial\tau} = A_i'(\tau)i_t + A_\pi'(\tau)\hat{\pi}_t + C'(\tau), \tag{1.41}$$

where $A_i'(\cdot)$, $A_\pi'(\cdot)$ and $C'(\cdot)$ denote derivatives of the affine price coefficients with respect to maturity.

Crucially, the model's predictions for how forward rates respond to policy rate surprises translate directly into predictions for the term premium. If a surprise tightening leads to a larger movement in forward rates than can be explained by changes in the expected path of the short rate alone, the term premium increases, mirroring the "overreaction" empirically observed in emerging markets. Conversely, a muted forward rate response corresponds to

a decline in the term premium, consistent with the findings for advanced economies.

The following proposition therefore mirrors the empirical strategy: I trace out the response of model-implied nominal term premia (i.e., the excess movement in forward rates over the change in the expected short rate path) to monetary policy surprises.

Proposition 4. *Consider a surprise increase of one-unit in the policy rate $dI_t = 1$. The response of the instantaneous forward rate is*

$$\frac{\partial f_t^{(\tau)}}{\partial dI_t} = 1 + K_i \frac{Z}{\kappa_\pi^\star} \left(1 - e^{-\kappa_\pi^\star \tau}\right). \tag{1.42}$$

The response of the expected future short rate is

$$\frac{\partial E_t[i_{t+\tau}]}{\partial dI_t} = 1 + K_i \frac{\phi}{\kappa_\pi + \phi\,\theta} \left(1 - e^{-(\kappa_\pi + \phi\,\theta)\tau}\right). \tag{1.43}$$

Thus, there is overreaction of the forward rate relative to the expected short rate

$$\frac{\partial f_t^{(\tau)}}{\partial dI_t} > \frac{\partial \mathbb{E}_t[i_{t+\tau}]}{\partial dI_t}, \tag{1.44}$$

if $K_i > 0$, which occurs if and only if

$$\frac{\phi}{\sigma_i^2} \frac{-\kappa_\pi + \sqrt{\kappa_\pi^2 + \left(\frac{\phi^2}{\sigma_i^2} + \frac{1}{\sigma_s^2}\right)\sigma_\pi^2}}{\frac{\phi^2}{\sigma_i^2} + \frac{1}{\sigma_s^2}} > \theta. \tag{1.45}$$

Conversely, there is underreaction of the forward rate relative to the expected short rate

$$\frac{\partial f_t^{(\tau)}}{\partial dI_t} < \frac{\partial \mathbb{E}_t[i_{t+\tau}]}{\partial dI_t}, \tag{1.46}$$

if $K_i < 0$.

Proof. See Appendix A.3.10. □

The proof of Proposition 4 rests on the fact that forward rates embed risk compensation through the market prices of risk, whereas expected short rates do not. When $K_i > 0$, a surprise tightening raises investors' inflation expectations, preferred-habitat demand for nominal bonds falls, and arbitrageurs are forced to absorb a riskier residual supply. Their required compensation for bearing interest-rate and inflation risk rises, pushing up both $\lambda_{i,t}$ and $\lambda_{\hat{\pi},t}$, and producing an overreaction of forward rates relative to the expected short rate (i.e. the nominal term premium increases). By contrast, when $K_i < 0$, a tightening lowers expected inflation, preferred-habitat demand increases, arbitrageurs hold less risky supply, and market prices of risk fall. Forward rates then move less than the expected short rate path, yielding underreaction.

In this way, the model shows that the information environment surrounding monetary policy announcements determines whether policy surprises amplify or dampen term premia. It also yields a straightforward, testable implication: when private signals are noisy, monetary policy rate changes become informative signals that prompt agents to revise their expectations about the inflation outlook. In the next section, I take this prediction to the data by examining the dispersion of professional inflation forecasts. I show that disagreement is substantially higher in emerging markets than in advanced economies, consistent with noisier private information and with the view that monetary policy surprises in EMs convey relatively greater information about future inflation.

1.5 Testing the Model's Information Channel

Uncertainty is a central determinant of economic agents' decision-making, and a large body of research has shown that it dampens investment and growth.[43] This issue is especially salient in emerging markets, where macroeconomic fluctuations tend to be larger. For instance, emerging markets exhibit greater inflation volatility (Ha et al., 2019).[44]

To test the main link between the model's mechanism and the empirical findings, I use data on analysts' forecasts from *Bloomberg*, starting in 2010 where available. For each country and period, I collect the cross-sectional standard deviation of forecasts, the number of analysts contributing, and the average forecast across respondents. Periods with fewer than five available forecasts are excluded. Although analysts may revise their submissions up to the release date, I only observe the final forecast. Over 90% of forecasts are submitted within the two weeks preceding the release, suggesting that analysts incorporate the most relevant information available at that time. Finally, I exclude clear typographical errors (e.g., negative unemployment forecasts), which represent less than 0.1% of observations and do not materially affect results.

Table 1.4 shows that forecast dispersion around economic releases is larger in emerging markets than in advanced economies. Comparing the median emerging market with the median advanced economy, dispersion in inflation forecasts is about 60 percent larger in EMs, pointing to greater disagreement among analysts about the inflation outlook. Looking at other macroeconomic variables yields the same implication. For real variables such as

[43] See, for example, Ramey and Ramey (1994); Bloom (2009); Haddow et al. (2013); Bekaert et al. (2013); Creal and Wu (2017); Bloom et al. (2018).

[44] The literature has also found that emerging markets exhibit greater volatility in fiscal variables (Marioli et al., 2024), and in output and consumption growth (Aguiar and Gopinath, 2007). Moreover, the index of central bank transparency constructed by Dincer et al. (2022) shows that emerging markets perform particularly poorly on the economic transparency component, which measures the extent to which central banks disclose the data and internal projections used to formulate monetary policy. This is also in line with the IAPOC (Independence and Accountability, Policy and Operational Strategy, and Communications) index of monetary policy frameworks in Unsal et al. (2022).

real GDP growth and unemployment, forecasters in EMs also display roughly twice the degree of disagreement compared with their AE counterparts.[45] These findings suggest that market participants in EMs operate with noisier priors, making it harder to separate pure monetary policy shocks from signals about fundamentals.

1.5.1 Validity of Bloomberg analysts' forecasts

A natural question is whether the forecasts provided by market analysts to Bloomberg are reliable and informative. The main advantage of using these forecasts is that they offer a homogeneous and internationally comparable sample. In contrast, national surveys of macroeconomic expectations often differ in structure, periodicity, and respondent composition. Bloomberg surveys, by design, mitigate these sources of heterogeneity: participants are professional analysts from major financial institutions, research firms, and consultancies. For each country and data release, analysts can submit and revise their forecasts at any time prior to the official announcement. Importantly, they have strong incentives to report accurate forecasts, since Bloomberg publicly displays the name of each contributing institution and maintains rankings of top-performing forecasters.

[45] Results are robust to excluding the pandemic period (2020-22; see Table A.10). Additionally, the time-series evolution of the standard deviation is presented in Figure A.7.

Table 1.4: Average Dispersion of Macroeconomic Forecasts, 2010–2022

A. Emerging Markets

	Annual inflation (bps)	Real GDP YoY% (bps)	Unemployment rate (bps)
Brazil	8.7	34.9	14.8
Chile	10.6	16.7	14.2
Colombia	9.1	69.8	42.2
Hungary	17.0	49.2	11.9
India	23.2	74.1	
Indonesia	16.2	30.0	
Malaysia	19.2	75.5	
Mexico	8.1	38.2	13.1
Nigeria	23.1		
Peru			24.1
Philippines	17.3	66.0	
Romania	17.6	59.0	
Russia	17.4	46.8	13.9
South Africa	13.5	39.9	54.3
Thailand	20.0	73.9	
Median	**17.1**	**49.2**	**14.5**

B. Advanced Economies

	Annual inflation (bps)	Real GDP YoY% (bps)	Unemployment rate (bps)
Australia	12.5	25.9	8.4
Canada	11.7	15.1	13.1
Czech Republic	9.9	42.4	6.8
Euro Area	9.8	21.5	5.9
Japan	9.4	21.7	5.6
New Zealand	11.0	27.7	9.4
Norway	11.9	31.5	13.9
South Korea	10.7	17.9	7.4
Sweden	9.3	20.9	9.4
Taiwan	11.0	56.4	3.5
United Kingdom	10.4	17.7	6.4
United States	8.4	19.1	9.5
Median	**10.5**	**21.6**	**7.9**

Note: This table presents the average cross-sectional standard deviation of analysts' forecasts for annual inflation, real GDP growth, and the unemployment rate over the period 2010–2022. Dispersion is expressed in basis points (bps) for all variables. Months with fewer than five available forecasts are excluded. Data is from *Bloomberg*.

To assess the validity of these forecasts, I cross-check the Bloomberg estimates with measures of forecast dispersion from national surveys for three emerging markets: Colombia, India, and Mexico.[46] Comparing the dispersion in national surveys to that from Bloomberg forecasts provides a useful benchmark for evaluating whether the latter reflect meaningful differences in expectations across analysts, rather than noise or data artifacts.

For these three countries, the standard deviation of the 12-month-ahead inflation forecast is roughly four to six times larger in the central bank surveys than in Bloomberg. This discrepancy may arise for several reasons. First, national surveys are typically conducted at fixed dates each month, while Bloomberg allows analysts to update their forecasts continuously until the official release, thereby incorporating the most recent macroeconomic and financial information. Second, respondents to central bank surveys cannot revise their submissions once sent. Third, the anonymity of most national surveys may reduce respondents' incentives to devote attention to their estimates.

Beyond these mechanical differences, the lower dispersion in Bloomberg forecasts may also reflect differences in information sets and professional incentives. Bloomberg participants (often economists at financial institutions) tend to process high-frequency data and central bank communication more actively, which may lead to faster convergence of beliefs around a common information set. By contrast, respondents to central bank surveys may comprise a broader mix of institutions with heterogeneous analytical capacities, forecasting horizons, and incentives, resulting in wider disagreement. Overall, the smaller dispersion in Bloomberg forecasts is best interpreted as evidence of higher information precision among professional forecasters, supporting their use as a reliable and informative measure of expectations in this paper.

[46] Country selection is based solely on data availability for the standard deviation of forecasts. For Colombia, I use the *Encuesta mensual de expectativas de analistas económicos* from Banco de la República. For India, I use the *Survey of Professional Forecasters* conducted by the Reserve Bank of India. For Mexico, I use the *Encuesta sobre las expectativas de los especialistas en economía del sector privado* from Banco de México.

1.6 Discussion

This paper documents a stark asymmetry in how unexpected policy tightenings transmit to the long end of the yield curve across country groups: in advanced economies, the nominal term premium falls; in emerging markets, it rises. Decomposing monetary policy surprises shows that the EMs' pattern is driven not by pure policy shocks but by the information embedded in surprises. A preferred-habitat model with asymmetric information rationalizes these facts: when private agents have noisy information about the inflation rate, monetary policy rate changes double as signals about the outlook, increasing inflation expectations and prompting portfolio rebalancing that raises the required compensation for holding long-duration nominal debt. Survey evidence (higher dispersion in inflation forecasts in EMs) supports this channel.

What does these results imply for policy? First, monetary policy surprises can be cost increasing in emerging markets. When information frictions are pervasive, a surprise hike can increase long-term borrowing costs by increasing term premia, even if the intended effect of the policy action is to damp inflation. That said, a rise in long-term yields is not necessarily undesirable from the standpoint of inflation stabilization alone, since tighter financial conditions at longer maturities may reinforce the contractionary effects of policy. The concern highlighted here is more specific: in emerging markets, the upward shift in the long end appears to reflect higher inflation expectations and inflation-risk compensation rather than a cleaner transmission of policy. In such environments, more hawkish surprises may tighten at the short end while simultaneously making duration financing more expensive for sovereigns and the private sector, thereby increasing the fiscal and macro-financial costs of disinflation.

Second, central bank communication itself should be viewed as an active policy instrument, not merely a complement to rate decisions. When markets operate under imperfect

information, the central bank's public statements, forecasts, and reaction function jointly shape how investors interpret policy actions. Opacity amplifies the informational content of monetary policy surprises, turning decisions into a signal about the central bank's private assessment of the economy. In an era of inflation targeting, if the private sector has diffuse information about the economy, and if this information is less accurate than the central bank's, full transparency is optimal (Amato et al., 2002; Walsh, 2005). Strengthening communication frameworks can lower the "information premium" embedded in the yield curve. Greater clarity about the central bank's outlook and policy rationale not only improves the efficiency of monetary transmission but also reduces the fiscal cost of debt financing. As per the International Monetary Fund, practical steps to improve central bank communication include:

- Releasing inflation and output projections together with conditional policy-rate paths to clarify how current actions relate to future objectives.
- Publishing detailed post-meeting minutes or summaries that explain the rationale for decisions and the data underlying them, reducing room for markets to infer private assessments.
- Maintaining coherent messages across speeches, reports, and press releases. Additionally, avoiding mixed signals and clearly distinguishing between objectives such as price stability and exchange-rate management.
- Disclosing foreign exchange interventions and clarifying the use of multiple instruments to demonstrate alignment with the inflation-targeting framework and prevent misinterpretation of actions as hidden signals.

Third, monetary authorities can reinforce policy credibility by coordinating more closely with public debt management offices. This coordination would signal policy coherence be-

tween monetary and fiscal authorities, helping markets interpret higher policy rates as consistent with a credible disinflation effort rather than as evidence of rising fiscal or inflation risk. Over the medium term, developing deeper inflation-indexed bond markets can further enhance the transmission of monetary policy and the credibility of the central bank's commitment to price stability.

Fourth, a broader implication of these findings concerns macro-financial stability. In many emerging markets, the presence of liability dollarization, short average debt maturities, and limited hedging markets makes the yield curve a key transmission channel between monetary policy and financial stability. When term premia rise abruptly in response to perceived information in policy announcements, long-term borrowing costs can increase sharply, amplifying refinancing and valuation risks for banks, firms, and the government itself. These inference-driven shifts in risk compensation can interact with exchange rate pressures and balance-sheet mismatches, potentially destabilizing financial conditions. Macroprudential frameworks should therefore internalize the informational dimension of monetary policy (i.e., by strengthening duration and liquidity requirements, tightening limits on unhedged FX exposures, and coordinating with monetary authorities to mitigate amplification through the term structure).

Chapter 2: Foreign Investors in Local-Currency Bond Markets: Implications for Bond Yields and Exchange Rates

Joint with Pierre De Leo, Lorena Keller, Mauricio Villamizar-Villegas, and Tomás Williams

2.1 Introduction

There is increasing recognition that financial intermediaries play a central role in determining bond yields and exchange rates (Gabaix and Maggiori, 2015; Vayanos and Vila, 2021b; Kekre et al., 2024b; Greenwood et al., 2023a; Gourinchas et al., 2022b; Kekre and Lenel, 2025b). In the presence of intermediation frictions, these prices respond to shifts in investors' demand for assets across maturities and currencies, which intermediaries must absorb. This paper shows that there are important differences in the bond demand of domestic and foreign investors, because they induce intrinsically different changes in intermediaries' positions. As a result, the composition of investors in local-currency bond markets affects the joint behavior of bond yields and exchange rates and, in turn, the transmission of asset purchase policies by central banks.

We begin by showing that foreign investors—a significant presence in many local-currency government bond markets—induce correlated flows across a country's long-term bond and currency markets. Using unique transaction-level data from Colombia's bond and foreign exchange markets, we show that trades by foreign investors in local-currency bonds are accompanied by tightly linked transactions in currency markets, while this is

not the case for domestic investors' trades. These systematically different patterns arise naturally from differences in the funding currency of domestic and foreign investors. Because foreign investors fund their purchases in foreign currency, acquiring local-currency bonds requires a simultaneous acquisition of local for foreign currency. As long as foreign investors don't fully hedge their currency exposures, their long-term bond purchases thus induce correlated changes in intermediaries' long-term bond and currency positions.

In the presence of intermediation frictions, foreign investors' bond purchases generate correlated revisions in term premia and currency premia. We formalize this mechanism by developing an equilibrium model of bond yields and exchange rates with frictions in financial intermediation. Our framework builds on Greenwood et al. (2023a) but incorporates a distinction between domestic and foreign investors. When foreign investors demand local-currency bonds, intermediaries' exposure to long-term bond price risk and currency risk contemporaneously declines. As a result, they require lower term premia and lower currency premia, leading to a decline in bond yields and an exchange rate appreciation.

The mechanism conveyed by our structural model accounts for the effects of an identified foreign inflow into Colombia's local-currency bond market. In March 2014, J.P. Morgan announced that Colombian long-term government bonds would be included in one of its flagship global bond indexes, triggering large bond purchases by benchmark-tracking foreign investors adjusting their portfolios to the new index. We show that these purchases were matched almost one-for-one by foreign investors' acquisitions of Colombian pesos in the spot market. During this episode, domestic intermediaries absorbed foreign investors' demand, offloading long-term bonds and currencies from their portfolios. Alongside, in line with the mechanism described above, bond prices increased and the exchange rate appreciated.

We also document that Colombian intermediaries' bond and currency positions are positively correlated, and that their portfolio returns reflect compensation for bearing bond and

currency risk. The correlation between intermediaries' holdings of long-term government bonds and Colombian pesos is around 40%, consistent with foreign investors participating in Colombian pesos bond markets without fully hedging their currency exposure. We also find that changes in intermediaries' bond and currency positions are positively associated with subsequent excess bond and currency returns. Taken together, these findings support the view that intermediaries are the marginal agents absorbing correlated imbalances across maturities and currencies, and that they are compensated through term premia and currency premia for doing so.

The composition of investors in local-currency bond markets affect the comovement between bond yields and exchange rates. In our model, this comovement depends on the relative importance of two sets of shocks, those to short-term interest rates and those to foreign investors' unhedged bond flows. Short-term interest rate shocks affect bond and currency returns in opposite directions. An increase in the domestic short-term interest rate causes a decline in the price of long-term bonds, through an expectations hypothesis channel, while appreciating the domestic currency, through an uncovered interest parity channel. Interest rate risk thus induces a *negative* correlation in bond and currency returns. In contrast, because foreign investors induce correlated changes in intermediaries positions, they naturally generate *positive* correlation in bond and currency returns, precisely because of the positively correlated revisions in term premia and currency risk premia. Thus, bond and currency risk can either offset or compound depending on the relative importance of foreign investors in local currency bond markets.

Existing models of frictional financial intermediation have largely emphasized short-term interest rate shocks, but assume that investors' bond demand shocks absorbed by intermediaries are independent, effectively abstracting from the presence of foreign investors (Greenwood et al., 2023a; Gourinchas et al., 2022b). In other words, all portfolio shifts in these models affect the net supply of domestic bonds without directly altering intermedi-

aries currency positions. As a result, in these models the comovement in term premia and currency premia is negative, as it exclusively reflect bond and currency returns' exposure to interest rate shocks. By contrast, our model can generate any covariance between bond and currency returns. For Colombia, for example, unhedged foreign investor flows raise the return correlation by about 8 percentage points, and it is essential to account for the overall positive correlation.

We show that countries in which unhedged foreign investor flows are more prevalent exhibit more positive comovement between bond and currency returns. In the model, the importance of unhedged foreign investor flows is governed by two forces. First, the relative volatility of bond purchases by foreign versus domestic investors. When a larger fraction of the variation in bond demand reflects foreign investors, changes in intermediaries' bond and currency positions are more correlated. Second, the extent of currency hedging by foreign investors. A higher tendency of foreign investors to hedge currency spot purchases, through currency forwards, naturally dampens the correlation in intermediaries' bond and currency positions.

Using cross-country data on local-currency bond holdings (Arslanalp and Tsuda, 2012, 2014) and estimates of foreign investors' hedging ratios (Cheema-Fox and Greenwood, 2024; Chen and Zhou, 2025), we document substantial heterogeneity in the incidence of foreign investors' flows in local bond markets and in their currency hedging behavior. Most importantly, countries where foreign investors drive a larger share of bond flows, or hedge less, display more positive comovement between bond and currency returns. Thus, foreign investor participation in local currency bonds appears to contribute to the observed cross-country differences in the comovement of bond yields and exchange rates (an empirical object recently emphasized in Lustig et al., 2019).

Our analysis also speaks to how asset purchase policies, such as quantitative easing and foreign exchange interventions, transmit to bond yields and exchange rates. In our

framework, intermediaries operate simultaneously in bond and currency markets, which are exposed to correlated risks. As a result, when the net supply of one asset shifts, intermediaries adjust their required premia in both markets. Whether bond purchases appreciate or depreciate the currency—or whether foreign exchange interventions raise or lower bond prices—depends on the sign of the covariance between bond and currency returns, which, in turn, reflects the relative importance of unhedged foreign investor flows versus short-term interest rate shocks, as described above.

We test these predictions using a regression discontinuity design based on two types of auctions conducted in Colombia between 2012 and 2014: Treasury auctions to issue local-currency government bonds and central bank auctions to buy U.S. dollars. For identification, we exploit the discontinuity at each auction's cutoff price to compare marginal winners and losers. We find that dealers required to absorb more government bonds subsequently trade both bonds *and* dollars at worse prices, consistent with simultaneously demanding higher term and currency premia. Analogously, dealers that sell dollars to the central bank subsequently trade both dollars *and* bonds at worse prices. These results highlight that term premia and currency premia comove positively in response to shifts in the net supply in a single market.

Related literature Models of frictional financial intermediation have gained prominence in international finance (Gabaix and Maggiori, 2015; Itskhoki and Mukhin, 2021; Kekre and Lenel, 2024; Chahrour et al., 2024; Gourinchas et al., 2022b; Greenwood et al., 2023a; Kekre and Lenel, 2025b). Moreover, there is mounting evidence that shocks to bond and currency demand cause movements in bond yields and exchange rates (Pandolfi and Williams, 2019; Ray et al., 2024; Hau et al., 2010; Beltran and He, 2024). We contribute to this literature by analyzing a new mechanism that links foreign investors' bond flows to the portfolios of financial intermediaries, and in turn to the joint dynamics of bond yields and

exchange rates. We thus offer a new interpretation of cross-country differences in the co-movement between bond yields and exchange rates (Lustig et al., 2019; Lloyd and Marin, 2024; Kekre and Lenel, 2024; Rebucci et al., 2025). We also trace out implications for the transmission of asset purchase and sale policies, and provide novel causal evidence based on intermediaries' behavior during Treasury and central bank auctions in Colombia.

Our proposed mechanism is related to, yet distinct from, several recent contributions. Some studies propose mechanisms whereby fluctuations in the U.S. dollar affect risk premia by revaluing the wealth of constrained intermediaries (Carstens and Shin, 2019; Bruno and Shin, 2015; Hofmann et al., 2022; Bertaut et al., 2024; Kekre and Lenel, 2025b). In contrast, we emphasize the role of foreign investors' financial flows in shaping intermediaries' positions across bond and currency markets. That is, we emphasize a channel that operates through changes in intermediaries' positions (a "portfolio balance" channel), rather than through changes in intermediaries' price of risk. Our proposed mechanism is thus closely related to studies of the joint determination of capital flows, exchange rates, and equity prices (Hau and Rey, 2006; Camanho et al., 2022; Rey and Stavrakeva, 2024).[1] Hau and Rey (2006) and Camanho et al. (2022) document that net equity flows into foreign markets are associated with currency appreciations, which echoes our empirical findings for Colombia's local currency bond markets. While our paper relies on an analogous mechanism, it instead focuses on government bond markets. In addiction, our paper uniquely highlights the relationship between foreign investors and bond–currency return comovement, and its implications for the transmission of asset purchase policies, both theoretically and empirically. In this context, we note that it is the relative volatility of foreign investors' bond flows, rather that their average bond holdings, that matters for the comovement between bond and currency returns. Our paper is thus related to recent work studying the

[1] Koijen and Yogo (2020) estimates a demand system to study exchange rates jointly with short-term rates, long-term yields, and equity prices.

differential volatility between foreign and domestic investors' financial flows, and in their sensitivities to global financial conditions (Zhou, 2024; Fu, 2023). While these papers investigate the sources of differences in investors' volatilities, we take the relative volatility of foreign investors as given, and highlight how it shapes the comovement between bond yields and exchange rates through their effects on intermediaries' portfolios.[2]

This paper also relates to studies that focus on the broader implications of foreign investors, following Calvo et al. (1993). Among many questions, this literature investigates the role of foreign investors in the dynamics of short-term market rates (De Leo et al., 2024b), deviations from covered interest parity (De Leo et al., 2024c), deviations from uncovered interest parity (Kalemli-Özcan, 2019; di Giovanni et al., 2021; Cormun and De Leo, 2024), sovereign and firm borrowing costs (Fang et al., 2024; Morelli et al., 2022; Moretti et al., 2024; Morais et al., 2019; Xu, 2024), and patterns of capital flows (Avdjiev et al., 2022).

This paper also contributes to the literature on the transmission of policies that alter the net supply of long-term bonds or foreign currency. In models of frictional financial intermediation, the effects of these policies, such as quantitative easing, depend on the equilibrium covariance between bond and currency returns Gourinchas et al. (2022b); Greenwood et al. (2023a); Kekre and Lenel (2025b). Empirical tests of these predictions are often based on high-frequency responses of asset prices to announced or implemented quantitative easing programs or dollar swap lines (Bauer and Neely, 2014; Neely, 2015; Swanson, 2021b; Bhattarai and Neely, 2022; Rebucci et al., 2022; Arslan et al., 2020; Toraman, 2025; Kekre and Lenel, 2025a,b).

Our contribution to this literature is twofold. First, we note that investor composition shapes the transmission of these policies by affecting the equilibrium covariance between

[2]Other papers, instead, study the determinants and dynamics of sovereign debt currency composition (Ottonello and Perez, 2019; Du and Schreger, 2022; Engel and Park, 2022; Lee, 2024).

bond and currency returns, and we extend these predictions beyond bond purchase policies to the context of interventions in the currency. Second, we provide clean, causal evidence based on intermediaries' behavior in Treasury auctions of local-currency government bonds and central bank auctions to purchase U.S. dollars. This setting is directly informative for these models' predictions, as both types of auctions alter the net supply of assets that intermediaries must absorb. Our evidence speaks directly to the channel of interest by examining the trading behavior of intermediaries whose balance sheet portfolios are differentially affected by these policies in an arguably exogenous way.

2.2 Foreign investors in local-currency bond markets

All local-currency government bond markets feature both domestic and foreign investors (Arslanalp and Tsuda, 2012, 2014). A key difference between these investors lies in their funding currency: domestic investors fund themselves in domestic currency, while foreign investors fund themselves in foreign currency. We show that this distinction gives rise to fundamental differences in the correlation of financial flows across bond and currency markets. To do so, we use transaction-level data from Colombia's bond and currency markets. The sample period goes from January 2013 through December 2020, unless otherwise noted. Colombia is a representative small open economy, with foreign investor participation in its government bond market at levels comparable to those observed in other small open economies.

2.2.1 Institutional details of Colombia's bond and currency markets

Government bond market Since the 2000s, most of Colombia's sovereign debt is denominated in its domestic currency, the Colombian peso (COP). By 2020, total fiscal debt is USD 177 billion, representing around 65% of GDP. Of this total, two-thirds was de-

nominated in COP, with 20% issued as inflation-linked bonds and 46% as standard peso-denominated instruments.

The Ministry of Finance designates financial institutions to participate in the "primary dealers" (PD) program for government bonds (*Títulos de Tesorería*, TES). Membership in this program is limited to a select group of institutions that meet the Ministry's eligibility criteria. On average, the primary market issues approximately COP 300 billion (less than USD 100 million) worth of TES bonds daily. In contrast, the secondary market for TES, described below, experiences a substantially higher daily turnover of around COP 2 trillion (USD 500 million).

Foreign exchange markets The COP-USD spot and forward interdealer market in Colombia is highly centralized. A single electronic trading platform, *SET-ICAP FX*, accounts for approximately 95% of total market volume. Offshore trading is restricted by regulatory measures.

Transactions in foreign exchange markets are restricted to authorized dealers (*Intermediarios del Mercado Cambiario*, IMC). Consequently, all market participants must conduct transactions through one of these 50 authorized dealers. These intermediaries include banking institutions, financial corporations, financial cooperatives, and stock brokerage firms. The average daily turnover for the spot market is USD 1.5 billion, while the forward market sees an average daily turnover of USD 4 billion.[3]

[3]These institutions are subject to FX mismatch regulations overseen by the central bank, *Banco de la República*. The two key metrics are the *Posición Propia de Contado* (PPC) and the *Posición Propia* (PP), both measured as the difference between foreign-currency assets and liabilities relative to bank equity. The PPC captures short-term spot foreign-currency mismatches, while the PP measures overall net FX exposure by also incorporating derivatives positions. Hence, the PPC limits short-run foreign-currency funding risk, whereas the PP constrains total currency risk on and off the balance sheet. During our sample, intermediaries were required to keep PP within $[-5, 20]$ percent of equity and PPC within $[0, 50]$ percent.

Data sources Our analysis uses several datasets. Transactions involving purchases and sales of bonds in the secondary market take place on two trading platforms: (i) the Colombian Electronic Market (MEC), operated by the Colombian Stock Exchange, and (ii) the Electronic Trading System (SEN), managed by the central bank. All trades executed on these platforms are registered in the Central Securities Depository (DCV), overseen by the central bank. We rely on the DCV, which records counterparties, to observe all transactions occurring in the secondary market. To identify trades involving foreign investors, we use data from the "Declarations of Foreign Exchange Transactions" (*Declaraciones de Cambio*), which records all transactions involving foreign exchange and is compiled by the Technical and Economic Information Department of the central bank. To analyze FX Spot, Next-Day, and Forward markets in the interdealer market, we use data from the *SET-ICAP FX* platforms.[4]

2.2.2 Bond purchases by foreign and domestic investors

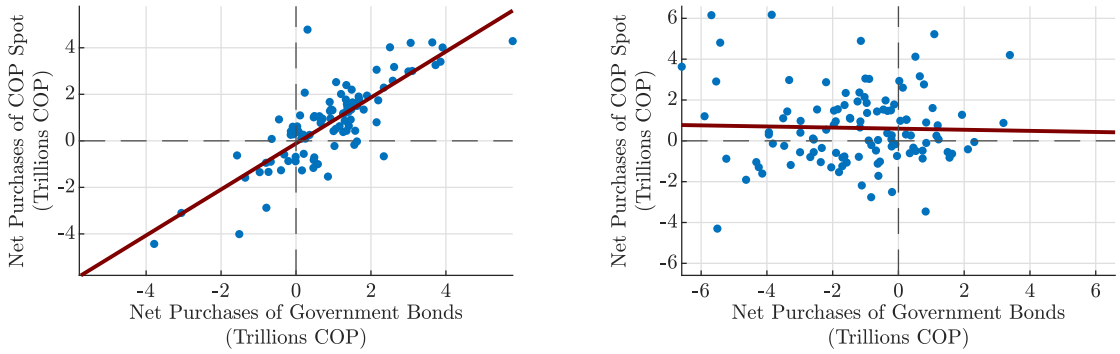
Foreign investors use foreign currency—typically the U.S. dollar—to fund their local-currency bond purchases, whereas domestic investors fund themselves in domestic currency, that is, COP. This simple observation implies that bond purchases by foreign and domestic investors generate fundamentally different financial flows in foreign exchange markets.

Figure 2.1a plots foreign investors' monthly net purchases of local-currency bonds (TES) in the secondary government bond market (x-axis) against their net purchases of domestic currency (COP) in the spot foreign exchange market (y-axis).[5] The data lie tightly

[4]While bond trades can be directly attributed to individual foreign investors, foreign exchange transactions may be conducted either directly by these investors or through foreign intermediaries, obscuring the identity of the ultimate buyer. Consequently, it is not possible to establish a precise link between bond and foreign exchange transactions trades at the individual transaction level.

[5]During this period, total outstanding TES debt was around COP 250 trillion, of which foreign investors held roughly one-quarter.

Figure 2.1: Investors' flows in bond and currency markets



(a) By foreign investors

(b) By domestic investors

Note: The figure plots monthly net purchases of government bonds (x-axis) against net purchases of COP in the spot market (y-axis). Panel 2.1a displays net purchases by foreign investors; Panel 2.1b displays net purchases by domestic investors. The fitted OLS regression line is shown in each panel.

along the 45-degree line. A regression of COP purchases on TES purchases for foreign investors yields a slope statistically indistinguishable from 1, with an R^2 of around 65%. A large share of foreign investors' net spot FX transactions is thus directly tied to their government bond market activity.

By contrast, domestic investors' net bond purchases are not systematically correlated with their transactions in the spot FX market (Figure 2.1b). A regression of COP purchases on TES purchases for domestic investors yields a slope indistinguishable from zero, with an R^2 close to zero.

These patterns arise naturally from differences in funding currency. A COP-funded investor purchasing X COP of a local-currency bond operates entirely within the bond market. The intermediary on the other side of the trade delivers X COP worth of bonds from inventory in exchange for COP cash, with no foreign exchange transaction. A U.S. dollar-funded investor, by contrast, must first convert USD into COP at the spot rate Q to acquire the same bond. The intermediary then delivers the bond from inventory while receiving X/Q USD in exchange, as a result of the combined foreign investor FX and bond transaction. From the intermediary's perspective, the transaction thus generates both a COP

bond outflow and a USD inflow.

Foreign and domestic investors' bond flows thus affect intermediary balance sheet risk in qualitatively different ways. In particular, foreign investor activity induces a correlated change in bond and currency positions for the intermediaries that absorb these flows.

We next formalize these insights in a model of bond and currency markets with frictional financial intermediation. The main purpose of the model is to derive the asset price implications arising from these differences in investor composition. In the model, we also consider the role of currency hedging by foreign investors. Following Greenwood et al. (2023a), we model foreign and domestic investors' bond demand as exogenous and inelastic. Although investor behavior is undoubtedly more complex in practice, the patterns documented in Figure 2.1 are driven by a fundamental difference in funding currency and would arise even in environments with fully endogenous portfolio choice.[6]

2.3 Implications for bond yields and exchange rates

We present a model in which government bond and foreign exchange markets, including spot and forward markets, are integrated with one another but segmented from other financial markets. A group of (domestic) intermediaries trades in all of these markets, conducting both the "yield-curve trade"—the trade that borrows short-term and lends long-term in domestic currency—and the "UIP trade"—the trade that borrows short-term in foreign currency and lends short-term in domestic currency. These two distinct trades are exposed to two distinct sources of risk: bond price risk and currency risk.

In the model, two main sets of shocks determine the equilibrium properties of returns on the yield curve trade and the UIP trade: shocks to short-term interest rates, and shocks

[6]Studies incorporating endogenous portfolio choice by foreign investors include Morelli et al. (2022); Akinci et al. (2022); Kekre et al. (2024b); Kekre and Lenel (2025b); Bruno and Shin (2015); Hofmann et al. (2022).

to domestic and foreign investors' demand for bonds and currencies. As emphasized by Greenwood et al. (2023a) and Gourinchas et al. (2022b), shocks to short-term interest rates affect the returns from holding bonds of different maturities and currencies through expectations hypothesis and uncovered interest parity channels, while shocks to demand alter intermediaries' balance sheet exposures. We emphasize that the origin of these demand shocks—whether from foreign or domestic investors—affects intermediary balance sheets in distinct ways, which is key to the equilibrium properties of returns on the yield curve trade and the UIP trade.

We now present the model and derive its equilibrium properties. We then outline several implications of the model for investors' positions in bond and currency markets, equilibrium bond yields and exchange rates, and their responses to asset purchase policies, and compare these implications to the Colombian data introduced earlier (Section 2.4 and Section 2.5). Finally, Section 2.6 examines the model's mechanism using cross-country data on investor composition and bond yield–exchange rate comovement.

2.3.1 Baseline Model

The discrete-time model includes domestic short- and long-term bonds, as well as foreign short-term bonds, denominated in their respective currency. There are two types of agents: domestic and foreign bond investors, and intermediaries. Foreign and domestic bond investors have a preference—resulting in inelastic demand—for assets of specific currencies and maturities. Intermediaries, on the other hand, are specialized investors who absorb the net supply of domestic long-term bonds and foreign exchange resulting from exogenous demand shocks.

We follow Greenwood et al. (2023a) and assume that asset prices (or yields) and ex-

pected returns are linear functions of a vector of state variables. To model fixed income assets, we (i) substitute log returns for simple returns throughout and (ii) use Campbell and Shiller (1988) linearizations of log returns.

Short-term bonds Short-term bonds in both currencies are supplied perfectly elastically, and short-term interest rates are determined exogenously according to AR(1) processes:

$$i_{t+1} = \bar{\imath} + \phi_i(i_t - \bar{\imath}) + \varepsilon_{i_{t+1}}; \tag{2.1}$$

$$i^\star_{t+1} = \bar{\imath} + \phi_i(i^\star_t - \bar{\imath}) + \varepsilon_{i^\star_{t+1}}, \tag{2.2}$$

where $\bar{\imath} > 0$, $\phi_i \in (0,1)$, $\text{var}_t[\varepsilon_{i_{t+1}}] = \text{var}_t[\varepsilon_{i^\star_{t+1}}] = \sigma_i^2 > 0$, and $\text{corr}_t[\varepsilon_{i_{t+1}}, \varepsilon_{i^\star_{t+1}}] = \rho \in [0,1]$. That is, we allow interest rate shocks to be correlated.

Long-term bonds Long-term bonds are default-free perpetuities whose payments decline geometrically. The domestic currency return on long-term domestic bonds from t to $t+1$ is

$$r^y_{t+1} = y_t - \frac{\delta}{1-\delta}(y_{t+1} - y_t) - g_t, \tag{2.3}$$

where y_t is the log yield-to-maturity on domestic long-term bonds, and parameter $\delta \in (0,1)$ is the bond's payment decline rate (see Appendix B.1.1). Equation (2.3) expresses the bond's return as consisting of three components: (i) a carry component, (ii) a capital gain component, and (iii) a stochastic, time-varying wedge to bond returns, respectively. The last term introduces a time-varying wedge in the price of domestic long-term bonds without directly influencing the fundamental value of the exchange rate. It can be interpreted as a time-varying tax on local-currency bond returns, or a time-varying convenience yield. It

follows an AR(1) process:

$$g_t = \phi_g g_{t-1} + \varepsilon_{g_t}, \tag{2.4}$$

where $\phi_g \in [0,1)$, and $\text{var}_{t-1}[\varepsilon_{g_t}] = \sigma_g^2 \geq 0$.

Let rx_{t+1}^y denote the excess return on domestic bonds, that is the return on the "yield-curve trade"—the trade that borrows short-term and lends long-term in domestic currency. Iterating equation (2.3) forward and taking expectations, one obtains:

$$y_t = (1-\delta) \sum_{j=0}^{\infty} \delta^j \, \text{E}_t[i_{t+j} + rx_{t+j+1}^y + g_{t+j}], \tag{2.5}$$

which indicates that the domestic long-term yield can be decomposed into the standard expectations hypothesis and term premium components, as well as a term that accounts for all current and future expected wedge ("taxes") on bond returns. The excess return on domestic bonds rx_{t+1}^y is determined in equilibrium.

Foreign exchange Let q_t denote the log nominal exchange rate, expressed as units of home currency per unit of foreign currency, in the spot market. The log excess return on the "UIP trade," the trade that borrows short-term in foreign currency and lends short-term in domestic currency, is:

$$rx_{t+1}^q = (i_t - i_t^\star) - (q_{t+1} - q_t). \tag{2.6}$$

Iterating this expression forward and taking expectations yields:

$$q_t = \sum_{j=0}^{\infty} \text{E}_t[(i_{t+j}^\star - i_{t+j}) + rx_{t+j+1}^q] + \lim_{k \to \infty} \text{E}_t \, q_{t+k}. \tag{2.7}$$

Equation (3.11) expresses the spot exchange rate as the sum of three components: (i) the expected future sum of interest rate differentials, *i.e.* the UIP component; (ii) a foreign

exchange risk premium; and (iii) the expected long-run nominal exchange rate. We assume that the long-term nominal exchange rate, $\bar{q}_t \equiv \lim_{k\to\infty} \mathrm{E}_t\, q_{t+k}$, is time-varying and follows a random walk. That is:

$$\bar{q}_t = \bar{q}_{t-1} + \varepsilon_{q_t}, \tag{2.8}$$

where $\mathrm{var}[\varepsilon_{q_t}] = \sigma_q^2 \geq 0$. Time variation in the long-term value of the nominal exchange rate introduces an additional, yet independent, source of risk for the UIP trade. While it is unmodeled here, time variation in the value of the long-term nominal spot exchange rate arises naturally in open-economy macroeconomic models with cross-country differences in inflation rates, while maintaining real exchange rate stationarity.

Intermediaries A subset of agents can trade all assets and enforce no arbitrage. Anticipating our empirical findings in Section 2.4, we refer to these agents as "intermediaries" or "domestic intermediaries." These agents maximize their next-period wealth through mean-variance preferences with a risk tolerance parameter τ. Let d_t^y denote the market value of the intermediary' holdings of long-term domestic bonds and d_t^q denote the value of the net intermediary's position in the borrow-foreign and lend-domestic UIP trade, all denominated in domestic currency. Defining $\mathbf{d}_t \equiv [d_t^y, d_t^q]'$ and $\mathbf{rx}_{t+1} \equiv [rx_{t+1}^y, rx_{t+1}^q]'$, intermediaries solve:

$$\max_{\mathbf{d}_t} \left\{ \mathbf{d}_t' \,\mathrm{E}_t[\mathbf{rx}_{t+1}] - \frac{1}{2\tau} \mathbf{d}_t' \,\mathrm{var}_t[\mathbf{rx}_{t+1}] \mathbf{d}_t \right\}, \tag{2.9}$$

Taking first-order condition yields the optimality condition faced by intermediaries:

$$\mathrm{E}_t[\mathbf{rx}_{t+1}] = \tau^{-1} \,\mathrm{var}_t[\mathbf{rx}_{t+1}] \mathbf{d}_t. \tag{2.10}$$

Equation (2.10) links the intermediaries' expected excess returns on different assets to their asset holdings. The variance-covariance matrix of excess returns governs the equi-

librium relationship of excess returns across the yield-curve trade and the UIP trade. In Appendix B.1.2, we show that d_t^q consists of the sum of long spot and long forward positions, and thus indeed represents the net position of the intermediaries on the UIP trade.

Domestic and foreign investors' bond and currency demand In each period, domestic long-term bonds are available in a given net supply, denoted as s_t^y, which is equal to their gross issuance minus the demand from bond investors. The net supply of domestic long-term bonds follow AR(1) process:

$$s_{t+1}^y = \phi_{s^y} s_t^y + \varepsilon_{s_{t+1}^y}, \tag{2.11}$$

where $\phi_{s^y} \in [0,1)$, $\varepsilon_{s_{t+1}^y} = \eta_{t+1} + \eta_{t+1}^\star$, $\text{var}_t[\varepsilon_{s_{t+1}^y}] = \sigma_\eta^2 + \sigma_{\eta^\star}^2 \geq 0$. We distinguish here between demand shocks coming from domestic investors (fluctuations in η_{t+1}) and foreign investors (fluctuations in $\eta_{t+1}^\star$) and we allow them to have different stochastic properties.

Let the net supply of home currency in the spot market (that is, net of the demand from domestic and foreign investors), denominated in units of domestic currency, follow a stochastic process such that

$$s_{t+1}^x = \phi_{s^q} s_t^x + \varepsilon_{s_{t+1}^x}, \tag{2.12}$$

where $\phi_{s^q} \in [0,1)$, $\varepsilon_{s_{t+1}^x} = \gamma_{t+1} + \gamma_{t+1}^\star$, and $\text{var}_t[\varepsilon_{s_{t+1}^x}] = \sigma_\gamma^2 + \sigma_{\gamma^\star}^2 \geq 0$, distinguishing between demand shocks coming from domestic investors, γ_{t+1}, and foreign investors, $\gamma_{t+1}^\star$.

Without loss of generality, we assume throughout that, within each investor group, bond and currency demand volatilities are identical:

$$\phi_{s^y} = \phi_{s^q} \in [0,1); \qquad \sigma_\eta^2 = \sigma_\gamma^2 \geq 0; \qquad \sigma_{\eta^\star}^2 = \sigma_{\gamma^\star}^2 \geq 0. \tag{2.13}$$

Departing from the framework of Greenwood et al. (2023a), and in line with the evidence

of Section 2.2, we assume that any purchase of domestic long-term bonds $\eta^\star_{t+1}$ by foreign bond investors is accompanied by a simultaneous purchase of home currency in the spot market $\gamma^\star_{t+1}$—of same local-currency amount. That is:

$$\text{cov}_t(\eta^\star_{t+1}, \gamma^\star_{t+1}) = \sigma^2_{\eta^\star} = \sigma^2_{\gamma^\star} \geq 0; \qquad \text{corr}_t(\eta^\star_{t+1}, \gamma^\star_{t+1}) = 1. \tag{2.14}$$

This correlation arises naturally, as we argued in Section 2.2, as foreign investors intending to purchase domestic long-term bonds must inevitably acquire a corresponding amount of the home currency.

To the contrary, bond and currency demand of domestic investors are uncorrelated:

$$\text{cov}_t(\eta_{t+1}, \gamma_{t+1}) = 0; \qquad \text{corr}_t(\eta_{t+1}, \gamma_{t+1}) = 0. \tag{2.15}$$

Currency hedging We assume that a constant fraction $\alpha \in [0,1]$ of foreign investor bond purchases are hedged via forward contracts. Specifically, foreign investors always purchase the full amount of domestic currency in the spot market to acquire local-currency bonds, and hedge a share α of this position by selling domestic currency in the forward market.

Let the net supply of home currency in the forward market (net of the demand from foreign investors), denominated in units of domestic currency, follow

$$s^f_{t+1} = \phi_{s^q} s^f_t + \varepsilon_{s^f_{t+1}} \tag{2.16}$$

where $\varepsilon_{s^f_{t+1}} = -\alpha\gamma^\star_{t+1}$ and thus $\text{var}_t[\varepsilon_{s^f_{t+1}}] = \alpha^2\sigma^2_{\gamma^\star}$. Intermediaries absorb flows both in the spot and forward FX markets. Taken together, the overall net supply of home currency (net of the demand from domestic and foreign investors), denominated in units of domestic currency, is

$$s^q_{t+1} = s^x_{t+1} + s^f_{t+1} = \phi_{s^q} s^q_t + \varepsilon^q_{t+1} \tag{2.17}$$

where $\varepsilon^q_{t+1} = \gamma_{t+1} + (1-\alpha)\gamma^\star_{t+1}$, and $\text{var}_t[\varepsilon_{s^q_{t+1}}] = \sigma^2_\gamma + (1-\alpha)^2\sigma^2_{\gamma^\star}$.

Market clearing The market clearing condition ensures that intermediaries absorb demand across bond and currency markets. That is:

$$\mathbf{s}_t = \mathbf{d}_t, \tag{2.18}$$

where $\mathbf{s}_t = [s^y_t, s^q_t]'$ denotes the vector of net supplies.

Correlated positions A key observation of our analysis is that foreign investors' local-currency bond demand induces correlated changes in intermediaries' bond and currency positions—and thus to the exposure to bond and currency risk. Formally, the correlation in intermediaries' positions is:

$$\Omega \equiv \text{corr}(d^y_t, d^q_t) = \frac{\text{cov}[s^y_t, s^q_t]}{\sqrt{\text{var}(s^y_t)}\sqrt{\text{var}(s^q_t)}} = \frac{(1-\alpha)\sigma^2_{\eta^\star}}{\sqrt{(\sigma^2_\eta + \sigma^2_{\eta^\star})}\sqrt{(\sigma^2_\eta + (1-\alpha)^2\sigma^2_{\eta^\star})}}, \tag{2.19}$$

which can be derived from equations (2.11)-(2.18).

The correlation in intermediaries positions, labeled Ω, is governed by the relative importance of unhedged foreign investors' local-currency bond demand. As highlighted by equation (2.19), both the relative volatility of foreign investors' demand—rather than their average portfolio share—and their propensity to hedge currency risk matter for this relationship. When foreign investors play a relatively greater role in aggregate flow volatility, the correlation in flows across markets, and thus in intermediaries' positions, is higher. The extent of hedging by foreign investors also shapes this correlation. Comovement is strongest when foreign investors do not hedge their FX positions ($\alpha = 0$). Conversely, when foreign investors fully hedge their exposure to currency risk ($\alpha = 1$) their bond purchases generate no net currency pressure, eliminating any correlation in intermediaries

bond and currency positions.

Equilibrium Equilibrium expected excess returns must satisfy the intermediaries' optimality condition (2.10) as well as the market clearing condition (2.18), implying:

$$\mathrm{E}_t[\mathbf{rx}_{t+1}] = \tau^{-1}\,\mathrm{var}_t[\mathbf{rx}_{t+1}]\mathbf{s}_t. \tag{2.20}$$

To pin down equilibrium bond yields and exchange rates, y_t and q_t, we follow Greenwood et al. (2023a) and conjecture that prices are linear functions of the state vector $\mathbf{z}_t$, which contains all stochastic processes, $\mathbf{z}_t = [i_t - \bar{\imath}, i_t^\star - \bar{\imath}, g_t, \bar{q}_t, s_t^y, s_t^q]'$. Appendix B.1.3 contains the full mathematical solution, provides a characterization of equilibrium bond yield and exchange rate, as well as a discussion on equilibrium multiplicity and selection.

2.4 Investors' flows and intermediaries' positions

We now outline several model predictions, and confront them with evidence from Colombia's bond and foreign exchange markets. Section 2.4.1 explores position and price dynamics during a period of foreign investor inflows into the Colombian local-currency sovereign bond market. Among several findings, this episode supports the view that intermediaries absorb imbalances across both markets and prices adjust accordingly. To further substantiate this interpretation, Section 2.4.2 analyzes the positions and returns of intermediaries and shows that they can indeed be regarded as the players responsible for enforcing no-arbitrage. Throughout, "intermediaries" are the Colombian intermediaries that are designated dealers in bond and currency markets, "foreign investors" are investors domiciled outside of Colombia, and "domestic investors" are the domestic financial market participants other than intermediaries.

The key friction in the model is intermediaries' limitied risk-bearing capacity. This links intermediaries positions to expected returns. For tractability, we derive our main analytical results under a simplified version of the model. In particular, we abstract from interest-rate shocks ($\sigma_i = 0$) to isolate the effects of correlated foreign-investor flows. We abstract from currency risk hedging ($\alpha = 0$), so foreign investors' purchases of local-currency bonds translate fully into simultaneous flows in the currency market. We take other special cases to obtain a sharp closed-form characterizations.

Assumption 1. *Short-term interest rates are deterministic ($\sigma_i = 0$), foreign investors do not hedge currency risk ($\alpha = 0$), asset-specific shocks are transitory ($\phi_g = 0$) and share the same volatility ($\sigma_g = \sigma_q$), and net-supply shocks are transitory ($\phi_{s_y} = \phi_{s_q} = 0$). In addition, long-term bonds are assumed to have near-infinite duration ($\delta \to 1$), and agents are sufficiently risk-tolerant (i.e., τ is large enough).*

2.4.1 The effects of foreign investors' bond purchases

We characterize the effects of a sudden increase in foreign investors' bond purchases.

Proposition 1. *Under Assumption 1, foreign investors' purchases of local-currency bonds causes*

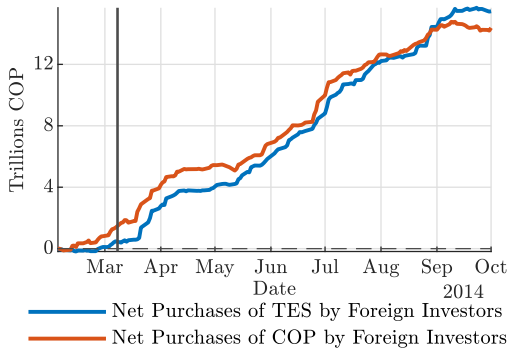
(a) a simultaneous acquisition of domestic currency and domestic bonds by foreign investors;

(b) a decline in intermediaries' bond and currency positions;

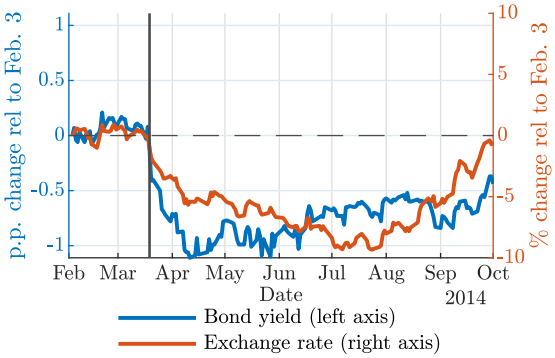
(c) an increase in long-term bond prices (a decline in bond yields for finite duration) and home currency appreciation.

Proof. See Appendix B.1.4. □

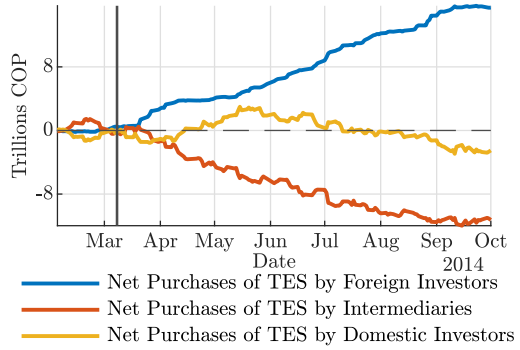
Figure 2.2: Positions and prices during J.P. Morgan Index Rebalancing



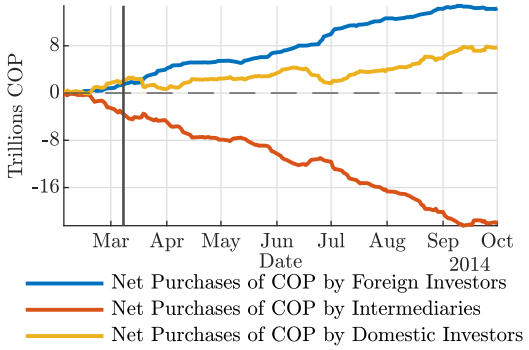
(a) Foreign investors' flows



(b) Bond price and exchange rate



(c) Intermediaries' TES holdings



(d) Intermediaries' COP holdings

Note: Figure 2.2a displays the evolution of the daily accumulated foreign purchases of TES bonds as well as with the daily accumulated foreign purchases of COP in the spot market during 2014. Figure 2.2c and Figure 2.2d display the evolution of the daily accumulated purchases of TES bonds and COP in the spot market, respectively, during 2014 for different investor groups (described in Section 2.2 and Section 2.4). The black vertical line denotes March 19, 2014, the day the rebalancing was announced.

This proposition embodies a key mechanism of the model. When foreign investors buy local-currency bonds, they exchange short-term dollars for long-term pesos (they obtain short-term pesos and immediately use them to purchase long-term peso bonds). As intermediaries sell these bonds for dollars, they now have a shorter position on both the yield-curve trade, and on the UIP trade. Consequently, they require a downward shift in the premium on long-term bonds relative to short-term bonds, and on pesos relative to dollars. In equilibrium, long-term bond yields fall and the home currency appreciates.

We confront these predictions against a major episode of foreign investor inflows into Colombia's local-currency government bond market (TES). In March 2014, J.P. Morgan announced the inclusion of five TES bonds in its flagship emerging-market local-currency index. The announcement triggered large index-driven purchases by foreign investors: by the time inclusion was completed in October 2014, foreign investors had acquired nearly 10% of outstanding TES bonds.[7] These purchases were mirrored almost one-for-one by foreign acquisitions of Colombian pesos (COP) in the spot FX market (Figure 2.2a). Consistent with Proposition 1, although the shock originates in the bond market, it simultaneously generates flows in the FX market because it is funded in dollars.

During this episode, Colombian intermediaries largely absorbed the resulting imbalances in both the bond and spot exchange markets (Figure 2.2c, Figure 2.2d); a similar pattern appears when considering net spot purchases minus forward sales (Appendix Figure B.4). Thus, the foreign investors purchases significantly changed intermediaries' bond and currency positions.

Price dynamics also align with the model's mechanism. Figure 2.2b reveals that, following the announcement, long-term yields fell and the peso appreciated. These patterns are consistent with the notion that intermediaries required lower future returns on both yield-curve and UIP trades as their exposures to these strategies' underlying risk has declined.

2.4.2 Intermediaries: positions and returns

We now document that the portfolios of intermediaries exhibit correlated bond and currency positions. We also show that intermediaries' positions and portfolio returns suggest that they indeed demand compensation for bearing bond and currency risks.

[7] Williams (2018) uses Colombia's inclusion in J.P. Morgan's emerging markets debt index to study how increased foreign access to sovereign debt markets boosts private credit availability.

Correlated positions of intermediaries In the model, foreign investors generate correlated exposures on the balance sheets of intermediaries.

Lemma 1. *As long as foreign investors participate in local-currency bond markets and that they do not hedge perfectly, intermediaries feature correlated bond and currency exposures.*

Proof. It follows from equation (2.19). If $\sigma^2_{\eta^\star} > 0$ and $\alpha < 1$, then $\Omega > 0$. □

We now examine whether Colombian intermediaries exhibit a positive correlation between their holdings of long-term bonds (TES) and Colombian pesos (COP). For the group of intermediaries, we compute holdings of TES and COP, both measured in domestic currency. For simplicity, we focus on intermediaries' spot FX exposures. This is motivated by the limited role of the forward market during the 2014 capital inflow episode (see Appendix Figure B.4) and, more generally, by the fact that foreign investors' forward FX transactions appear largely uncorrelated with their spot FX transactions (see Appendix Figure B.5). We scale intermediaries' TES and COP positions by total TES outstanding and by M1, respectively.[8]

Consider the following regression:

$$\text{COP}_t = \alpha + \beta \text{TES}_t + \varepsilon_t \tag{2.21}$$

where COP_t and TES_t are spot and bond holdings by intermediaries. We perform this regression at both monthly and quarterly frequencies and report the estimated coefficients in Table 2.1. The estimates indicate a positive relationship between intermediaries bond and currency positions, with a correlation around 0.4. Overall, the results support the view that correlated positions are a salient feature of intermediary portfolios.

[8]TES holdings are taken directly as a share of total TES outstanding. COP holdings are constructed by taking total liquid assets in the financial system as of March 2012 (the start of our sample), accumulating monthly FX spot purchases, and normalizing by M1 money supply in Colombia.

Table 2.1: Portfolio positions of intermediaries

	COP holdings	
	Monthly	Quarterly
TES holdings	2.23***	1.99**
	(0.72)	(0.81)
Correlation	0.45	0.40
Observations	95	32
R-Squared	0.20	0.16

Note: The table presents OLS estimates of regression (2.21) of COP holdings on TES holdings by intermediaries. Driscoll-Kraay robust standard errors are reported in parenthesis. *, **, and *** denote statistically significant at the 10%, 5%, and 1% level, respectively.

Intermediaries' portfolios and returns A key feature of models with frictional financial intermediation is that intermediaries—the agents that enforce no-arbitrage—require positive expected excess returns to absorb any imbalances in bond and foreign exchange markets. The following proposition illustrates how this principle operates in our model.

Proposition 2. *Under Assumption 1, intermediaries earn positive excess returns, on average, on both their bond and currency trades. That is,* $\mathbb{E}(\Delta d_t^y rx_{t+1}^y) \geq 0$ *and* $\mathbb{E}(\Delta d_t^q rx_{t+1}^q) \geq 0$*, where* rx_{t+1}^y *and* rx_{t+1}^q *denote, respectively, the one-period excess returns on the bond and currency positions, and* $\mathbb{E}$ *is the unconditional expectation operator.*

Proof. See Appendix B.1.4. □

We test whether intermediaries' trades earn positive excess returns in the two markets by constructing their monthly-level trade returns, effectively the realized excess returns associated with their marginal trades. Specifically, the UIP trade returns are computed as:

$$\text{UIP Trade Returns}_{t,h} = \Delta d_t^q \times rx_{t+h}^q, \qquad h = 1, 3, 6, 12, \tag{2.22}$$

where Δd_t^q is a change in intermediary position associated with selling USD to buy COP (measured in USD), and rx_{t+h}^q denotes excess currency returns over a holding period of h

Table 2.2: UIP and Yield-Curve Trade Returns on Intermediaries' Portfolios

	UIP Trade Returns				Yield-Curve Trade Returns			
Horizon	1M	3M	6M	12M	1M	3M	6M	12M
	3.4	15.3**	38.8***	74.4***	4.5	13.2**	18.6***	24.9***
	(2.4)	(7.0)	(11.8)	(21.6)	(5.2)	(6.1)	(7.0)	(8.2)
Observations	83	81	78	72	120	120	120	120

Note: This table reports results from mean t-tests of UIP trade and yield-curve trade returns for intermediaries. UIP trade returns are computed following equation (2.22) and yield-curve trade returns are computed using equation (2.23). The 1M, 3M, 6M, and 12M columns denote test for average returns with holding periods of 1, 3, 6, and 12-month, respectively. The null hypothesis is that average returns are equal to zero. *, **, and *** denote statistically significant at the 10%, 5%, and 1% level, respectively.

months.

Intermediaries' yield-curve trade returns are computed as:

$$\text{Yield-Curve Trade Returns}_{t,h} = \Delta d_t^y \times rx_{t+h}^{(n)}, \qquad h = 1,3,6,12, \tag{2.23}$$

where Δd_t^y is the change in intermediaries' position towards the acquisition of long-term domestic government bonds, and $rx_{t+h}^{(n)}$ denotes the corresponding excess bond returns. For simplicity, we use the 10-year local-currency bond and measure returns in domestic currency. After calculating trade returns for each holding period, we perform mean t-tests to assess whether the returns are significantly different from zero.

Table 2.2 reveals that UIP and yield-curve trade returns are, on average, positively related to changes in intermediaries' positions. Returns are noisier for 1-month horizon, but all average returns are statistically significant for 3-, 6- and 12-month holding periods. Consistent with Proposition 2, these findings suggest that intermediaries—both primary dealers and FX market intermediaries—are the agents enforcing no-arbitrage. This evidence aligns with Haddad and Sraer (2020), who show that, for the United States, banks' interest rate

exposure predicts U.S. bond returns.[9]

2.5 Bond yields and exchange rates, and asset purchase policies

We now characterize the equilibrium relationship between bond yields and exchange rates, focusing on the correlation between excess bond returns and excess currency returns, and quantify the role of unhedged foreign investor bond demand in the observed correlation in Colombia's bond and currency returns (Section 2.5.1). We then derive the resulting implications for the transmission of asset purchase policies and confront them with identified effects of government bond issuance and dollar purchases (Section 2.5.2).

2.5.1 Equilibrium bond and currency returns

Our mechanism implies a tight connection between the presence of unhedged foreign investor flows and the correlation in bond and currency returns. We next formalize this connection in the simplified version of our model.

Proposition 3. *Under Assumption 1,* $\text{corr}_t(rx^y_{t+1}, rx^q_{t+1}) > 0$ *if and only if* $\Omega > 0$.

Proof. See Appendix B.1.4. □

Absent interest rate risk, the correlation between excess bond and currency returns is positive whenever unhedged foreign investors operate in the local-currency bond market, and thus risk exposures are positively correlated. Absent interest rate shocks, the fundamental sources of risk are asset-specific *independent* shocks (fluctuations in the bond return wedge and long-run nominal exchange rate). If demand fluctuations were independent, term premia and currency premia would also be uncorrelated. However, when demand shocks are

[9] For some other emerging economies, other studies that highlight the role of local banks in financial intermediation and asset pricing include di Giovanni et al. (2021); Gutierrez et al. (2023); Keller (2024); De Leo et al. (2024c); Fendoglu et al. (2019); Gonzalez et al. (2021); De Leo et al. (2024b).

correlated, as it is the case in the presence of unhedged foreign investors, their comovement is reflected in the joint distribution of premia.[10] Correlated positions thus beget correlated returns as they directly link the net supply in bond and currency markets that intermediaries must absorb, and, as a result, the premia they require. (Note that this effect emerges even if intermediaries managing bonds and currency were distinct entities.)

We now turn to examine the determinants of the bond-currency return comovement in a version of the baseline model that is calibrated to Colombia, and that accounts for short-term interest rate risk.

Calibration Our calibration approach does not target directly the empirical comovement in excess bond and currency returns. Instead, it relies on the time series properties of short-term interest rates and net supply volumes, as well as the standard deviations of excess bond returns and excess currency returns. The model is calibrated on Colombian data at a quarterly frequency. Table B.1 reports the values of the baseline parameters. We use 3-month interbank rates from Colombia and the United States to discipline the processes of short-term interest rates (ϕ_i, σ_i, ρ). We construct quarterly bond positions of market participants to discipline the persistence of the net supply shocks ($\phi_{s^y} = \phi_{s^q}$) and the volatility of domestic and foreign investors' bond demand shocks ($\sigma_\eta = \sigma_\gamma$ and $\sigma_{\eta^\star} = \sigma_{\gamma^\star}$).[11] The symmetric calibration of net supply processes across bond and currency markets (following eq. (2.13)) appears a reasonable approximation of the data. We set the risk aversion of intermediaries ($1/\tau$) to 0.03, in line with the range of values chosen by Gourinchas et al. (2022b). We calibrate the unobservable asset-specific shocks (σ_g, σ_q) to match the standard deviation of excess 10-year Colombian bond returns and excess currency returns.

[10] Gabaix and Maggiori's (2015) model illustrates how capital flows into *short-term* local-currency bond markets can cause currency appreciations. We argue that these capital inflows occur primarily in the *long-term* local-currency bond markets, not only changing the net supply of different currencies, but also changing the net supply of local-currency bonds of different maturities.

[11] The TES (COP) holdings are computed as a fraction of the total value of TES outstanding bonds (M1 money aggregate).

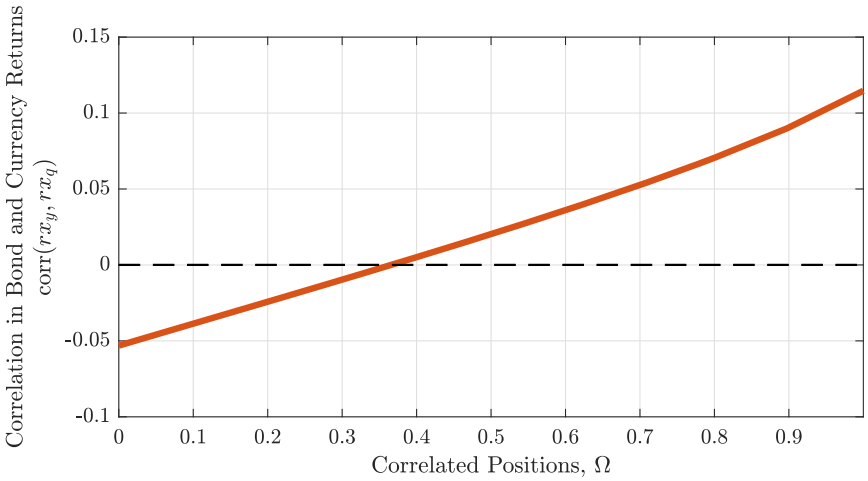
In this baseline calibration we set the currency-risk hedge ratio, α, to 0. We argue that this is a valid approximation of the data. Appendix Figure B.5 documents that there is no consistent relationship between foreign investors net purchases of COP in the spot market and their net sales of COP in the forward market. This indicates that foreign investors do not systematically hedge their Colombian bond positions.

Quantifying the role of correlated positions To quantify the contribution of correlated intermediaries' positions to the comovement of bond and currency excess returns, we compare two models that differ in the degree of correlated positions (see Ω in (2.19)). The first calibration corresponds to the empirically relevant level of correlated net supply fluctuations for Colombia, $\Omega = 0.46$. The second is a counterfactual economy in which foreign investors' flows are uncorrelated across markets, implying $\Omega = 0$, while holding all other parameters fixed, *i.e.*, leaving the overall volatility of demand shocks unchanged.

Figure 2.3 depicts the correlation in bond and currency returns as a function of the extent of correlated positions in intermediaries' portfolios. In the counterfactual economy without correlated positions, the model implies a negative correlation between excess bond and currency returns (around -6%). In this case, as in Greenwood et al. (2023a), short-term interest rate shocks are the only set of shocks that induce correlated risk, and yield-curve and UIP trades load on these shocks with opposite signs. An increase in domestic short-term rates results in lower domestic long-term bond prices, which is detrimental to the yield-curve trade, but it simultaneously appreciates the domestic currency, benefiting the UIP trade.[12] By contrast, in the baseline model with $\Omega = 0.46$, correlated net supply fluctuations introduce a source of risk that moves term premia and currency premia in the same direction (around 2%), overturning the negative comovement generated by interest

[12] Short-rate risk only influences the correlation in returns if short-rate fluctuations are not perfectly correlated across countries. Correlated short-rate fluctuations limit the variation in the short-rate *differential*, thereby reducing the impact of interest rate risk on the UIP trade.

Figure 2.3: Asset positions and asset returns



Note: The figure depicts the equilibrium correlation in bond and currency excess returns for different levels of correlation in net supplies, $\Omega \equiv \text{corr}(s_t^y, s_t^q)$ (achieved by different relative importance of foreign to home standard deviations of supply shocks, see eq. (2.19)). Table B.1 reports the baseline calibration.

rate risk alone.

Thus, the difference in correlations between these two calibrations of the model ($\Omega = 0$ and $\Omega = 0.46$) isolates the portion of bond-currency comovement attributable to unhedged foreign investors flows (and thus to correlated positions). Quantitatively, the increase in the correlation of excess returns when moving from the counterfactual to the baseline equilibria accounts for approximately half of the empirical bond-currency return correlation (which is around 14%). That is, nearly half of the observed comovement between excess bond and currency returns can be explained by correlated demand fluctuations associated with foreign investor flows.

2.5.2 Asset purchase policies, bond yields and exchange rates

The equilibrium variance-covariance of bond and currency returns is a key determinant of the effects government's asset purchases or sales, *i.e.*, policies that alter the position of intermediaries in either the bond or currency market. This class of policies, including quantitative easing/tightening and foreign exchange interventions, are increasingly preva-

lent across economies. The following proposition formalizes the effects of these policies on bond yields and exchange rates.

Proposition 4.

(a) A government's purchase of local-currency long-term bonds, modeled as a reduction in η_t in equation (2.11), increases the price of long-term bonds. In addition, it leads to an appreciation of the home currency if and only if $\text{cov}_t(rx^y_{t+1}, rx^q_{t+1}) > 0$.

(b) A domestic central bank's purchase of foreign currency, modeled as an increase in γ_t in equation (2.17), leads to a depreciation of the home currency. In addition, it causes a decline in the price of long-term bonds if and only if $\text{cov}_t(rx^y_{t+1}, rx^q_{t+1}) > 0$.

Proof. See Appendix B.1.4. □

Even if government bond purchases/sales are confined to a single market—the government bond market—they affect not only equilibrium bond yields but also exchange rates. Likewise, central bank purchases/sales of foreign currency affect not only the exchange rate but also the price of long-term bonds. These "cross-market" effects arise from two central features of our small open economy framework and are shaped by the structure of shocks.

First, intermediaries operate in both the bond and currency markets, which are exposed to correlated risks. As a result, a shift in the net supply of one asset effectively induces revisions in both term premia and currency premia, as it alters the intermediaries' exposure to correlated bond and currency risks. Second, the magnitude and direction of these effects depend on the stochastic properties of bond and currency returns, captured by their covariance. As discussed in Section 2.5.1, this covariance reflects the relative importance of unhedged foreign investor flow shocks versus short-term interest rate shocks. In economies where unhedged foreign investor flows are dominant, the risks associated with the UIP trade and the yield-curve trade are positively correlated, and so is the response of bond yields and

exchange rates to asset purchases/sales.[13]

We now turn to the empirical evidence on these predictions.

Empirical setting We consider a set of auctions conducted by either the Treasury or the Central Bank in Colombia. As we describe below, these auctions allow us to identify the effects of asset purchases or sales that alter the net supply of intermediaries in either the bond market or foreign exchange markets. As we explain below, we restrict this analysis to the sample period from March 2012 through December 2014.

Treasury auctions in the local-currency bond market The Treasury conducts weekly auctions of local-currency government bonds (TES) as the core instrument of domestic public debt financing.[14] Participation is restricted to institutions included in the Ministry of Finance's Primary Dealer Program for Public Debt Securities, whose members are designated on a fiscal-year basis. In exchange for privileged access to primary auctions, these institutions, the "intermediaries", are required to underwrite a minimum share of total annual issuance.

Our sample period (March 2012 to December 2014) comprises 491 auctions in which the Treasury issued TES at standard benchmark maturities of 1, 5, and 10 years, corresponding to the short-, medium-, and long-term segments of the domestic yield curve. In practice, issuance is concentrated through repeated re-openings of existing benchmark securities, which enhances liquidity and facilitates direct comparability between primary issuance and secondary market trading. Secondary market transactions take place through

[13] In this model, asset purchases/sales should be interpreted as a government interventions carried out under the full variance-covariance structure implied by all exogenous processes, rather than in a counterfactual setting where these interventions are the sole source of variation.

[14] These auctions serve a dual role: they ensure the placement of government securities needed to finance the annual budget, while supporting liquidity in the secondary market. Issuance takes place within the annual financing framework defined by the budget law and the government's financial plan for each fiscal year.

the electronic platforms described in Section 2.2, allowing us to link auction outcomes with contemporaneous and subsequent market prices.

TES auctions follow a multi-unit, uniform-price format. Prior to each auction, the Treasury announces the total amount to be issued along with the characteristics of the security. Bids are submitted within a fixed 30-minute window (9:30–10:00 a.m.), during which authorized participants can submit and revise price–quantity pairs. Each participant may submit up to five bids. At the close of the auction, bids are ranked in descending order of price, and allocations proceed sequentially until the preannounced issuance amount is exhausted. All winning bids are executed at a common clearing price (the cutoff price).

On average, each auction involves approximately ten participating institutions, predominantly private commercial banks, with an average of 3.7 bids per participant. This relatively concentrated but competitive structure mirrors the institutional setup of other sovereign bond markets. For context, the outstanding stock of TES at end-2014 amounted to approximately COP 192.9 trillion (about USD 100 billion), representing the bulk of central government domestic debt (see Table 2.3 for additional descriptive statistics on treasury auctions).

Central bank auctions in the foreign exchange market The central bank conducted foreign exchange auctions as part of a daily preannounced program of dollar purchases aimed at accumulating international reserves. Although such operations were conducted between 2008 and 2014, we focus on the 2012–2014 period, when auctions became effectively daily and when detailed secondary-market transaction data are available, allowing us to link auction outcomes to market prices.[15]

We analyze proprietary data covering 650 multi-unit, uniform-price foreign exchange

[15] Auctions were intermittent prior to 2012 (June 24–October 6, 2008; March 3–June 30, 2010; January 3–September 30, 2011). Continuous daily operations begin in 2012, coinciding with the availability of secondary market data.

auctions conducted by the central bank. Prior to each auction, the central bank announced a purchase quota for U.S. dollars, exactly two minutes before the auction began. This announcement was largely operational, as it implemented a stable purchase pace maintained over extended periods and already anticipated by market participants, rather than reflecting a discretionary day-by-day choice. At the same time, it conveyed the timing of the operation within the day. Realized purchases could fall short if submitted offers were insufficient, although in practice allocations averaged 99.2% of the announced quota. Auctions were conducted over a three-minute window, during which authorized financial institutions could submit and revise bids specifying both a price (COP/USD) and a quantity of dollars offered. Each participant was allowed to submit a single bid. At the close of the auction, bids were ranked in ascending order of price, and allocations were made sequentially starting from the lowest-priced offers until the preannounced purchase amount was exhausted. All accepted bids were settled at a common clearing price, defined as the highest accepted ask price (the cutoff price), consistent with a uniform-price auction mechanism.

On average, each auction involved around eight participating financial institutions, primarily private commercial banks drawn from the set of authorized dealers. The central bank purchased an average of USD 23 million per day, with purchases occasionally reaching USD 50 million. For context, during the same period the average daily turnover in the COP/USD spot market was approximately USD 950 million, with an average individual transaction size of about USD 0.785 million (see Table 2.3 for additional descriptive statistics on central bank auctions).

Identification We identify the effects of the auction by exploiting the discontinuity induced by the cutoff price, comparing the subsequent trading behavior of marginal winners and marginal losers in both bond and foreign exchange markets. The key assumption is that bidders close to the threshold—those who barely win and those who barely lose—

Table 2.3: TES Bonds and U.S. Dollar Auctions: Summary Statistics

	Mean	Std. Dev.	Min	P25	Median	P75	Max
Panel A: TES Bond Auctions (491 auctions, March 2012–December 2014)							
Number of Participants	10.3	3.3	1.0	8.0	11.0	13.0	17.0
Winning Participants	8.3	3.0	0.0	6.0	9.0	11.0	14.0
Losing Participants	2.0	2.5	0.0	0.0	1.0	3.0	13.0
Bids per Auction	37.5	27.4	1.0	11.0	38.0	53.0	153.0
Winning Bids per Auction	17.0	11.1	0.0	9.0	15.0	22.0	73.0
Losing Bids per Auction	20.5	19.7	0.0	0.0	20.0	33.0	127.0
Cutoff Yield	4.93	1.41	1.50	3.90	4.84	6.09	7.85
Bid Range per Auction	0.44	0.34	0.00	0.00	0.45	0.66	1.80
TES Purchase by Marginal Winning bid$^{(a)}$	4.66	5.97	0.02	0.57	2.61	5.22	112.60
Panel B: U.S. Dollar Auctions (650 auctions, March 2012–December 2014)							
Number of Participants	8.14	2.58	2.0	6.0	8.0	10.0	15.0
Winning Participants	4.91	2.10	1.0	3.0	5.0	6.0	11.0
Losing Participants	3.23	2.20	0.0	2.0	3.0	5.0	10.0
Bids per Auction	8.14	2.58	2.0	6.0	8.0	10.0	15.0
Winning Bids per Auction	4.91	2.10	1.0	3.0	5.0	6.0	11.0
Losing Bids per Auction	3.23	2.20	0.0	2.0	3.0	5.0	10.0
Cutoff Price	1,885	93.9	1,756	1,804	1,882	1,932	2,372
Bid Range per Auction	1.71	1.40	0.10	0.90	1.30	2.00	9.18
USD Purchase by Marginal Winning bid$^{(a)}$	4.69	4.53	0.10	1.50	3.00	6.60	32.00

Note: This table presents summary statistics for all 491 TES auctions and 650 U.S. Dollar auctions and in our sample. $^{(a)}$ Amounts are expressed in millions of U.S. dollars. In TES auctions, each dealer may submit up to five bids, whereas in U.S. dollar auctions only one bid is permitted. Aggregating at the dealer level, the average purchase by the marginal winning dealer in TES auctions is USD 9.5 million, nearly twice the corresponding amount in U.S. dollar auctions.

are ex ante similar. Under this assumption, differences in secondary-market asset prices within a narrow window around the cutoff reflect only the differential impact of the auction on winners' and losers' balance sheets. We provide evidence supporting the premise that marginal bidders are ex ante comparable. Table B.3 shows that participation and outcomes are not systematically different across institutions, whereas Table B.4 documents limited persistence in auction outcomes.

In foreign exchange auctions, the running variable—the variable determining treatment assignment—is the rank of each dealer's bid relative to the cutoff, with zero corresponding to the marginal winning bid and negative (positive) values indicating losers (winners). In bond auctions, where dealers can submit multiple bids, we define the running variable at the

dealer level using the most competitive bid and construct ranks analogously around the cut-off price. Under this definition, observations near the threshold reflect similar valuations, enabling a localized comparison of marginal winners and losers.

Formally, we estimate the following regression discontinuity design (RDD) centered around the auction cutoff price:

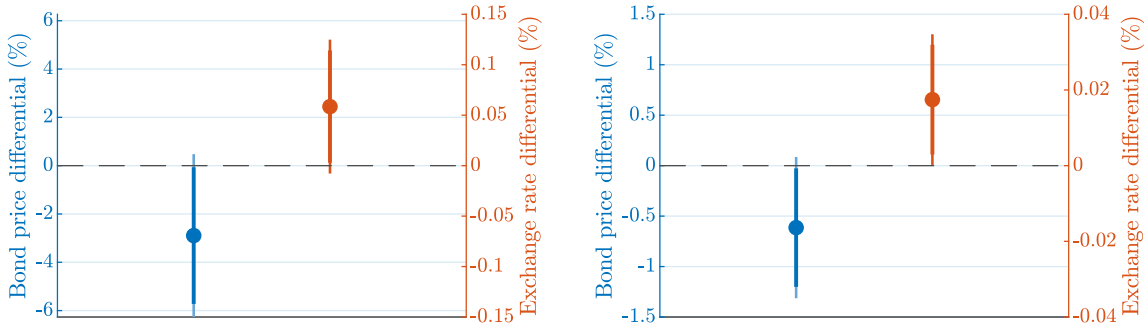
$$\arg\min_{\boldsymbol{\gamma}} \sum_{i=1}^{I} \left[y_i - \gamma_0 - \gamma_1 D_i - \gamma_2 (\text{bid}_i - c) - \gamma_3 (\text{bid}_i - c) \times D_i \right]^2 K\left(\frac{\text{bid}_i - c}{h} \right) \tag{2.24}$$

where y_i denotes either bond or dollar prices (in logs) in their respective secondary markets, measured immediately following the auction, but during the same trading day.[16] The variable $(\text{bid}_i - c)$ is the distance of bid i from the auction cutoff price c (positive for winners, negative for losers). The variable $D_i = \mathbb{1}\{\text{bid}_i \geq c\}$ is a treatment indicator equal to one if bid i was accepted; and $K(\cdot)$ is a triangular kernel function with bandwidth h. The parameter of interest in equation (2.24) is γ_1, as it captures the local average treatment effect of winning the auction on subsequent market outcomes.

Asset price effects of asset purchases/sales We document that asset purchase policies affect bond price and exchange rates in a manner consistent with Proposition 4. Figure 2.4 reports the estimates of γ_1 in (2.24), while Appendix Table B.5 reports the estimates of all coefficients of (2.24). The issuance of local-currency government bonds leads to an increase in bond yields *and* a depreciation of the Colombian peso against the U.S. dollar. Quantitatively, auction winners—dealers that purchase TES—subsequently trade bonds at prices 2.9% lower and dollars at prices 0.059% higher than auction losers. These magnitudes are economically large, corresponding to roughly one standard deviation of daily

[16] To avoid issues related to unit scaling over time, we express each secondary market trade as a % deviation (log-difference) from the median spot price of the corresponding day.

Figure 2.4: The responses of bond prices and exchange rates to asset purchases/sales



(a) Local-currency bond sales

(b) Foreign currency purchases

Note: This figure reports the estimates of γ_1 in equation (2.24), along with 90% (darker) and 95% (lighter) confidence intervals constructed using robust standard errors. All specifications use a bandwidth of two ranks, consistent with Calonico et al. (2014). Appendix Table B.5 reports the full set of estimates of equation (2.24).

changes in both exchange rate and bond prices.[17]

This pattern is consistent with Proposition 4(a), noting that it describes the effects of bond purchases rather than sales: intermediaries perceive bond and currency returns to be positively correlated. As they are required to absorb more local-currency government bonds, they demand not only a higher term premium, as in, *e.g.* Greenwood and Vayanos (2014b), but also a higher currency premium, leading in equilibrium to lower bond prices and a depreciation of the exchange rate.

Central bank purchases of U.S. dollars also have significant effects on both the exchange rate and bond prices. Dollar auction winners—dealers that sell USD to the central bank—subsequently trade FX at prices 0.017% higher and bonds at prices 0.61% lower than auction losers. These effects are economically meaningful, corresponding to roughly one-third of the standard deviation of daily exchange rate changes and one-fifth of the daily bond price changes.[18] These effects are consistent with the prediction in Proposition 4(b).

[17]The marginal winning bids amount to USD 4.66 million in TES purchases, and, accounting for multiple bids per dealer, this aggregates to roughly USD 9.5 million at the marginal winner level (see Table 2.3), corresponding to approximately 0.3–0.5% of average daily TES turnover.

[18]The marginal winners sell on average USD 4.69 million to the central bank (see Table 2.3), equivalent to roughly 0.5% of average daily FX turnover.

Because bond and currency returns are positively correlated, greater exposure to Colombian peso risk induced by winning the dollar auction leads intermediaries to demand not only a higher currency premium (a weaker peso) but also a higher term premium (lower long-term bond prices).

2.6 Foreign investors and asset returns across countries

Do countries with a higher prevalence of unhedged foreign investor flows exhibit a higher correlation between bond and currency returns? To answer this question, we construct a measure capturing the extent to which a country is exposed to unhedged foreign investor flows, and study whether cross-country variation in this measure is associated with differences in the joint behavior of bond yields and exchange rates.

2.6.1 Bond and currency returns across countries

We begin by extending the analysis of bond and currency returns in Lustig et al. (2019) to a larger panel of economies. We document that there is substantial amount of cross-country heterogeneity in correlation between excess bond returns and excess currency returns.

Definitions We denote $P_t^{(n)}$ as the price of a zero-coupon bond of maturity n in local-currency terms at time t, with the continuously compounding yield on this bond given by $y_t^{(n)} = -\frac{1}{n}\log P_t^{(n)}$. The short-term risk-free rate r_t^f is the yield on a one-period bond.

The local-currency bond excess return on the domestic zero-coupon bond in local currency $rx_{t+1}^{(n)}$ is defined as:

$$rx_{t+1}^{(n)} = -(n-1)y_{t+1}^{(n-1)} + ny_t^{(n)} - r_t^f.$$

It represents the excess return on the "yield-curve trade"—the trade that borrows short-term and lends long-term in domestic currency. It is the return from holding a long-term bond for one period, financed by borrowing at the short-term rate in domestic currency.

As in Section 2.3, Q_t is the nominal spot exchange rate in terms of domestic currency per U.S. dollar, and $\Delta q_{t+1} = \log(Q_{t+1}/Q_t)$ is the rate of domestic currency depreciation.

The excess return on home currency, rx^q_{t+1}, is

$$rx^q_{t+1} = r^f_t - r^{f,US}_t - \Delta q_{t+1}.$$

It represents the return on the "UIP trade"—the trade that borrows short-term in foreign currency and lends short-term in domestic currency over one period.

Data We select all economies with bond benchmark indexes on *Datastream*. We use monthly data from January 2006 to December 2021.[19] The dataset includes spot exchange rates, 10-year government bond total return indexes, and 3-month deposit rates.[20] Our sample includes Australia, Canada, Colombia, India, Indonesia, Japan, Mexico, New Zealand, Poland, South Africa, South Korea, Sweden, Switzerland, Thailand, and the United Kingdom.

Analysis Consider the following country-j-specific regression of excess currency returns on excess bond returns:

$$rx^q_{j,t+3} = \alpha + \beta rx^{(10y)}_{j,t+3} + \varepsilon_{j,t+3}. \tag{2.25}$$

Table 2.4 (Columns (1)-(2)) reports the estimated β of equation (2.25) and the associated correlation coefficients for our panel of countries. There is vast heterogeneity across

[19] While some countries' sample starts in December 1999, the data is sparse before 2006, and thus we start our analysis in 2006. We exclude data beyond December 2021 to avoid incorporating recent inflation dynamics.

[20] Long maturity bond return data from *Datastream* may pertain to coupon government bonds.

Table 2.4: Bond and currency excess returns (3-month excess returns)

			Residualized excess returns	
Country	Regression coefficient	Correlation	Regression coefficient	Correlation
Australia	-0.88^{***}	-0.47	-0.68^{***}	-0.41
	(0.17)	$[-0.57, -0.35]$	(0.18)	$[-0.54, -0.26]$
Canada	-0.69^{***}	-0.44	-0.74^{***}	-0.51
	(0.11)	$[-0.55, -0.32]$	(0.13)	$[-0.62, -0.38]$
Colombia	0.13^{**}	0.14	0.18^{***}	0.12
	(0.06)	[0.00, 0.28]	(0.07)	$[-0.04, 0.28]$
India	0.02	0.02	0.07	-0.01
	(0.13)	$[-0.13, 0.17]$	(0.12)	$[-0.18, 0.16]$
Indonesia	0.56^{***}	0.62	0.57^{***}	0.21
	(0.08)	[0.49, 0.72]	(0.10)	[0.02, 0.39]
Japan	1.37^{***}	0.35	1.12^{***}	0.29
	(0.30)	[0.21, 0.47]	(0.38)	[0.13, 0.43]
Korea	-0.06	-0.04	-0.15	-0.02
	(0.16)	$[-0.22, 0.14]$	(0.18)	$[-0.21, 0.18]$
Mexico	0.29^{**}	0.20	0.36^{***}	0.17
	(0.11)	[0.03, 0.35]	(0.12)	$[-0.02, 0.34]$
New Zealand	-0.61^{***}	-0.34	-0.45^{**}	-0.27
	(0.16)	$[-0.46, -0.21]$	(0.18)	$[-0.42, -0.11]$
Poland	0.09	0.04	-0.01	0.11
	(0.14)	$[-0.10, 0.18]$	(0.13)	$[-0.06, 0.27]$
South Africa	0.61^{***}	0.30	0.74^{***}	0.24
	(0.16)	[0.16, 0.42]	(0.20)	[0.07, 0.39]
Sweden	-1.01^{***}	-0.53	-0.98^{***}	-0.55
	(0.14)	$[-0.63, -0.42]$	(0.14)	$[-0.66, -0.43]$
Switzerland	-0.31^{*}	-0.16	-0.23	-0.12
	(0.17)	$[-0.30, -0.02]$	(0.18)	$[-0.28, 0.04]$
Thailand	0.51^{**}	0.31	0.51^{**}	0.20
	(0.21)	[0.00, 0.56]	(0.21)	$[-0.11, 0.47]$
United Kingdom	-0.69^{***}	-0.45	-0.74^{***}	-0.51
	(0.12)	$[-0.56, -0.33]$	(0.15)	$[-0.63, -0.38]$

Note: Column (1) reports the estimated slope coefficient of regression equation (2.25). Column (2) reports the correlation between $rx^{q}_{j,t+3}$ and $rx^{(10y)}_{j,t+3}$. Columns (3)-(4) repeat the analysis using the excess returns residualized by current and lag probability of default in the 10-year sovereign bonds (as retrieved from *Refinitiv*), that is: $rx^{m}_{j,t+3} = \sum_{k=0}^{1} \phi_k(\text{Prob. default}_{j,t-k}) + \varepsilon^{rx^m}_{j,t+3}$, where $m = \{10y, q\}$. Stars denote significance: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. Values in parenthesis are standard errors, and values in square brackets are 95 percent confidence interval.

countries in the comovement between bond and currency returns. Both regression and correlation coefficients are either significantly negative, significantly positive, or statistically insignificant. These conclusions apply also when using 1-month excess returns, as documented in Appendix Table B.2.

We provide evidence that rules out time-varying default risk as an explanation for the (positive) association between bond and currency excess returns. We proxy for default risk using probabilities of default on 10-year sovereign bonds from *Refinitiv*. We residualize

both bond and currency excess returns with respect to contemporaneous and lagged default probabilities, and then regress residualized excess currency returns ($\varepsilon^{rx^q}_{j,t+3}$) on residualized excess bond returns ($\varepsilon^{rx^y}_{j,t+3}$). Table 2.4 (Columns (3)–(4)) shows that the residualized returns exhibit comovement patterns very similar to those observed in the raw data (Columns (1)–(2)).

2.6.2 Foreign investors and currency hedging across countries

Our mechanism links the comovement of bond and currency returns to a country's exposure to unhedged foreign investor flows, and thus to the extent to which intermediaries' positions are positively correlated across bond and currency markets. This degree of correlated positions is governed by Ω in equation (2.19). We now turn to estimating Ω for each country in our sample.

We use the standard deviations of the shares of foreign and non-bank domestic investors, respectively, in local-currency government debt securities to proxy for the volatility of foreign and domestic investors' flows, $\sigma_{\eta^\star}$ and σ_η, in equation (2.19). These measures are based on data on quarterly investor holding shares of sovereign debt for the sample period 2006-2021 (Arslanalp and Tsuda, 2012, 2014). To do so, we assume that the total outstanding amount of local-currency bonds is constant, so as to use investor holding shares, rather than levels, to construct estimates of $\sigma_{\eta^\star}$ and σ_η.[21] We identify domestic banks as the relevant "intermediaries," a choice consistent with the analysis and evidence in Section 2.4.

To measure hedging ratios, α, we rely on estimates from the existing literature. For advanced economies, we use hedge ratios for global fixed-income investors from Cheema-

[21] Non-bank domestic investors exclude domestic bank and central bank holdings. For emerging economies, we construct these shares using the share of non-bank domestic investors in total sovereign debt, as data disaggregated by currency denomination are not available.

Table 2.5: Measure of correlated positions across countries

Country	Corr. Exp. Ω	Global Inv. (s.d., %)	Local Inv. (s.d., %)	Hedge Ratio α
Australia	0.09	6.24	9.45	0.75
Canada	0.08	3.39	5.32	0.76
Colombia	0.46	8.29	8.72	0.05
India	0.09	1.16	4.00	−0.20
Indonesia	0.81	9.27	4.63	−0.09
Japan	0.03	2.70	9.87	0.56
Korea	0.14	4.07	11.56	−0.33
Mexico	0.70	12.21	8.88	−0.24
New Zealand	0.20	8.59	8.03	0.74
Poland	0.59	8.44	8.14	−0.38
South Africa	0.70	10.62	6.70	0.09
Sweden	0.04	7.94	9.99	0.91
Switzerland	0.12	7.54	9.47	0.75
Thailand	0.63	4.52	3.97	−0.36
UK	0.001	2.10	20.41	0.93

Note: Column 1 computes our estimated measure of overall correlation in net supplies, based on eq. (2.19). The table also contains the standard deviation in the share of local-currency sovereign debt securities by foreigners (Column 2) and domestic non-bank investors (Column 3) using data from Arslanalp and Tsuda (2012, 2014). Column 4 computes the hedging ratio for each country, following Cheema-Fox and Greenwood (2024) for advanced economies, and Chen and Zhou (2025) for emerging economies.

Fox and Greenwood (2024). For emerging economies, we use the estimates of hedge ratios for US mutual funds documented in Chen and Zhou (2025). While the evidence on hedging ratios for emerging economies is based on US mutual funds, we note that international mutual funds are the main foreign investor type in Colombia's local currency government bond market (see Appendix B.2.1).[22]

Table 2.5 reports, for each country, the estimated importance of unhedged foreign investors' bond flows, labeled Ω, along with its components $\sigma_{\eta^\star}$, σ_η, and α. We observe substantial cross-country heterogeneity in Ω, which ranges from close to zero in the United Kingdom to roughly 80 percent in Indonesia. These differences reflect variation across countries in both the relative importance of foreign versus domestic investor flows and in foreign investors' currency-hedging behavior.

[22]The hedging ratio estimates correspond to averages over 1998-2023 from Cheema-Fox and Greenwood (2024) and 2010–2023 from Chen and Zhou (2025).

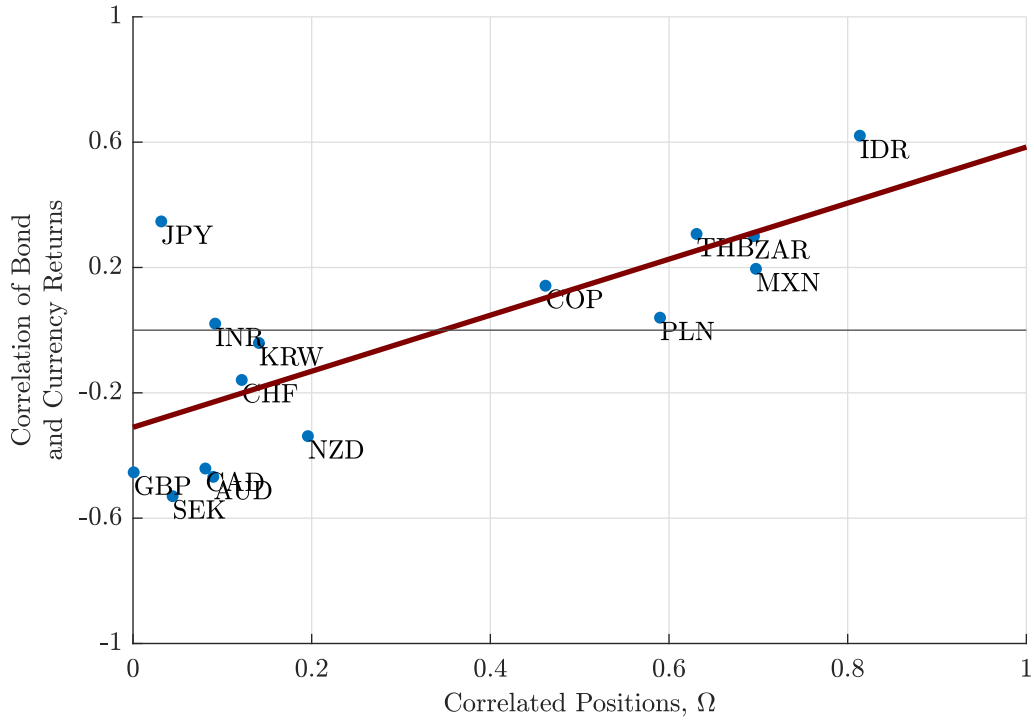
2.6.3 Correlated positions and correlated bond–currency returns

We document that, in line with our model's mechanism, countries with a higher estimated correlation in intermediaries' positions exhibit, on average, a more positive comovement between bond and currency returns, as illustrated in Figure 2.5. Although our empirical measure of correlated positions, Ω, is admittedly simple, we view this evidence as indicative of a meaningful relationship between the relative importance of unhedged foreign flows in total flow volatility and the comovement of bond and currency returns. We note that both the differential presence of foreign investors' and the differential propensity to hedge currency risk independently contribute to cross-country variation in bond–currency return comovement (see Appendix Figure B.3).

It is worth noting that hedge ratios of U.S. mutual funds investing in emerging economies are sometimes negative. This reflects the tendency of U.S. mutual funds to further increase, rather than reduce, their exposure to emerging market currencies, and it is consistent with the models in De Leo et al. (2024c) and Chen and Zhou (2025), where forward demand partly reflects currency speculation. Our model can easily accommodate "negative hedging ratios," as α need not be restricted between 0 and 1. That said, in Appendix Figure B.2, we recompute Ω after truncating the distribution of hedge ratios so that $\alpha \geq 0$, and the substantive conclusions of this analysis remain unchanged.

Advanced vs. emerging economies Figure 2.5 highlights a distinction between advanced and emerging economies. With a few notable exceptions, emerging economies exhibit both high correlated positions and a positive comovement between bond and currency returns. As shown in Table 2.5, the large correlated positions observed in emerging economies can be attributed to the prominent role of foreign investors *and* their limited use of currency hedging.

Figure 2.5: Correlated Positions and Bond-Yield Comovement



Note: This figure displays the correlation of bond and currency excess returns (y-axis) and the Ω from eq. (2.19) (x-axis).

Systematic differences in hedge ratios between emerging and advanced economies are not surprising. Foreign investors are typically attracted to emerging markets by substantially higher interest rate differentials relative to advanced economies. Hedging currency risk in such environments is costly, as it effectively requires giving up the interest rate differential (assuming covered interest parity approximately holds).[23] In addition, many emerging economies—both sovereigns and corporates—issue debt in U.S. dollars, providing an alternative to investing in local-currency assets that require costly currency hedging, which may further contribute to the observed differences in hedging behavior across the two groups of countries.[24]

[23] This consideration has led many investment advisors, such as Meketa Investment Group (2022), to recommend avoiding currency hedging for emerging market investments, while viewing hedging as more appropriate for advanced economies.

[24] De Leo et al. (2024c) argue that foreign investors oriented toward emerging economies primarily use the forward market for currency speculation rather than for hedging purposes.

2.7 Conclusion

Local-currency government bond markets worldwide are populated by investors with different funding currencies. We show that investor composition shapes the distribution of flows that financial intermediaries must absorb—and, consequently, the premia they require—across both bond and currency markets. This simple mechanism helps explain the effects of foreign capital inflows into local-currency bond markets, the patterns of positions and returns across different investors, the transmission of central bank asset purchase policies, and cross-country differences in the joint dynamics of bond yields and exchange rates.[25]

[25] A natural extension would allow for heterogeneity in investor characteristics such as risk tolerance, portfolio elasticity, and investment horizons. Sovereign bonds are held by a diverse set of investors across countries (Fang et al., 2022; Zhou, 2024), and preferences over currency denomination may further segment investor demand (Faia et al., 2024).

Chapter 3: Optimal Seigniorage: With and Without Foreign Exchange Controls

3.1 Introduction

Paradigms in matters related to international policy often change in the wake of financial crises and their accompanying recessions. One clear example of such a reversal concerns the free flow of capital across borders. Prudential capital and exchange controls can be tools to combat excessive volatility in external flows and to bolster monetary autonomy. The word 'prudential', however, is not always suitable for a particular set of developing countries and their usage of such restrictions.

Most countries have seemingly overcome chronically high inflation. A common pattern arises when we look at those who still have not: persistent fiscal deficits.[1] What is problematic is the inability—and often unwillingness—to close the fiscal gap. Revenue collection via money printing (*seigniorage*) is presented as the unavoidable consequence amid weak tax systems, shallow financial markets, and little-to-no access to international capital (Keynes, 1924; Click, 1998).

This problem does not go unnoticed by agents, evidently. Persistently high inflation leaves scars. Once individuals lose confidence in their own money, currency substitution (CS) emerges (Giovannini and Turtelboom, 1992). The reduction in the real value of the

[1] See, for example, Sargent (1986), Kiguel and Liviatan (1988), Dornbusch and Edwards (1990), Calvo and Végh (1992), Miller et al. (2019).

stock of base money due to inflation (the *inflation tax*) creates incentives to substitute away from the domestic currency and into more stable counterparts, like the dollar. Therefore, excessive reliance on seigniorage not only exacerbates inflationary pressures, by widening the gap between money supply and money demand (Rojas-Suarez, 1992), but also has high welfare costs for society as inflation acts as a tax on money holdings (Bailey, 1956).

In this paper I describe a typical situation faced by governments in developing economies: an inherent tilt towards inflationary finance. The use of conventional taxes to collect revenue is often neither effective nor politically desirable.[2] On top of this, external borrowing is costly and often limited. To make matters worse, reliance on the inflation tax causes agents to substitute away from the ever-weakening currency, hindering the authority's capacity to collect revenue from seigniorage. The government may then consider whether it can ever be desirable to impose restrictions on the use of foreign assets, and if so, when. It has long been argued that restrictions on the use of foreign monies (*exchange controls*) could allow governments to rely more heavily on the inflation tax (Nichols, 1974). This work studies how seigniorage can be optimally deployed, both with and without exchange controls.

I develop an infinite-horizon optimal taxation framework with two key ingredients. First, the government has access to conventional taxation, subject to increasing marginal costs of tax collection. Second, the setting is an open economy that allows for foreign currencies to serve as a store of value and medium of exchange. These two components, when put together, yield a novel result: for given (positive) revenue needs, the optimal inflation tax is always positive, irrespective of the foreign nominal interest rate. Here, convex tax-collection costs play a key role: they break the optimality of a zero domestic nominal interest rate when the foreign rate is zero.

[2] It is well documented that tax evasion and the informality is pervasive in developing countries. See, for example, Gerxhani (2004), López (2017).

I solve the model with and without exchange controls. I show that, because of currency substitution, the model without foreign exchange restrictions always calls for a higher optimal inflation tax. Additionally, the domestic nominal interest rate is increasing in the foreign nominal rate and in the tax-collection capacity of authorities. This last point yields policy implications. First, the decision to impose exchange controls is independent of the effectiveness of the tax system. That is, a weak tax-policing structure is broadly so, irrespective of the foreign exchange restrictions in place. Second, it can be desirable to limit the use of other currencies whenever the foreign nominal interest rate is high. Doing so greatly reduces inflation. Lastly, and probably the clearest implication, is that authorities should invest in actively reducing the inefficiencies of tax collection.

Related Literature Bailey (1956) and Cagan (1956) who pioneered the theoretical literature on seigniorage. Cagan highlighted monetary growth as the cause of very high inflation episodes, while Bailey initiated the consideration of inflation as a tax, switched the emphasis from monetary to fiscal factors as the main cause of very high inflation episodes. Phelps (1973) inspired the modern literature on the optimal collection of seigniorage from a public finance point of view. Phelps shows that, by integrating the revenues from money creation into a public finance framework, an optimal combination of all possible revenue sources is called for—including the inflation tax. This sparked an extensive debate regarding the optimality of the Friedman (1969) rule, which advocates for monetary policy that produces a nominal interest rate of zero—which implies an inflation tax equal to minus the real interest rate. This rule relies on the simple idea that the social cost of creating fiat money should equal the opportunity cost of holding it. Since the marginal cost of creating additional bills is approximately zero, the nominal interest rate should be zero. This result, however, is derived in a first-best environment where lump-sum taxes are available. The literature has, thus, typically rejected the Friedman rule, especially for low and middle income countries

where tax systems are weaker and conventional taxes less effective (Aizenman, 1987; Fuest and Riedel, 2009).

(Nichols, 1974) addressed currency substitution and capital exchange controls, arguing that the ready availability of stable alternative currencies makes the demand for real domestic money more elastic, resulting in a higher inflation rate for a given deficit.[3] Even in the presence of currency substitution, the literature remains divided between proponents of an optimal positive inflation tax (Végh, 1989a; Guidotti and Végh, 1993), and those advocating for an optimal inflation tax of zero (Kimbrough, 1991). These differing conclusions result from the different ways of introducing money into the model and the assumptions made.

This paper builds a theoretical framework that relies on two key ingredients: (i) two currencies are introduced and are required to conduct transactions; and (ii) the government has access to conventional taxes that carry collection costs. Previous work kas introduced money through a liquidity-in-advance constraint (Holman and Neanidis, 2006), with a monies-in-the-utility specification (Bufman and Leiderman, 1993), or using a shopping-time or transactions costs model (Végh, 1989a; Kimbrough, 1991). Meanwhile, several studies consider tax-collection costs in a closed-economy inflationary finance setup (Végh, 1989b; Canzoneri and Diba, 1992; De Fiore, 2000; Koreshkova, 2006). The motivation for costly tax collection is the strong prevalence of the underground economy in poorer economies, as shown in Schneider and Enste (2000), which enables evasion of conventional taxation.

The modeling novelty of this study is that currency substitution and convex tax-collection costs are modeled jointly alongside foreign exchange controls in a public finance setting. A closely related work is Hercowitz and Sadka (1987). They question whether it is socially desirable to impose foreign exchange controls to tax the national currency. In their model,

[3]For an early discussion of inflationary finance under currency substitution see Calvo and Végh (1992).

in which conventional taxation carries no collection costs (which is more suitable for developed economies), it is never desirable to ban the use of foreign monies. Additionally, inflationary finance is also never optimal. The model showcased here departs from both results. Using a shopping-time specification for both monies and incorporating increasing marginal collection costs to conventional taxation, I demonstrate that the optimal inflation tax is (always) positive, both with and without exchange controls. Additionally, I show here that a high foreign nominal interest rate could call for the presence of restrictions on the use of foreign currency.

Two existing studies incorporate tax-collection costs in an open economy: Aizenman (1987) and Holman and Neanidis (2006). Aizenman (1987) analyzes how the effectiveness of tax collection affects the optimal inflation tax. However, that work does not allow currency substitution, and assumes that the inflation tax carries very little deadweight losses. Meanwhile, Holman and Neanidis (2006) assume that both currencies serve as medium of exchange and the agent can evade a constant (exogenous) fraction of what the government levies as income taxes. They derive welfare implications for both sources of taxation and find that seigniorage financing has stronger negative implications. My methodological combination allows me to obtain new results on the complementarity and optimality of the different taxes.

Outline The paper is structured as follows. Section 3.2 presents several facts about countries that rely heavily on seigniorage and how they compare to those that do not. Section 3.3 introduces an optimal taxation framework in an open economy with conventional taxes with increasing marginal collection costs, demonstrating the influence of exchange controls on determining the optimal inflation tax. Finally, Section 3.4 gives closing remarks.

3.2 Anatomy of Seigniorage-Reliant Countries

I show here four facts that distinguish countries that use the inflation tax as a revenue device. I only include emerging and developing economies, as these countries are the main focus of this work.[4] These facts are:

1. Fiscal deficits are positively correlated with seigniorage revenue.

2. There exists a positive correlation between inflation and seigniorage revenue.

3. Deposit dollarization is negatively correlated with seigniorage revenue.

4. Countries generally collect some revenues from seigniorage, though it is more prevalent in countries that are more financially closed.

Using annual data from 1980 for over 70 countries, Figure 3.1 shows that the average fiscal deficit is twice as large for countries and periods in which seigniorage revenue is substantial (highest quintile), compared to when it is lower (lowest quintile). Monetary financing is thus more prevalent when there are large fiscal needs. Instead of extinguishing the fiscal deficit fires, money printing merely adds more paper to the flames.

Cagan (1956) showed theoretically that sustained relentless monetary growth leads to high inflation episodes. Figure 3.2 shows a positive correlation between inflation and seigniorage revenue, possibly implying that seigniorage-reliant economies experience higher inflation rates. This observation should come as no surprise: relentless money printing to finance expenditures is a self-inflicted malady. While many countries have overcome such unlawful practices by addressing their fiscal balances, others still use the inflation tax as a policy instrument. Finally, the fact that inflation is higher in countries with larger seigniorage revenue need not be solely because governments print more money. It can also reflect

[4] A detailed description of the data and the countries included can be found in Appendix C.1.

a higher elasticity of demand for domestic balances, or ineffectiveness of if exchange controls.

The inflationary consequences of a given budget deficit are aggravated as the public switches from domestic to foreign monies or other foreign assets. This becomes more important as the inflation rate accelerates, because the public tends to adjust their portfolios faster (Calvo and Végh, 1992). In other words, currency substitution exacerbates high inflation because it represents a loss of monetary autonomy for the authorities. While a central bank can keep track of its domestic money supply relatively easily, that is not the case for foreign currency. Figure 3.3 exhibits a negative relationship between the share of foreign currency deposits in total deposits and seigniorage revenue. Intuitively, when a surge in national money supply is not met with an equal increase in national money demand, real seigniorage revenue collection is hindered. A caveat of this chart is that data on currency substitution is not readily available for most countries, so this figure must rely on estimates using foreign currency deposits.

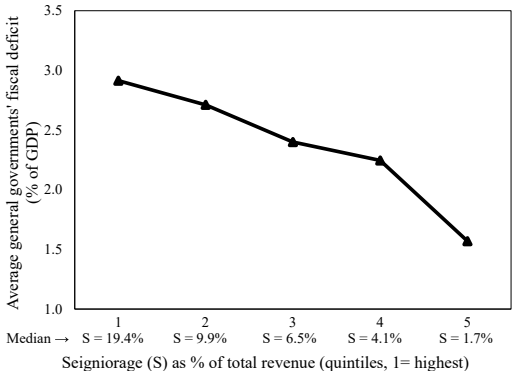


Figure 3.1: Fiscal deficit

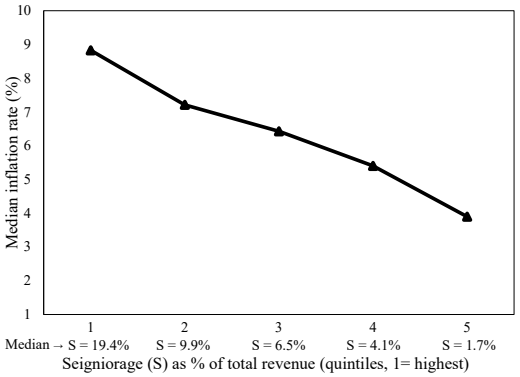


Figure 3.2: Inflation

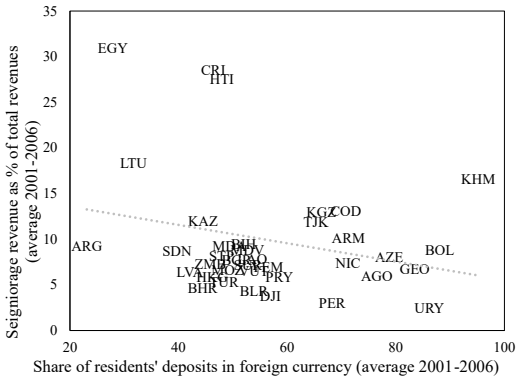


Figure 3.3: Deposit dollarization

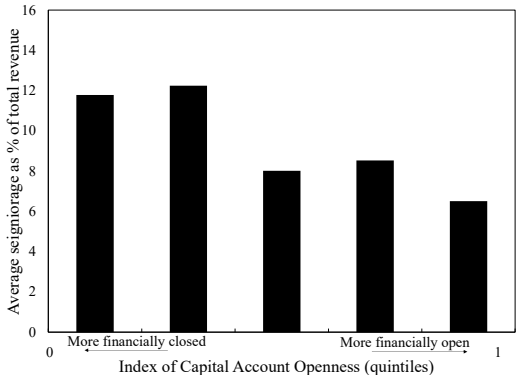


Figure 3.4: Capital account openness

Annual data since 1980, when available, is used for over 70 emerging and developing countries. Seigniorage is calculated as the change in the monetary base as a share of government's total revenues. The values below the x-axis in Figures 3.1 and 3.2 represent the median seigniorage value for each quintile. In Figure 3.1 (Figure 3.2), every country-year observation is organized in quintiles and each dot represent the average (median) value of the fiscal deficit (inflation) for each quintile. Figure 3.3 computes residents' average share of deposits denominated in foreign currency for the period 2001-06 (Honohan, 2008). Figure 3.4 allocates every country-year observation with respect to its capital account openness index (Chinn and Ito, 2006) into quintiles and computes the quintiles' average seigniorage. Figures use data from the IMF, The World Bank, and national sources.

For various reasons not analyzed here, governments might look into limiting the free

flow of foreign monies and/or assets.[5] This enables more control over who pays the inflation tax, and greater revenue collection. Figure 3.4 documents, using Chinn and Ito's (2006) index of capital account openness, that countries with relatively higher seigniorage revenue also have tighter capital and exchange controls in place. In other words, the imposition of restrictions on the use of foreign assets makes seigniorage more relevant and less wasteful—in the sense that a lower inflation tax is required for given revenue needs.

By allowing foreign assets to move freely in and out of the economy, an open financial account can facilitate economic integration. However, in contexts where currency substitution is widespread, this openness also restricts the government's capacity to resort to the inflation tax by diminishing control over the monetary base (Brock, 1984). On the other hand, by having greater control over the tax base through tight capital and foreign exchange restrictions (Alesina et al., 1993), enables governments to rely more heavily on monetary financing, all the while having a smaller effect on inflation than would otherwise prevail in a context of currency substitution. Therefore, authorities can use controls to salvage seigniorage revenue.

In summary, large fiscal deficits are positively correlated with seigniorage revenues, which are in turn positively correlated with inflation. Currency substitution negatively affects the scope for revenue collection through seigniorage by reducing the tax base, and capital and exchange controls can be deployed to counteract the adoption of alternative currencies. Not surprisingly, and whether by choice or inevitably, those countries with tighter capital and foreign exchange restrictions seem to rely more on seigniorage as a revenue device.

[5] It is worth noting that, while some countries voluntarily opt out of foreign exchange and international capital markets, others have no choice but to restrict their capital account due to, for example, sovereign defaults or IMF programs, among other reasons. For two good reviews of capital and exchange controls see Magud et al. (2011) and Rebucci and Ma (2019).

3.3 Optimal Inflation Tax in a Public Finance Framework

There exist multiple models in which the optimal inflation tax is different from zero. To begin with, standard monetary models assume that money either (a) reduces transactions costs/time, or (b) offers liquidity services. Examples can be found in Chari and Kehoe (1999). The model developed here combines two mechanisms that, under certain assumptions, can be shown to break the optimality of the Friedman rule. Doing so obtains a novel open-economy result: the optimal inflation tax is different from zero for *any* positive revenue needs. The two mechanisms are multiple currencies, and imperfect taxation. On the one hand, Végh (1989a) and Guidotti and Végh (1993) demonstrate that, in open economies where domestic and foreign monies coexist, the optimal inflation tax is positive whenever the foreign nominal rate is non-zero. On the other hand, assuming that alternative revenue sources reduce welfare (imperfect taxation) yields a positive optimal inflation tax. For instance De Fiore (2000) assumes tax collection costs in a one-currency shopping-time model. Furthermore, Koreshkova's (2006) work emphasizes the existence of a large tax-avoiding informal sector.

3.3.1 Setup

I present here a one-good small open economy model with flexible exchange rates.[6] The only factor of production is labor, n_t, under a CRS technology given by $y_t = n_t$. Producing one unit of the good requires one unit of labor, and the wage rate is normalized to 1. A representative household is endowed with one unit of time in every period and can allocate it among leisure (h_t), labor (n_t), and shopping time (s_t), such that $h_t + n_t + s_t = 1$ for every period t.

There are three assets in this economy: an internationally traded bond, B^* with exoge-

[6]The setup here is similar to Végh (1995), with convex tax-collection costs as in De Fiore (2000).

nous real return r, and domestic (m) and foreign (f) money. Currency facilitates carrying out transactions, in that it saves time devoted to shopping activities.[7] In fact, both monies serve as media of exchange:

$$s_t = V\left[\frac{m_t}{c_t}, \frac{f_t}{c_t}\right] c_t, \tag{3.1}$$

where

$$V \geq 0, \quad V_1, V_2 \leq 0, \quad V_{11}, V_{22} > 0, \quad V_{12} > 0. \tag{3.2}$$

where c_t denotes consumption of the good in period t. $V_{12} > 0$ ensures that currencies are substitutes. Allowing the consumer to choose the media of exchange based on the currencies' underlying costs is a desired feature of this study. Denote $X^m = m/c$ and $X^f = d/c$ as domestic and foreign relative money balances, respectively. Increased relative money holdings reduce shopping time at a diminishing rate. I assume that there exists a satiation point $(X^{m,max}, X^{f,max})$ at which no more gains in terms of reduced shopping time can be obtained. That is, $V_1(X^{m,max}, X^{f,max}) = 0$ and $V_2(X^{m,max}, X^{f,max}) = 0$. Without loss of generality, I assume that, at this locus of points, it also holds that $V(X^{m,max}, X^{f,max}) = 0$. In other words, shopping requires no time from the household. I show in the upcoming discussion that this satiation property allows me to establish an equivalence between the two model economies that I analyze and compare.

Household Maximization Problem The household chooses $\{c_t, h_t, m_t, f_t\}_{t=0}^{\infty}$ to maximize a twice-continuously differentiable, increasing strictly concave utility function:

$$\sum_{t=0}^{\infty} \beta^t U(c_t, h_t), \tag{3.3}$$

[7] Work by Correia and Teles (1999) has shown that this transactions function is isomorphic to a money-in-the-utility function under certain assumptions. Additionally, McCallum (1983) shows that a cash-in-advance constraint is a particular case of this transactions technology.

subject to

$$P_t c_t + M_t + \mathscr{E}_t F_t + \phi_t \mathscr{E}_t F_t + \mathscr{E}_t B^*_{t-1}(1+i^*) = (1-\tau_t)P_t(1-h_t-s_t) + \mathscr{E}_t B^*_t + M_{t-1} + \mathscr{E}_t F_{t-1}; \tag{3.4a}$$

$$B^*_{-1} = M_{-1} = F_{-1} = 0; \tag{3.4b}$$

$$\lim_{t\to\infty} \frac{B^*_{t+1}}{P_{t+1}} = 0, \tag{3.4c}$$

where s_t is modeled as in (3.1) and (3.2). P_t denotes the price of the consumption good in period t, $\mathscr{E}_t$ is the nominal exchange rate, and i^* is the exogenous nominal foreign interest rate. τ_t is the labor income tax, which is levied on n_t; and lastly ϕ_t is a tax on the holdings of foreign currency. I assume that exchange controls are in full action if $\phi \to \infty$, which drives $F \to 0$. While it could be argued that a total ban on foreign money is implausible, evasion of FX controls is beyond the scope of this paper and left for future work. Using (3.4b) and (3.4c), I can iterate forward (3.4a) (in real terms) to formulate the lifetime budget constraint:

$$\sum_{t=0}^{\infty} \left[\frac{1}{1+r}\right]^t (1-\tau_t)(1-h_t-s_t) = \sum_{t=0}^{\infty} \left[\frac{1}{1+r}\right]^t [c_t + I_t m_t + (I^* + \phi_t) f_t], \tag{3.5}$$

where $I_t = \frac{i_t}{1+i_t}$ and $I^* = \frac{i^*}{1+i^*}$ are the domestic and foreign inflation tax, respectively. The term i_t is the domestic nominal interest rate, which satisfies

$$1 + i_t = (1+i^*)\frac{\mathscr{E}_{t+1}}{\mathscr{E}_t}, \tag{3.6}$$

while the real foreign interest rate is derived as

$$1 + r = \frac{1+i^*}{1+\pi^*}. \tag{3.7}$$

The last two terms in the right-hand side of equation (3.5) represent *after-inflation-tax* money holdings. Intuitively, these terms represent the opportunity cost of holding m units of real money. For instance, by holding M_t units of money in t, the consumer forgoes $i_t M_t$ in interest that he could have accrued in $t+1$. The real value of this expression at time t is

$$\frac{1}{1+r}\left(\frac{i_t M_t}{P_t(1+\pi_t)}\right) = \frac{i_t}{1+i_t} m_t.$$

Expenditures thus consist of the single consumption good and the real costs of holding the two currencies.

The household therefore maximizes (3.3) subject to (3.5) and (3.1). Note that this problem can be solved in two stages, observing that real money holdings do not enter the utility function. Thus, in a first stage, the household chooses $\{s_t, m_t, f_t\}$. First-order derivatives yield

$$m_t: \qquad -V_1(X_t^m, X_t^f) = I_t(1+\tilde{\tau}_t) \qquad t = 0,1,2,...; \tag{3.8a}$$

$$f_t: \qquad -V_2(X_t^m, X_t^f) = (I^* + \phi_t)(1+\tilde{\tau}_t) \qquad t = 0,1,2,..., \tag{3.8b}$$

where I have used that $(1+\tilde{\tau}) = \frac{1}{1-\tau}$ for simpler exposition. Equations in (3.8) equate the gains in terms of reduced shopping time of holding an extra unit of (domestic or foreign) currency to their opportunity cost.

As can readily be seen in equations (3.5) and (3.8), the labor income tax affects not only time devoted to working but also the prices of consuming and holding currency. Denote $R = I(1+\tilde{\tau})$ and $R^* = (I^* + \phi)(1+\tilde{\tau})$ as the *effective* cost of holding domestic and foreign relative money balances, respectively. The strict convexity of $V(X^m, X^f)$ allows us to fully

describe the money demand functions:

$$X_t^m = X^m(R_t, R_t^*) \qquad t = 0,1,2,... \qquad X_R^m < 0, \;\; X_{R^*}^m > 0; \tag{3.9a}$$

$$X_t^f = X^f(R_t, R_t^*) \qquad t = 0,1,2,... \qquad X_R^f > 0, \;\; X_{R^*}^f < 0. \tag{3.9b}$$

Intuitively, an increase in the domestic inflation tax, I, increases the effective cost of holding domestic relative money balances, R, inducing households to substitute towards foreign currency. Hence, this model allows foreign currency to fulfil two common roles of money: store of value and medium of exchange.[8]

I can now solve for the second stage of this maximization problem by plugging (3.9) into (3.5), obtaining:

$$\sum_{t=0}^{\infty} \left[\frac{1}{1+r}\right]^t \left[\frac{(1-h_t)}{1+\tilde{\tau}_t}\right] = \sum_{t=0}^{\infty} \left[\frac{1}{1+r}\right]^t \left[\frac{q_t c_t}{1+\tilde{\tau}}\right], \tag{3.10}$$

where $q_t = q(\tilde{\tau}, R_t, R_t^*)$ is the distortion generated by distortionary taxes, both domestic and foreign:

$$q_t = 1 + \tilde{\tau} + V\left[X^m(R_t, R_t^*), X^f(R_t, R_t^*)\right] + R_t X^m(R_t, R_t^*) + R_t^* X^f(R_t, R_t^*). \tag{3.11}$$

Households then choose $\{c_t, h_t\}_{t=0}^{\infty}$ to maximize (3.3) subject to (3.10), using (3.11). Assuming that $\beta = 1/(1+r)$, it obtains that

$$\frac{U_c(c_t, h_t)}{U_h(c_t, h_t)} = q_t \qquad t = 0,1,2,... \tag{3.12}$$

This optimality condition allows us to compute optimal consumption and leisure choices

[8] Another noteworthy difference between this work and Hercowitz and Sadka's (1987) is that foreign money serves only as a store of value in their model.

as a function of their relative price, $c_t = c(q)$ and $h_t = h(q)$ for all t. It follows that, given the nature of the distortion, $c' < 0$ and $h' < 0$. Consequently, consumer's welfare decreases as q rises.

Government's Problem The government faces an exogenous path of spending, assumed to be constant over time ($g_t = g$ for all t). Lump-sum taxation is not feasible, which is at the core of this problem. The government must finance its expenditures through distortionary taxes. The three instruments the government has access to are a labor income tax (τ), the inflation tax (I), and the tax on foreign money balances (ϕ). The budget constraint reads[9]

$$g = \tau y + Im + \phi f - \kappa(\tau y), \tag{3.13}$$

where $\kappa(\cdot)$ denotes the tax-collection costs, which satisfy $\kappa'(\cdot) > 0$ and $\kappa''(\cdot) \geq 0$. Convex costs imply that the higher tax rate is, the harder it is for the government to police tax collection. Note that a model with tax-collection costs is isomorphic to one where there are two sectors: (i) a formal tax-paying sector, and (ii) and an informal tax-evading one, where higher taxes induce substitution towards the informal sector. I assume away problems of time inconsistency in what follows.

3.3.2 The Optimal Taxation Problem

I now look at two limiting cases: (i) fully enforceable exchange controls ($\phi \to \infty$); and (ii) the free flow of foreign monies ($\phi \to 0$). The former totally bans the use of the foreign currency, effectively reducing this problem to a closed economy. It is worth mentioning that this work is not particularly interested in analyzing cases where $\phi \in (0, \infty)$, but some

[9]Note that the nature of the problem makes it so that the optimal taxation structure will be constant over time, so I can drop time subscripts. The economy is always at the steady state and will adjust immediately to any changes.

considerations about the desirability of exchange controls will be addressed towards the end of this section. Endogenizing the exchange controls parameter is left for future work. Furthermore, an outright ban on the use of foreign currencies could generate incentives for a parallel market for foreign exchange, but that is also left for future work.

For now, it serves to rewrite the government's budget constraint in terms of $\tilde{\tau}$, R, R^*, and ϕ by using $y = 1 - h - s$ and equation (3.5):

$$g = c\left[q(\tilde{\tau}, R, R^*; I^*)\right] \Omega(\tilde{\tau}, R, R^*; I^*), \tag{3.14}$$

where

$$\Omega(\tilde{\tau}, R, R^*; I^*) = \tilde{\tau} + RX^m + (R^* - I^*)X^f - \frac{\kappa(\tilde{\tau} y)}{c}. \tag{3.15}$$

Intuitively, equation (3.14) describes the government's budget balance in terms of consumption units.

Instead of using a dual approach to the optimal Ramsey tax problem, which would entail choosing $(\tilde{\tau}, R)$ to maximize the indirect utility function subject to the government's budget constraint, I can make use of the fact that welfare is decreasing in the distortion, $q(\tilde{\tau}, R, R^*; I^*)$. The problem can then be recast into one where the distortion (equation (3.11)) is minimized subject to the government's budget constraint (equation (3.14)). The following optimality condition is obtained:

$$\frac{\dfrac{\partial q}{\partial R}}{\dfrac{\partial q}{\partial \tilde{\tau}}} = \frac{\dfrac{\partial \Omega}{\partial R}}{\dfrac{\partial \Omega}{\partial \tilde{\tau}}}. \tag{3.16}$$

Equation (3.16) is the key equality that breaks the optimality of the Friedman rule. Because conventional taxes carry collection costs, it will be optimal to resort to the inflation tax. Currency substitution then introduces an additional layer when exchange controls are

inactive. The interplay between the cost of holding domestic and foreign relative money balances has bearings on the optimality and size of the different taxes. At an optimum and for given revenue needs, the government equates the marginal rate of substitution between the two tax instruments to the marginal rate of transformation per unit of consumption.

I now have a system of two equations in (3.14) and (3.16), which can be used to solve for the optimal values of $\tilde{\tau}$ and R as functions of the exogenous variables. Thereafter, one can solve for the optimal inflation tax, I, and labor income tax, τ. I begin by establishing an equivalence between the two exchange-control scenarios.

Equivalence Between Open and Closed Economy Models The goal of this paper is to analyze how the imposition of FX controls affects optimal seigniorage in an otherwise open economy with currency substitution. Moving forward, I will be referring to an economy without exchange controls ($\phi = 0$) as an *open economy*, while an economy with exchange controls ($\phi \to \infty$) will be a *closed economy*.

Lemma 1 (equivalence between models when $I^* = 0$). *Whenever the foreign nominal interest rate is zero, the model without exchange controls is equivalent to the model with controls.*

The above lemma arises as an outcome of the model because of the satiation property of the shopping-time technology, $V(\cdot)$. Intuitively, when $I^* = 0$, meaning that $R^* = 0$, the agent is at his satiation point in terms of foreign relative monetary balances, and thus any increase in the domestic nominal interest rate will not generate any further substituting away from the domestic currency. It is argued, therefore, that under this scenario all the implications of the model are analogous to those obtained in a closed economy setting.[10]

[10] A mathematical proof can be found in Appendix C.2.1

In other words, the two conditions that need to be satisfied (equations (3.14) and (3.16)) are the same between the two models. The usefulness of this lemma lies in the fact that, when comparing both models, it suffices to study the open economy version when (i) $I^* = 0$, and (ii) when $I^* > 0$. For the following analysis, I assume that $\phi = 0$.

Optimal Inflation Tax for Different Spending Levels Two propositions will be shown here, of which one yields a novel result: The Friedman rule is never optimal in an open economy whenever revenue needs are different from zero ($g > 0$). Previous literature has found that a positive foreign nominal interest rate yields an (optimally) positive domestic interest rate. Here, I show that the presence of convex tax-collection costs produces a positive inflation tax even when $I^* = 0$.

For the following propositions, it serves to present both the government budget constraint and the optimality condition (written out) again:

$$g = c\left[q(\tilde{\tau}, R, R^*)\right]\left[\tilde{\tau} + RX^m + \tilde{\tau}(R^* - I^*)X^f - \frac{\kappa(\tilde{\tau}y)}{c}\right]; \tag{3.17}$$

$$\frac{X^m}{1 + I^* X^f} = \frac{X^m + RX^m_R + \tilde{\tau} I^* X^f_R - \frac{1}{c}\frac{\partial \kappa(\tilde{\tau}y)}{\partial R}}{1 + I^* X^f + RI^* X^m_{R^*} + \tilde{\tau} I^{*2} X^f_{R^*} - \frac{1}{c}\frac{\partial \kappa(\tilde{\tau}y)}{\partial \tilde{\tau}}}, \tag{3.18}$$

where

$$\tilde{\tau}y = c\left(\tilde{\tau} + \frac{\tilde{\tau}}{1 + \tilde{\tau}} RX^m + \tilde{\tau} I^* X^f\right). \tag{3.19}$$

Proposition 1 (No revenue needs). *The Friedman rule holds when there are no expenditures to be financed, regardless of the foreign nominal interest rate. Additionally, it follows that* $\tau = 0$.

Proof. It is easy to show that, if $g = 0$, equations (3.17) and (3.18) are both satisfied when $\tilde{\tau} = R = 0$. This, in turn, implies that $\tau = I = 0$. □

Logically, since both tax instruments are welfare-reducing, it is never optimal to tax domestic relative money balances. Setting $I = 0$ implies that the private and social costs of holding relative money balances are equal. While Proposition 1 has been thoroughly documented in the literature, the next result offers a novel result.

Proposition 2 (**Friedman rule is never optimal when** $g > 0$)**.** *The optimal inflation tax is positive, whenever* $g > 0$*, regardless of the foreign nominal interest rate.*

Proof. I demonstrate this sequentially by showing that no other combination of taxes is desired and/or attainable. Let I^* take any value, *including* zero. Consider first the case when $R = 0$. If, in addition, $\tilde{\tau} = 0$, equation (3.17) is violated because $g \not> 0$, for any value of I^*. Similarly, if $\tilde{\tau} > 0$, equation (3.18) boils down to

$$\frac{X^m}{1+I^*X^f} < \frac{X^m\left[1-\frac{\tilde{\tau}}{1+\tilde{\tau}}\kappa'(\cdot)\right]+\tilde{\tau}I^*X^f_R[1-\kappa'(\cdot)]}{(1+I^*X^f)[1-\kappa'(\cdot)]+\tilde{\tau}I^{*2}X^f_{R^*}[1-\kappa'(\cdot)]}.$$

If $I^* > 0$ it can be shown that the right-hand side will always be bigger since $\kappa'(\cdot) > 0$ and because the right-most terms on each side of the division both make the right-hand side term larger. Alternatively, when $I^* = 0$, equation (3.18) when $I^* = 0$ reads

$$X^m < \frac{X^m\left[1-\frac{\tilde{\tau}}{1+\tilde{\tau}}\kappa'(\cdot)\right]}{1-\kappa'(\cdot)},$$

where it follows that, since $\kappa'(\cdot) > 0$, the right-hand side is always larger.

The presence of convex tax-collection costs breaks the equality in both cases. It is thus established that $R > 0$. Suppose now that $\tilde{\tau} = 0$. Equation (3.18) now yields

$$\frac{X^m}{1+I^*X^f} > \frac{X^m+RX^m_R}{1+I^*X^f+RI^*X^m_{R^*}}.$$

Let $I^* > 0$. It is easily shown that the right-most terms in the numerator and denominator

of the right-hand side make the fraction smaller, and therefore the left-hand side is always larger. Using the same reasoning for the case when $I^* = 0$, it can also be shown that the same inequality arises:

$$X^m > X^m + RX_R^m.$$

This sequential proof thus shows that $\tilde{\tau}$ and R are positive at an optimum, which in turn imply that the inflation tax, I, and the labor income tax, τ, are also positive.

□

Proposition 2 yields a novel result that follows from the nature of the collection costs: *the inflation tax is optimally greater than zero for any level of* $g > 0$. Imagine a situation in which the foreign nominal interest rate is zero. In this case, the private and social marginal cost of holding foreign relative money balances are both equal to zero. While other models of optimal seigniorage and currency substitution would call for a domestic nominal interest rate of zero, the presence of increasing marginal tax collection costs generates an incentive for deadweight-loss-sharing among the two taxes. On the flip side, a non-zero foreign nominal interest rate coupled with a positive labor income tax generates a wedge between the private (R^*) and social (I^*) marginal costs of holding foreign currency, which can be addressed by raising I. In other words, following an increase in the labor income tax, foreign money holdings are initially too low. A higher nominal interest rate is thus called for.

Moreover, and following the equivalence established previously, it follows that positive financing needs call for a positive (optimal) inflation tax, both with and without foreign exchange controls.

So far, we have results on the signs of the different taxes, but not their magnitude. The following results aim to shed some light on the different dynamics at play in both economies.

Effect of an Increase in Government Spending Will the two economies meet their financing needs equally? I now show that the presence of the foreign nominal interest rate, whenever positive, will drive a higher optimal domestic inflation tax via a currency substitution motive.

Proposition 3 (Effects of an increase in government spending). *In an initial equilibrium in which both government spending and the foreign nominal interest rate are zero, the optimal inflation tax directly reacts to changes in government spending. Additionally, around an initial equilibrium in which $g = 0$ and $I^* > 0$, the optimal inflation tax is an increasing function of government spending.*

Proof. It has already been shown in Proposition 1 that $I(g = 0, I^* = 0) = 0$ and $\tau(g = 0, I^* = 0) = 0$. Therefore, by totally differentiating equations (3.17) and (3.18) with $I^* = 0$, evaluating them at $R = \tilde{\tau} = 0$, and combining, we obtain:

$$\left.\frac{\partial R(g=0,I^*=0)}{\partial g}\right|_{R=\tilde{\tau}=0} = \frac{\kappa''(\cdot)}{c\left[\kappa''(\cdot)X^m - X^m_R\right]} > 0; \tag{3.20a}$$

$$\left.\frac{\partial \tilde{\tau}(g=0,I^*=0)}{\partial g}\right|_{R=\tilde{\tau}=0} = -\frac{X^m_R}{c\left[\kappa''(\cdot)X^m c - X^m_R\right]} > 0, \tag{3.20b}$$

since $X^m_R < 0$. For the second part, it has also been shown that $I(g = 0, I^* > 0) = 0$ and $\tau(g = 0, I^* > 0) = 0$. Again, totally differentiating equations (3.17) and (3.18), but allowing I^* to be positive, yields

$$\left.\frac{\partial R(g=0,I^*>0)}{\partial g}\right|_{R=\tilde{\tau}=0} = \frac{I^* X^f_R(1+I^*X^f) - X^m I^{*2} X^f_{R^*} + \kappa''(\cdot)(1+I^*X^f)^2}{\Theta} > 0; \tag{3.21a}$$

$$\left.\frac{\partial \tilde{\tau}(g=0,I^*>0)}{\partial g}\right|_{R=\tilde{\tau}=0} = \frac{I^* X^f_R X^m - X^m_R(1+I^*X^f)}{\Theta} > 0, \tag{3.21b}$$

where

$$\Theta = C\Big[X^m I^* X_R^f (1+I^* X^f) + X^m I^* X_{R^*}^m (1+I^* X^f) - (X^m I^*)^2 X_{R^*}^f \\ - X_R^m (1+I^* X^f)^2 + X^m \kappa''(\cdot)(1+I^* X^f)^2\Big] > 0, \tag{3.22}$$

and where these equations have been signed using the money demand properties outlined in (3.9). Finally, both τ and I can be recovered, yielding the same results. □

Proposition 3 yields an insightful result about two main components of the model. Whenever the foreign nominal interest rate is zero, the convexity of the tax collection costs drives the optimal positive inflation tax. The reason behind this result lies in the satiation property of the shopping-time function, V.

The second part shows what happens when currency substitution kicks in. The labor income tax is raised following an increase in spending, which, with a positive foreign nominal interest rate, increases the wedge between the private and social cost of holding foreign relative monetary balances. To counteract this, the government finds it optimal to increase the inflation tax, effectively inducing the consumer to use more foreign money.

This model is able to reproduce two findings of the literature—albeit found in closed economy frameworks. First, equations (3.20a) and (3.20b) show that, for models with constant marginal tax-collection costs, an increase in government spending will be fully financed with the labor income tax. Note that, while Friedman rule fails to hold in these models because of the presence of tax-collection costs, the (positive) optimal inflation tax is not affected BY changes in g.

Second, equations (3.20a) and (3.20b) also allow us to see how the optimal combination of the two taxes changes as the tax system becomes weaker—proxied here by an increase in $\kappa''(\cdot)$. Namely, countries with weaker tax systems will rely more heavily on the inflation tax because the deadweight loss generated by conventional taxes is greater. The present

model can shed some light for the open economy case: weaker tax systems yield a higher optimal inflation tax. This can more easily be seen in a model with an informal tax-evading sector, such as Koreshkova (2006). Conventional taxes become less effective at collecting revenue as the informal sector grows, and thus the inflation tax serves as a tax on the informal sector—which mainly functions with cash.

There is another key implication of this model that has already been documented in open economy models without tax-collection costs. Equation (3.22) shows that a higher foreign nominal interest rate brings about a higher optimal inflation tax. The mechanism is both through the convex tax-collection costs and through the combination of the medium of transaction and store of value motives that incite the household to substitute between currencies. Intuitively, for a higher foreign nominal interest rate, an increase in g, via the subsequent increase in τ, brings about a larger wedge between the private (R^*) and social (I^*) costs of holding foreign currency. This in turn calls for an even higher optimal inflation tax than what would otherwise prevail with a lower foreign nominal interest rate. On top of this effect, the deadweight loss generated by a higher labor income tax also contributes to a higher I. Summing up, there are two forces driving the increase in the optimal domestic nominal interest rate in the open economy.

Lastly, and following the discussion presented in this section, I can derive one last result, which compares countries with and without foreign exchange controls.

Proposition 4 (**Optimal inflation tax with and without exchange controls**). *Given the equivalence of the economies with and without foreign exchange controls whenever $I^* = 0$, and the fact that the optimal inflation tax is an increasing function with respect to the foreign nominal interest rate, the following result holds:*

$$I^{FXc} \geq I^{noFXc}, \tag{3.23}$$

where I^{FXc} and I^{noFXc} are the prevailing optimal domestic nominal interest rates in the economies with and without foreign exchange controls, respectively.

Proposition 4 states that, by imposing a ban on the use of foreign currency, effectively reducing the problem to a closed economy one, the government has greater control over the monetary base, and can thus achieve a given level of spending with a lower interest rate. On the contrary, as the economy opens up and multiple currencies coexist, the presence of a foreign interest rate will have additional implications on the optimality of the domestic rate of interest.

3.3.3 On the Desirability of Foreign Exchange Controls

When could it be desirable, in light of this model, to impose controls on the use of foreign exchange? The answer is not straightforward. However, what this framework suggests is that exchange controls could be advantageous during periods of high foreign nominal interest rates. Since a high I^* implies that increases in government spending are financed mostly with the inflation tax, exchange controls would allow the economy to isolate itself from the rest of the world, effectively reducing the inflationary burden.

This has negative implications, however, which are not modeled here and could play a counteracting role. Firstly, in a tightly restricted economy, there will be a welfare loss caused by the forceful misallocation of household's wealth. In other words, by shutting down currency substitution, a higher domestic interest rate will leave agents with suboptimal holdings of the domestic currency. Secondly, it is expected that agents will have incentives to bypass the aforementioned restrictions and acquire foreign currency in a parallel market as the cost of holding domestic money increases. This is left for future work.

With respect to the tax system, the decision of whether or not to limit the use of foreign currency is orthogonal to the weakness of the tax system. That is, a country with an in-

efficient tax-collecting capacity cannot improve its situation by adding exchange controls. Instead, strengthening institutions via a reduction in the inefficiencies of the tax system can have important implications with respect to output and welfare (Aruoba, 2021).

Finally, a richer model that endogenizes the tax on foreign money holdings, ϕ, could yield further insight into this question. In this paper I only focus on two limiting cases. However, in the data, many countries impose some restrictions on the use of foreign currency, but usually do not go as far as an outright ban on other currencies (IMF, 2022). I conjecture that intermediate levels of foreign exchange controls would yield similar but attenuated results.

3.4 Conclusion

This work has looked at the desirability of imposing restrictions on the use of foreign currencies using an open-economy optimal taxation framework. The Friedman rule is never optimal because conventional taxation carries collection costs. Furthermore, an economy with free flow of foreign monies optimally calls for a higher inflation tax than would otherwise prevail in a setting with exchange controls. The bottom line of this analysis is that limiting the use of foreign currency could become desirable as the foreign interest rate is high, since it concurrently calls for a high domestic nominal interest rate. Exchange controls would, for given revenue needs, effectively reduce the prevailing inflation rate.

Future research could look into black markets for foreign exchange once controls are in place. The assumption that agents lose total access to foreign currency is not entirely realistic, as an unavoidable inflation tax would create incentives to bypass said controls (Schmitt-Grohé and Uribe, 2023). One could also think of a richer model with more realistic firms, with more tax instruments available for the government, or one in which the exchange controls parameter is endogenous. Additionally, I have disregarded time incon-

sistency throughout, but one could argue that authorities would want to reassess their tax choices as agents substitute away from the domestic currency.

Appendix A: Appendix to Chapter 1

A.1 Data Appendix

This appendix provides additional details on the data, methodologies, and country coverage underlying the empirical analysis in Sections 1.2 and 1.3. Appendix A.1.1 describes the bond-level dataset, while Appendix A.1.2 outlines the parametric Svensson fit used to construct zero-coupon yield curves. Appendix A.1.3 then explains the estimation of nominal term premia. Appendix A.1.4 presents the high-frequency change in nominal term premia around monetary policy meetings. Appendix A.1.5 discusses index-linked bond data, and Appendix A.1.6 discusses the data sources of the macroeconomic and financial variables used to construct monetary policy shocks. Finally, Appendix A.1.7 presents additional figures and tables.

A.1.1 Sovereign bond data and interest rates

Data on sovereign local-currency nominal bonds are obtained from *Bloomberg* and *Refinitiv* since 2000. For each bond, I collect issue and maturity dates, coupon rates, payment schedules, and other contractual features (i.e., whether the bond is callable or has been called, etc). In addition, I extract daily bond prices and yields to maturity. The dataset is restricted to bonds with a remaining maturity of at most ten years; longer-term bonds enter the sample only once their residual maturity falls below this threshold. When available, zero-coupon yields are used directly to avoid re-estimation. Only bonds issued by

the government are included. Bonds issued by central banks or other public institutions are removed. Following much of the literature, callable bonds are excluded. Table A.1 summarizes the country coverage and start dates of the bond-level data used to construct daily zero-coupon nominal yield curves. The dataset initially covers 70 countries, of which 36 are classified as emerging markets according to the International Monetary Fund. Due to the availability of monetary policy surprise data, the final analysis sample includes 42 countries, of which 15 are classified as emerging markets, as presented in Table 1.1. The yield-curve data are available upon request.

Data on monetary policy rates is extracted from *Bloomberg* and *Refinitiv*.

Table A.1: Country Sample and Start Dates

Emerging Markets		**Advanced Economies**	
Country	Start Date	Country	Start Date
Armenia	May-2017	Australia	Jan-2000
Bangladesh	Jul-2016	Austria	Jan-2000
Botswana	Oct-2010	Belgium	Jan-2000
Brazil	Dec-2007	Canada	Jan-2000
Bulgaria	Dec-2010	Croatia	Feb-2011
Chile	Aug-2011	Cyprus	Apr-2014
China	Jan-2005	Czech Republic	Feb-2007
Colombia	Jul-2008	Denmark	Jan-2000
Egypt	Apr-2012	Finland	Jan-2000
Georgia	Jan-2013	France	Jan-2000
Ghana	Jan-2010	Germany	Jan-2000
Hungary	Sep-2010	Greece	Jan-2000
India	Dec-2008	Hong Kong	Feb-2011
Indonesia	Jan-2006	Iceland	Apr-2007
Kenya	Jan-2010	Ireland	Jan-2000
Malaysia	May-2009	Israel	Feb-2007
Mauritius	Feb-2011	Italy	Jan-2000
Mexico	Jul-2002	Japan	Jan-2000
Morocco	Nov-2014	Lithuania	Feb-2011
Namibia	Apr-2015	Malta	Jan-2000
Nigeria	Jan-2010	Netherlands	Jan-2000
Pakistan	Oct-2010	New Zealand	Jan-2000
Peru	May-2009	Norway	Feb-2000
Philippines	Dec-2010	Portugal	Jan-2000
Poland	Jan-2000	Singapore	Jan-2007
Romania	Feb-2007	Slovakia	Nov-2006
Russia	Oct-2006	Slovenia	Feb-2011
Serbia	Feb-2011	South Korea	Feb-2011
South Africa	Feb-2011	Spain	Jan-2000
Sri Lanka	Oct-2010	Sweden	Jan-2000
Thailand	Apr-2004	Switzerland	Jan-2000
Türkiye	Mar-2007	Taiwan	Sep-2008
Uganda	Feb-2011	United Kingdom	Jan-2000
Venezuela	Jul-2010	United States	Jan-2000
Vietnam	Jan-2006		
Zambia	Nov-2014		

A.1.2 Parametric Yield Curve Fitting

A standard approach to constructing smooth zero-coupon yield curves from bond price data is to use parsimonious parametric functions. The most widely applied are the models of Nelson and Siegel (1987) and its extension by Svensson (1994). These models are particularly attractive because they capture the main stylized shapes of the yield curve (monotonic, humped, or S-shaped) with only a handful of parameters, while ensuring smoothness and economic interpretability. It is because of these reasons that central banks around the world rely on these methods to construct yield curves (BIS, 2014).

Nelson-Siegel (1987) The Nelson and Siegel (1987) specification assumes that the instantaneous forward rate at maturity m can be written as

$$f(t,t+m;b) = \beta_0 + \beta_1 \exp\left(-\frac{m}{\tau_1}\right) + \beta_2 \left(\frac{m}{\tau_1}\right) \exp\left(-\frac{m}{\tau_1}\right), \tag{A.1}$$

where β_0 captures the long-run level of rates, β_1 governs the short-end slope, and β_2 controls the magnitude of a medium-term "hump" in the curve. The parameter τ_1 is a decay factor that determines the location of this hump. The corresponding spot rate on a zero-coupon bond maturing in m years is obtained by integrating the forward curve and then dividing the result by the remaining time to maturity:

$$i(t,t+m;b) = \beta_0 + \beta_1 \frac{1-\exp(-m/\tau_1)}{m/\tau_1} + \beta_2 \left(\frac{1-\exp(-m/\tau_1)}{m/\tau_1} - \exp(-m/\tau_1)\right). \tag{A.2}$$

Svensson (1994) To improve flexibility and allow for more complex curve shapes, Svensson (1994) introduced an additional term, adding a second hump:

$$f(t,t+m;b) = \beta_0 + \beta_1 \exp\left(-\frac{m}{\tau_1}\right) + \beta_2 \left(\frac{m}{\tau_1}\right) \exp\left(-\frac{m}{\tau_1}\right) + \beta_3 \left(\frac{m}{\tau_2}\right) \exp\left(-\frac{m}{\tau_2}\right). \tag{A.3}$$

Here β_3 and τ_2 introduce a second curvature component, enabling the model to fit yield curves with both short- and long-term humps or U-shapes. Because of its flexibility and tractability, the Svensson (1994) specification has become a standard in central banks and international financial institutions for constructing zero-coupon yield curves. The corresponding spot rate is

$$i(t,t+m;b) = \beta_0 + \beta_1 \frac{1-\exp(-m/\tau_1)}{m/\tau_1} + \beta_2 \left(\frac{1-\exp(-m/\tau_1)}{m/\tau_1} - \exp(-m/\tau_1) \right) + \beta_3 \left(\frac{1-\exp(-m/\tau_2)}{m/\tau_2} - \exp(-m/\tau_2) \right). \tag{A.4}$$

The discount function is related to the spot rate by

$$d(t,t+m;b) = \exp\left(-\frac{i(t,t+m;b)}{100} \cdot m \right). \tag{A.5}$$

The discount function is used to compute the estimated (theoretical) bond price. The price of a coupon bond can be defined, in analogy to the traditional valuation of an investment, as the sum of discounted future coupon payments C and the present value of the face value paid after m years. It follows that the price of the bond $P(t,t+m)$ on trade day t is approximated by

$$P^e(t,t+m;b) = \sum_{k=1}^{m} C \cdot d(t,t+k;b) + 100 \cdot d(t,t+m;b). \tag{A.6}$$

For coupon bonds, yields to maturity are often quoted. The yield to maturity is the internal rate of return that equates the present value of the coupon payments and the face value with the observed price of the bond. Thus, the price of a coupon bond can equivalently be written as a function of the yield to maturity $y(t,t+m)$:

$$P(t,t+m) = \sum_{k=1}^{m} C \cdot \exp\left(-\frac{y(t,t+m)}{100} k \right) + 100 \cdot \exp\left(-\frac{y(t,t+m)}{100} m \right). \tag{A.7}$$

Data selection The input data consist of daily observations of sovereign bond yields and their coupon structures, with maturities ranging up to ten years. Very short maturities are covered by money-market instruments, while callable bonds and instruments with residual maturity below one year are excluded. Longer-dated coupon bonds enter the sample only once their remaining maturity is below ten years. For each day and each country, the bond universe provides the cross-section of prices and yields used in the estimation.

Estimation procedure The discount function is estimated for each trade date by minimizing either the sum of squared price errors or the sum of squared yield errors. Following standard practice, the estimation is constrained so that the forward rate curve (and hence the spot rate curve) starts at the short end. In other words, the term $(\beta_0 + \beta_1)$ is set equal to the prevailing short rate (monetary policy rate), thereby anchoring the curve at the policy-controlled end of the term structure.

For each trade day, the parameter vector $(\beta_0, \beta_1, \beta_2, \beta_3, \tau_1, \tau_2)$ is chosen to minimize the weighted sum of squared bond price errors:

$$\min_{\beta_0,\beta_1,\beta_2,\beta_3,\tau_1,\tau_2} \sum_{j=1}^{n} \left\{ \frac{P_j - P_j^e(\beta_0, \beta_1, \beta_2, \beta_3, \tau_1, \tau_2)}{\Phi_j} \right\}^2, \tag{A.8}$$

where P_j is the observed price of bond j on trade date t (from (A.7)), and P_j^e is the corresponding model-implied price (from (A.6)). The weighting factor Φ_j is defined as

$$\Phi_j = \frac{duration_j \cdot P_j}{1 + y_j}, \tag{A.9}$$

where $duration_j$ refers to the time to maturity and y_j is the observed yield to maturity of bond j. This weighting accounts for the sensitivity of prices to yield changes across maturities and normalizes errors across bonds.

The resulting zero-coupon nominal yield curves are then compared to existing bench-

mark curves (e.g., Gürkaynak et al. (2007) for the U.S. or the Bank for International Settlements for selected European countries) to validate the fit, ensure plausibility at the short and long ends, and guarantee consistency with established international standards.

Interpretation The estimated Svensson (1994) curve provides a smooth and flexible representation of the sovereign term structure. By construction, it nests the Nelson and Siegel (1987) specification as a special case ($\beta_3 = 0$). Its additional curvature term allows for the presence of multiple humps or U-shapes along the maturity profile, which improves the empirical fit when yield curves display complex dynamics. The resulting zero-coupon nominal yields are used to estimate daily nominal term premia.

A.1.3 Affine Term Structure Model of Interest Rates

This subsection describes how daily nominal term premia are constructed from the fitted zero-coupon yield curves obtained in Appendix A.1.2. For each country, I apply the regression-based affine term structure model of Adrian et al. (2013) to decompose zero-coupon nominal yields into two components: the model-implied average expected future path of short-term interest rates and a residual term-premium component. The model is estimated separately by country using the time series of fitted zero-coupon yields across maturities. For each date t and maturity n, the nominal term premium is defined as the difference between the fitted n-period zero-coupon yield and the corresponding expectations component implied by the model. This procedure produces the daily term-premium series used throughout Chapter 1.

The implementation proceeds in four steps: factor extraction from the zero-coupon yield curve, estimation of factor dynamics, recovery of risk prices from bond excess returns, and decomposition of fitted yields into expectations and term-premium components.

Model setup Let X_t denote a $K \times 1$ vector of pricing factors. In the empirical implementation, these factors are the first K principal components extracted from the cross section of fitted zero-coupon yields:

$$X_{t+1} = \mu + \Phi X_t + v_{t+1}, \qquad v_{t+1} \sim \mathcal{N}(0, \Sigma), \tag{A.10}$$

where μ is a constant, Φ is a $K \times K$ transition matrix, and Σ is the covariance matrix of factor innovations.

Bond prices in affine term structure models are consistent with an exponentially affine stochastic discount factor. Let $P_t^{(n)}$ denote the price at time t of an n-period zero-coupon bond. Under no-arbitrage, there exists a stochastic discount factor (pricing kernel) M_{t+1}

such that

$$P_t^{(n)} = \mathbb{E}_t\left[M_{t+1}P_{t+1}^{(n-1)}\right]. \tag{A.11}$$

The pricing kernel is assumed to take the exponentially affine form

$$M_{t+1} = \exp\left(-r_t - \tfrac{1}{2}\lambda_t'\lambda_t - \lambda_t'\Sigma^{-1/2}v_{t+1}\right), \tag{A.12}$$

where r_t denotes the continuously compounded one-period short rate, v_{t+1} are factor innovations with covariance matrix Σ, and λ_t denotes the market price of risk. Following Duffee (2002), the market price of risk is specified to be affine in the state variables:

$$\lambda_t = \Sigma^{-1/2}\left(\lambda_0 + \lambda_1 X_t\right), \tag{A.13}$$

with λ_0 capturing the constant component and λ_1 the time-varying component of risk premia.

The log excess holding-period return on an n-period zero-coupon bond held for one period is defined as

$$rx_{t+1}^{(n-1)} = \ln P_{t+1}^{(n-1)} - \ln P_t^{(n)} - r_t, \tag{A.14}$$

where r_t is the short rate. Under no-arbitrage and assuming joint normality of $\{rx_{t+1}^{(n-1)}, v_{t+1}\}$, expected excess returns are affine in the state variables:

$$\mathbb{E}_t\left[rx_{t+1}^{(n-1)}\right] = \beta_t^{(n-1)\prime}(\lambda_0 + \lambda_1 X_t) - \tfrac{1}{2}\beta_t^{(n-1)\prime}\Sigma\beta^{(n-1)}, \tag{A.15}$$

where the factor loadings are given by

$$\beta_t^{(n-1)} = \mathrm{cov}_t\left(rx_{t+1}^{(n-1)}, v_{t+1}'\right)\Sigma^{-1}. \tag{A.16}$$

The unexpected excess return can be decomposed into a component spanned by the factor innovations v_{t+1} and an idiosyncratic pricing error:

$$rx_{t+1}^{(n-1)} - \mathbb{E}_t\left[rx_{t+1}^{(n-1)}\right] = \beta_t^{(n-1)\prime} v_{t+1} + e_{t+1}^{(n-1)}, \tag{A.17}$$

where $e_{t+1}^{(n-1)}$ are return pricing errors, assumed conditionally i.i.d. with variance σ^2.

Stacking across maturities and time yields the panel regression system

$$rx = \beta'\left(\lambda_0 \iota_T' + \lambda_1 X_-\right) - \tfrac{1}{2}(B^\star \operatorname{vec}(\Sigma) + \sigma^2 \iota_N)\iota_T' + \beta' V + E, \tag{A.18}$$

where:

- rx is the $N \times T$ matrix of excess returns across N maturities and T periods,
- $\beta = [\beta^{(1)}\ \beta^{(2)}\ \ldots\ \beta^{(N)}]$ is the $K \times N$ matrix of factor loadings,
- ι_J denotes a $J \times 1$ vector of ones,
- $X_- = [X_0, X_1, \ldots, X_{T-1}]$ is the $K \times T$ matrix of lagged pricing factors,
- $B^\star = [\operatorname{vec}(\beta^{(1)}\beta^{(1)\prime}), \ldots, \operatorname{vec}(\beta^{(N)}\beta^{(N)\prime})]$ is an $N \times K^2$ matrix,
- V is the $K \times T$ matrix of estimated factor innovations,
- E is the $N \times T$ matrix of pricing errors.

Estimation of the affine term structure model Following Adrian et al. (2013), model parameters are estimated using a computationally efficient three-step OLS procedure:

1. **VAR estimation of factors**

Estimate the factor dynamics

$$X_{t+1} = \mu + \Phi X_t + v_{t+1}$$

via OLS. This yields estimates $\hat{\mu}, \hat{\Phi}$, factor innovations $\hat{v}_{t+1}$, and $\hat{\Sigma} = \frac{1}{T}\sum \hat{v}_{t+1}\hat{v}'_{t+1}$.

2. **Return regressions**

   Regress excess returns across maturities on a constant, factor innovations, and lagged factors:

$$rx = \mathbf{a}\,\iota_T' + \mathbf{fi}'\hat{V} + \mathbf{c}X_{-} + E, \tag{A.19}$$

   where $\hat{V}$ stacks the estimated innovations, X_{-} stacks lagged factors, and β are the factor loadings. Estimation yields $\hat{\mathbf{a}}, \hat{\beta}, \hat{\mathbf{c}}$ and return pricing errors E.

3. **Cross-sectional regression for prices of risk**

   Using the identities

$$\mathbf{a} = \beta'\lambda_0 - \tfrac{1}{2}(B^* \operatorname{vec}(\Sigma) + \sigma^2 \iota_N), \qquad \mathbf{c} = \beta'\lambda_1,$$

   the market price of risk parameters (λ_0, λ_1) are obtained by regressing estimated exposures onto the fitted coefficients from step (2):

$$\hat{\lambda}_0 = (\hat{\beta}\hat{\beta}')^{-1}\hat{\beta}\left(\hat{\mathbf{a}} + \tfrac{1}{2}(\hat{B}^* \operatorname{vec}(\hat{\Sigma}) + \hat{\sigma}^2 \iota_N)\right), \tag{A.20}$$

$$\hat{\lambda}_1 = (\hat{\beta}\hat{\beta}')^{-1}\hat{\beta}\hat{\mathbf{c}}. \tag{A.21}$$

   where $\hat{\sigma}^2 = \operatorname{tr}(\hat{E}\hat{E}')/NT$, and $\hat{B}^\star$ is constructed from $\hat{\beta}$.

Affine yields and term premia Given parameter estimates $(\Phi, \Sigma, \lambda_0, \lambda_1)$, affine recursions generate fitted bond prices:

$$A_n = A_{n-1} + B'_{n-1}(\mu - \lambda_0) + \tfrac{1}{2}(B'_{n-1}\Sigma B_{n-1} + s^2) - \delta_0, \tag{A.22}$$

$$B'_n = B'_{n-1}(\Phi - \lambda_1) - \delta'_1, \tag{A.23}$$

with initial conditions $A_0 = 0, B_0 = 0$. The model-implied zero-coupon yield is then

$$y_t^{(n)} = -\tfrac{1}{n}(A_n + B'_n X_t). \tag{A.24}$$

Risk-neutral yields are obtained by setting $\lambda_0 = \lambda_1 = 0$, so that bond prices reflect only the expected future path of short rates under the model and not compensation for bearing interest-rate risk. Accordingly, for each maturity n, the nominal term premium is computed as

$$TP_t^{(n)} = y_t^{(n)} - y_t^{(n),RN}, \tag{A.25}$$

where $y_t^{(n),RN}$ denotes the corresponding risk-neutral yield. Equivalently, this object measures the component of the fitted yield that is not accounted for by the model-implied expected path of future short-term interest rates.

Data and implementation For each country, I estimate two-, three-, four-, and five-factor specifications using the zero-coupon yield data constructed in Appendix A.1.2. The cross section consists of fitted zero-coupon yields with maturities $n = 3, \ldots, 120$ months, from which I extract principal components to serve as pricing factors. Excess holding-period returns are then computed for zero-coupon bonds with maturities $n = 12, 24, \ldots, 120$ months using (A.14), yielding a cross section of $N = 10$ maturities.

All models are estimated separately by country over the period 2000-2022, when data

are available, using daily observations. Conditional on the selected number of pricing factors, I estimate the parameters $(\Phi, \Sigma, \sigma, \beta, \lambda_0, \lambda_1)$ using the three-step procedure described above. I then estimate the short-rate parameters (δ_0, δ_1) by regressing the one-month yield on the pricing factors. Finally, I use the estimated parameters in the affine recursions in (A.22) and (A.23) to obtain model-implied yields and risk-neutral yields at each date and maturity, from which daily nominal term premia are constructed.

Model specification tests For each country, I select the number of pricing factors by comparing the average daily root mean squared error (RMSE) of the model-implied 10-year yield relative to its observed counterpart across alternative specifications. Using daily data over an extended sample alleviates concerns about small-sample bias highlighted by Bauer et al. (2014). The results are broadly consistent with the existing literature on term premia (Turner, 2014; Naudon and Yany, 2016; Cohen et al., 2018). Table A.2 reports, for each country, the number of principal components selected and the corresponding average daily RMSE for the 10-year bond yield.

Table A.2: Principal Components Selected and RMSE by Country

Emerging Markets			**Advanced Economies**		
Country	PCs	RMSE	Country	PCs	RMSE
Armenia	4	0.28	Australia	4	0.17
Bangladesh	5	0.53	Austria	3	0.18
Botswana	3	0.57	Belgium	2	0.18
Brazil	2	0.30	Canada	3	0.20
Bulgaria	2	0.41	Croatia	2	0.32
Chile	2	0.13	Cyprus	4	0.37
China	4	0.13	Czech Republic	2	0.33
Colombia	2	0.38	Denmark	4	0.16
Egypt	3	0.55	Finland	4	0.17
Georgia	3	0.90	France	4	0.17
Ghana	5	1.40	Germany	4	0.18
Hungary	2	0.29	Greece	2	2.74
India	3	0.11	Hong Kong	2	0.08
Indonesia	3	0.32	Iceland	3	0.21
Kenya	2	0.44	Ireland	2	0.42
Malaysia	3	0.11	Israel	2	0.24
Mauritius	2	0.27	Italy	4	0.17
Mexico	2	0.63	Japan	5	0.17
Morocco	2	0.20	Lithuania	2	0.33
Namibia	2	0.30	Malta	3	0.25
Nigeria	2	0.96	Netherlands	3	0.19
Pakistan	2	0.70	New Zealand	3	0.21
Peru	2	0.42	Norway	3	0.18
Philippines	2	0.23	Portugal	4	0.49
Poland	2	0.31	Singapore	4	0.14
Romania	2	0.29	Slovakia	2	0.24
Russia	2	0.35	Slovenia	3	0.52
Serbia	4	0.35	South Korea	2	0.11
South Africa	2	0.48	Spain	3	0.22
Sri Lanka	5	0.79	Sweden	5	0.13
Thailand	3	0.18	Switzerland	2	0.12
Türkiye	3	0.42	Taiwan	3	0.07
Uganda	5	2.08	United Kingdom	2	0.15
Venezuela	4	10.76	United States	5	0.13
Vietnam	3	0.89			
Zambia	4	26.32			

A.1.4 High-Frequency Event-Study Around Monetary Policy Announcements

To verify that the panel regression results are not driven by model specification or by particular countries, I complement the monthly panel analysis with a high-frequency event-study approach. Specifically, I study the two-day change in the 10-year nominal term premium around monetary policy announcements.[1]

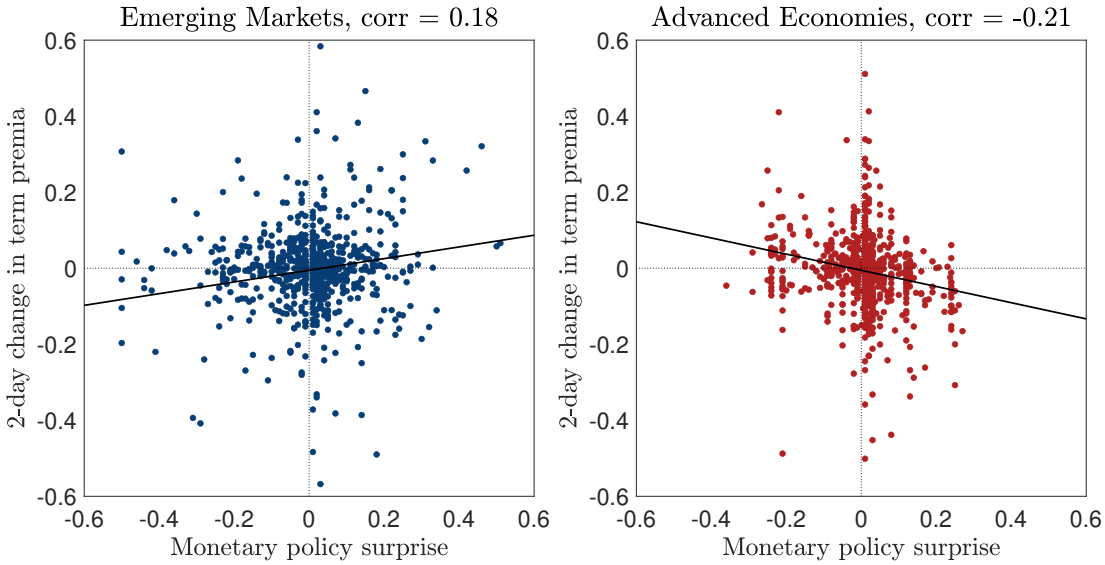
First, Figure A.1 shows that, when looking across all policy announcement days with identified surprises, the aggregate results mirror those from the panel regressions in Table 1.2: a surprise tightening is associated with an increase in term premia for emerging markets, and with a decline in advanced economies.[2] Second, Figure A.2 shows that this pattern holds widely across individual emerging markets. In nearly all cases, monetary policy surprises are positively associated with changes in term premia, with only Chile, Mexico, and Romania showing negative correlations.[3] This underscores that, unlike in advanced economies where surprises tend to reduce perceived risk, in emerging markets they are more often interpreted as heightening uncertainty and raising the compensation investors demand to hold long-term debt.

[1] Analyzing the change in financial variables in narrow windows around policy decisions is a standard approach to isolate the immediate effects of monetary policy surprises from other macroeconomic news. For example, Witheridge (2024) examines the two-day change in market-based inflation expectations (the breakeven inflation rate) around monetary policy announcements.

[2] See Table A.9 for regression results, confirming these patterns.

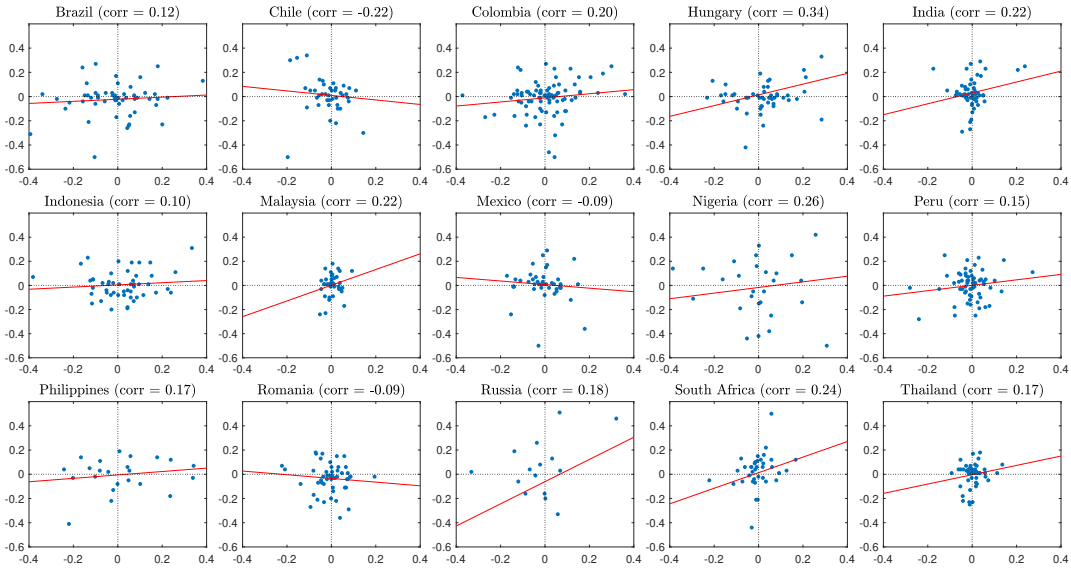
[3] In Figures A.5 and A.6, I replicate the same analysis for advanced economies and find that most countries individually exhibit negative correlations between surprises and nominal term premia.

Figure A.1: Change in 10-Year Nominal Term Premium Around Policy Announcements



Note: These figures plot the two-day change (in basis points) in the 10-year nominal term premium around monetary policy announcement dates, separately for emerging markets (left, blue) and advanced economies (right, red). Each dot corresponds to a policy announcement; the solid black line is the fitted regression. The sample is restricted to meetings with non-zero surprises and covers the period from 2010 onward. See Appendix A.1 for details on data, methodology, and country selection, and Table A.9 for regression results.

Figure A.2: Change in 10-Year Nominal Term Premium Around Policy News, by Country



Note: These figures plot the two-day change (in basis points) in the 10-year nominal term premium around monetary policy announcement dates, separately for each of the 15 emerging markets in the sample. Each blue dot corresponds to a policy announcement; the solid red line is the fitted regression for that country. The sample period begins in 2010. See Appendix A.1 for details on data, methodology, and country selection.

A.1.5 Index-linked Bonds

Data on sovereign local-currency index-linked bonds are obtained from *Bloomberg* and *Refinitiv*, starting in 2000 where available. For each bond, I collect issue and maturity dates, coupon rates, payment schedules, and other contractual features (e.g., whether the bond is callable or has been called). As with the nominal sample, the dataset is restricted to bonds with a remaining maturity of at most ten years; longer-term bonds are included only once their residual maturity falls below this threshold. Only bonds issued directly by central governments are retained, while bonds issued by central banks or other public institutions are excluded. Following common practice, callable bonds are also excluded.

Due to data availability, index-linked yield curves are constructed for a smaller set of countries: six emerging markets (Brazil, Chile, Colombia, Mexico, South Africa, and Turkey) and eleven advanced economies (Australia, Belgium, Canada, France, Germany, Italy, Japan, South Korea, Spain, the United Kingdom, and the United States). For each of these countries, I apply the Svensson parametric methodology described in Appendix A.1.2 to obtain daily real zero-coupon yield curves for maturities up to ten years. The index-linked yield-curve data are available upon request.

Table A.3 reports descriptive statistics on government debt composition for these countries. While index-linked bonds have gained modest relevance in several advanced economies, their role is more pronounced in some emerging markets—particularly in Latin America—where episodes of high and volatile inflation have increased demand for inflation protection. In the selected emerging markets, about one-fourth of all government debt securities are index-linked, compared with about 8% on average in advanced economies.

The empirical analysis in Section 1.2.4 is conducted for five advanced economies (Australia, France, Italy, the United Kingdom, and the United States) and five emerging markets (Brazil, Chile, Colombia, Mexico, and South Africa). Among advanced economies, I re-

strict the sample to countries where index-linked bonds account for at least 5% of total government debt securities. For emerging markets, I focus on countries where ILB data are both available and reliable; accordingly, Türkiye is excluded from the final sample.

Table A.3: Government Debt Securities, Selected Countries (May 2025)

	Government debt securities		Foreign currency	Index-linked bonds	
	USD bn	% GDP	% of gov. DS	% of gov. DS	% of LC gov. DS
Brazil	696.2	32%	6.2%	22.7%	24.2%
Chile	166.0	49%	27.0%	33.1%	45.4%
Colombia	227.4	63%	19.9%	18.6%	23.2%
Mexico	731.6	41%	14.5%	39.9%	46.7%
South Africa	267.7	70%	9.2%	15.7%	17.3%
Türkiye	196.1	18%	38.6%	14.3%	23.3%
EM (weighted average)		**37%**	**14.9%**	**27.0%**	**31.7%**
Australia	561.1	32%		5.0%	5.0%
Belgium	549.9	85%	1.7%	0.3%	0.3%
Canada	883.4	41%	1.6%	3.9%	4.0%
France	2751.6	90%		9.4%	9.4%
Germany	2214.1	49%		3.4%	3.4%
Italy	2725.3	118%	0.6%	14.7%	14.8%
Japan	7626.5	181%		1.9%	1.9%
South Korea	964.7	56%	0.9%	0.5%	0.5%
Spain	1621.3	100%		4.8%	4.8%
United Kingdom	3341.2	99%		16.0%	16.0%
United States	22388.7	81%		9.3%	9.3%
AE (weighted average)		**86%**	**0.1%**	**8.0%**	**8.0%**

Note: This table presents a disaggregated view of government debt securities outstanding as of May 2025 for selected advanced and emerging economies. Data on government debt securities is sourced from *Refinitiv* and *Bloomberg*, while nominal GDP figures for 2024 are from the IMF's *World Economic Outlook*. Weighted averages are computed using total government debt as weights within each group.

A.1.6 Monetary policy shocks

To extract the information effects of monetary policy surprises, I follow Checo et al. (2024) and extract macroeconomic and financial data for each country, before running the specification in equation (1.6).

Price indicators Data on headline inflation and nominal wage growth are extracted from the International Financial Statistics (IMF), complementing with national sources when unavailable. One-year-ahead inflation expectations are from Consensus Economomics and Bloomberg. The inflation rate of imported commodities is sourced from Gruss and Kebhaj (2019).

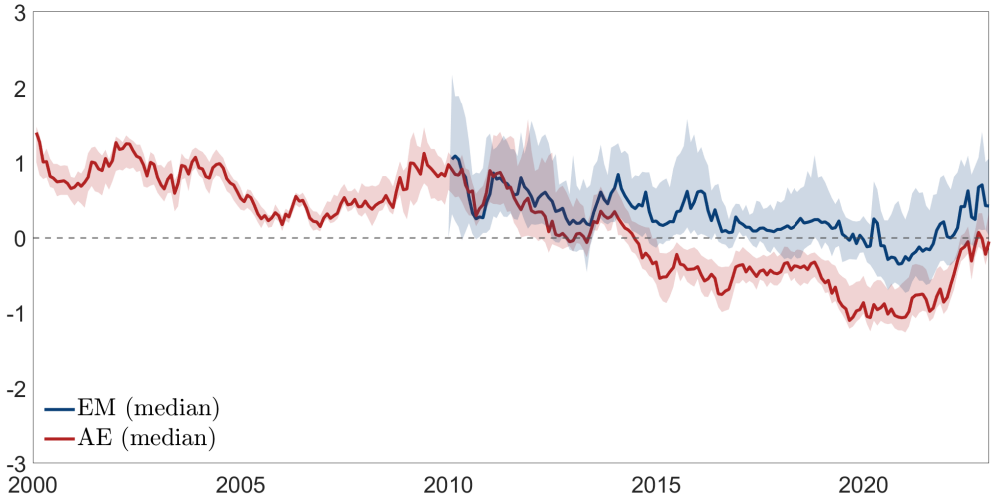
Activity indicators Data on the growth rate of industrial production and the unemployment rate are extracted from the International Financial Statistics (IMF), complementing with national sources when unavailable. The one-year-ahead growth rate of industrial production is from Consensus Economics.

Financial market variables The change in the exchange rate and in stock-market prices is sourced from Bloomberg and Refinitiv. For stock-market prices, the main stock-market index for each country is used. The one-year-ahead change in the exchange rate is from Consensus Economics.

A.1.7 Additional Figures and Tables

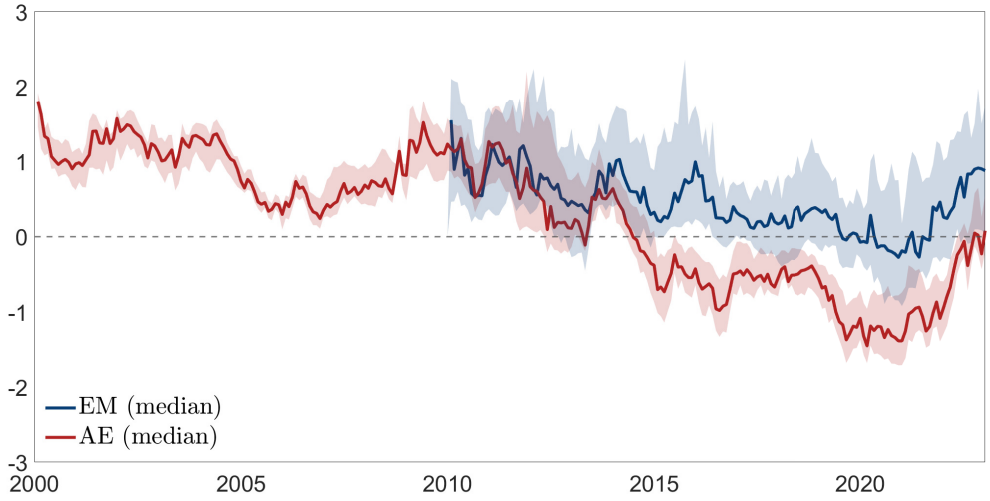
This section presents supplementary figures and tables that complement the main analysis in the paper.

Figure A.3: Median Nominal Term Premium on a 5-Year Sov. Bond, by Country Group



Note: This figure presents the median nominal term premium on a 10-year sovereign bond for emerging markets (blue line) from 2010 to 2022, and advanced economies (red line) from 2000 to 2022. The shaded bands indicate the interquartile range for each group. Term premia are estimated using the three-step no-arbitrage framework of Adrian et al. (2013) based on daily zero-coupon yield curves constructed from sovereign bond prices. See Appendix A.1 for further details on the sample, estimation methods, and country classifications.

Figure A.4: Median Nominal Term Premium on a 7-Year Sov. Bond, by Country Group



Note: This figure presents the median nominal term premium on a 10-year sovereign bond for emerging markets (blue line) from 2010 to 2022, and advanced economies (red line) from 2000 to 2022. The shaded bands indicate the interquartile range for each group. Term premia are estimated using the three-step no-arbitrage framework of Adrian et al. (2013) based on daily zero-coupon yield curves constructed from sovereign bond prices. See Appendix A.1 for further details on the sample, estimation methods, and country classifications.

Table A.4: Response of Term Premia to Surprise Tightenings (with default risk)

	Emerging Markets			Advanced Economies		
	5y	*7y*	*10y*	*5y*	*7y*	*10y*
Surprise tightening	0.316^{***}	0.303^{***}	0.285^{***}	-0.431^{***}	-0.511^{***}	-0.584^{***}
(1 p.p.)	(0.056)	(0.065)	(0.103)	(0.052)	(0.054)	(0.057)
Observations	1930	1930	1930	4023	4023	4023
Countries	15	15	15	27	27	27
R^2	0.72	0.73	0.61	0.92	0.94	0.95
Controls	Yes	Yes	Yes	Yes	Yes	Yes

Note: This table presents the estimated response of nominal term premia at 5-, 7-, and 10-year maturities to a 1 percentage point surprise monetary policy tightening, for emerging markets and advanced economies. The nominal term premium is defined as the excess yield investors require to hold sovereign long-term nominal bonds relative to short-term ones. All specifications include country fixed effects and three lags of the dependent variable and the policy rate surprise. They also control for sovereign default risk (measured by the government's default probability from *Refinitiv*) included contemporaneously and with three lags. The sample period begins in 2010. Significance stars indicate: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. See Appendix A.1 for further details on data and methodology.

Table A.5: Response of Term Premia to Surprise Tightenings (with global controls)

	Emerging Markets			Advanced Economies		
	5y	*7y*	*10y*	*5y*	*7y*	*10y*
Surprise tightening	0.307^{***}	0.278^{***}	0.247^{**}	-0.427^{***}	-0.511^{***}	-0.583^{***}
(1 p.p.)	(0.057)	(0.066)	(0.105)	(0.051)	(0.056)	(0.058)
Observations	1930	1930	1930	4023	4023	4023
Countries	15	15	15	27	27	27
R^2	0.72	0.73	0.61	0.92	0.94	0.95
Controls	Yes	Yes	Yes	Yes	Yes	Yes

Note: This table presents the estimated response of nominal term premia at 5-, 7-, and 10-year maturities to a 1 percentage point surprise monetary policy tightening, for emerging markets and advanced economies. The nominal term premium is defined as the excess yield investors require to hold sovereign long-term nominal bonds relative to short-term ones. All specifications include country fixed effects and three lags of the dependent variable and the policy rate surprise. They also control for the U.S. term premium (Naudon and Yany, 2016) and the VIX, included contemporaneously and with three lags. The sample period begins in 2010. Significance stars indicate: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. See Appendix A.1 for further details on data and methodology.

Table A.6: Response of Term Premia to Surprise Tightenings (with one lag)

	Emerging Markets			Advanced Economies		
	5y	*7y*	*10y*	*5y*	*7y*	*10y*
Surprise tightening	0.285^{***}	0.261^{***}	0.216^{**}	-0.432^{***}	-0.510^{***}	-0.567^{***}
(1 p.p.)	(0.052)	(0.061)	(0.097)	(0.052)	(0.054)	(0.057)
Observations	1962	1962	1962	4055	4055	4055
Countries	15	15	15	27	27	27
R^2	0.73	0.72	0.60	0.92	0.94	0.95

Note: This table presents the estimated response of nominal term premia at 5-, 7-, and 10-year maturities to a 1 percentage point surprise monetary policy tightening, for emerging markets and advanced economies. The nominal term premium is defined as the excess yield investors require to hold sovereign long-term nominal bonds relative to short-term ones. All specifications include country fixed effects and one lag of the dependent variable and the policy rate surprise. The sample period begins in 2010. Significance stars indicate: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. See Appendix A.1 for further details on data and methodology.

Table A.7: Response of Term Premia to Surprise Tightenings (with six lags)

	Emerging Markets			Advanced Economies		
	5y	*7y*	*10y*	*5y*	*7y*	*10y*
Surprise tightening	0.308^{***}	0.299^{***}	0.303^{***}	-0.429^{***}	-0.523^{***}	-0.591^{***}
(1 p.p.)	(0.061)	(0.071)	(0.112)	(0.052)	(0.054)	(0.057)
Observations	1882	1882	1882	3975	3975	3975
Countries	15	15	15	27	27	27
R^2	0.71	0.72	0.62	0.92	0.94	0.95

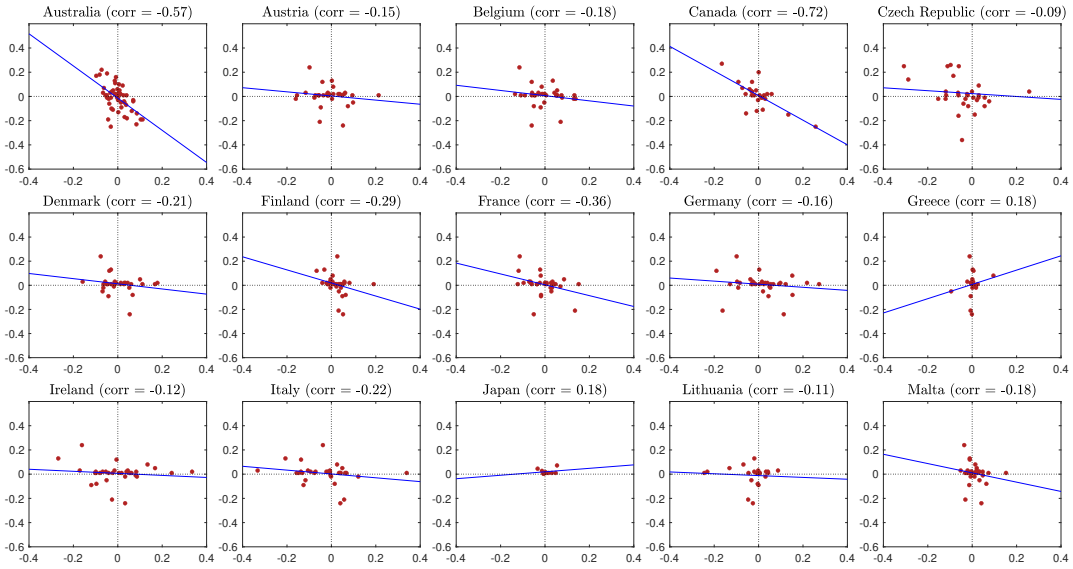
Note: This table presents the estimated response of nominal term premia at 5-, 7-, and 10-year maturities to a 1 percentage point surprise monetary policy tightening, for emerging markets and advanced economies. The nominal term premium is defined as the excess yield investors require to hold sovereign long-term nominal bonds relative to short-term ones. All specifications include country fixed effects and six lags of the dependent variable and the policy rate surprise. The sample period begins in 2010. Significance stars indicate: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. See Appendix A.1 for further details on data and methodology.

Table A.8: Response of Term Premia to Surprise Tightenings (alternative sample)

	Emerging Markets			Advanced Economies		
	5y	*7y*	*10y*	*5y*	*7y*	*10y*
Surprise tightening	0.237^{***}	0.277^{***}	0.245^{*}	-0.502^{***}	-0.615^{***}	-0.764^{***}
(1 p.p.)	(0.083)	(0.102)	(0.130)	(0.102)	(0.112)	(0.117)
Observations	1390	1390	1390	3153	3153	3153
Countries	15	15	15	27	27	27
R^2	0.74	0.75	0.71	0.92	0.94	0.94

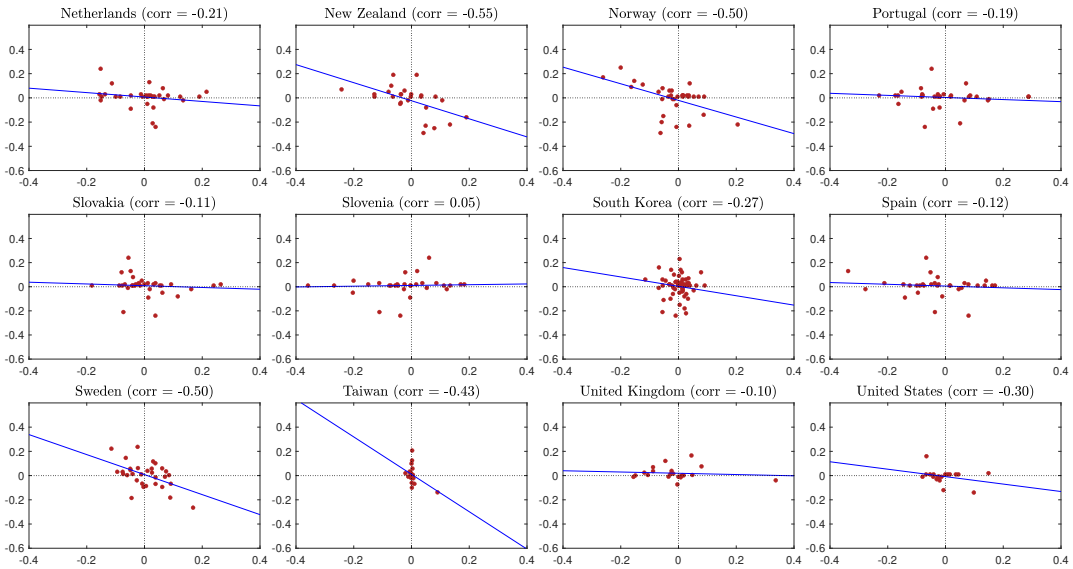
Note: This table presents the estimated response of nominal term premia at 5-, 7-, and 10-year maturities to a 1 percentage point surprise monetary policy tightening, for emerging markets and advanced economies. The nominal term premium is defined as the excess yield investors require to hold sovereign long-term nominal bonds relative to short-term ones. All specifications include country fixed effects and three lags of the dependent variable and the policy rate surprise. The sample period begins in 2010 and ends in 2019. Significance stars indicate: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. See Appendix A.1 for further details on data and methodology.

Figure A.5: Change in 10-Year Nominal Term Premium Around Policy News, by Country



Note: These figures plot the two-day change (in basis points) in the 10-year nominal term premium around monetary policy announcement dates, separately for 15 advanced economies in the sample. Each red dot corresponds to a policy announcement; the solid blue line is the fitted regression for that country. The sample period begins in 2010.

Figure A.6: Change in 10-Year Nominal Term Premium Around Policy News, by Country



Note: These figures plot the two-day change (in basis points) in the 10-year nominal term premium around monetary policy announcement dates, separately for 12 advanced economies in the sample. Each red dot corresponds to a policy announcement; the solid blue line is the fitted regression for that country. The sample period begins in 2010.

Table A.9: Change in 10-Year Nominal Term Premia Around MP Surprises

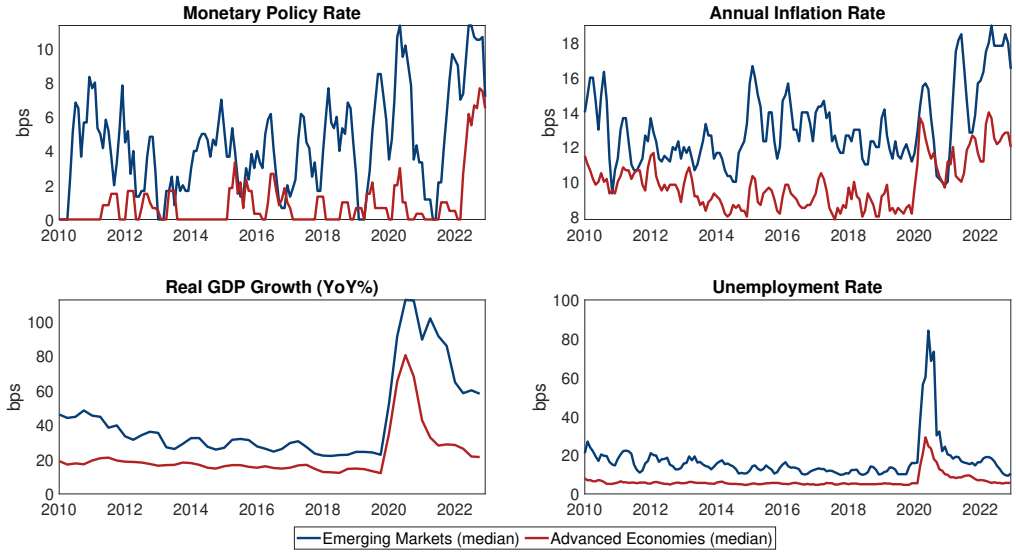
	Emerging Markets		Advanced Economies	
	(1)	(2)	(3)	(4)
Surprise tightening	0.149^{***}	0.154^{***}	-0.213^{***}	-0.211^{***}
(1 p.p.)	(0.032)	(0.032)	(0.035)	(0.035)
Observations	710	710	844	844
Countries	15	15	27	27
R^2	0.10	0.08	0.09	0.06
Country FE	Yes	No	Yes	No

Note: This table presents the results from regressing monetary policy surprises on the two-day change in the 10-year nominal term premium around monetary policy meeting dates (see Figure A.1). The sample is restricted to meetings with non-zero surprises and covers the period from 2010 onward. Significance stars indicate: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. See Appendix A.1 for further details on data and methodology.

A.2 Central Bank Communication in Emerging Markets

Managing private expectations has become a central aspect of monetary policy, as private agents' interpretation of central bank actions and communications plays a key role in shaping their beliefs (Woodford, 2005). In principle, clear communication by the central bank increases the predictability of policy actions and helps to anchor inflation expectations over the long run (Blinder, 2009). While early research largely focused on the impact of communication with financial markets and professional forecasters (Siklos 2002; Fujiwara 2005; Lucca and Trebbi 2009), more recent attention has shifted toward engaging the broader

Figure A.7: Dispersion of Macroeconomic Forecasts (median st. dev. of forecasts)



Note: Each panel presents the median cross-sectional dispersion (standard deviation) of analysts' forecasts across countries within each group (EMs or AEs). For a given month, I compute the standard deviation of forecasts for each country, and then take the median across countries. All series are shown as three-month moving averages for smoothness. See Section 1.5 for further information on the data and methodology.

public.[4] As Blinder et al. (2008) observed, "it may be time to pay some attention to communication with the general public." Haldane and McMahon (2018) presents compelling evidence supporting this shift.[5]

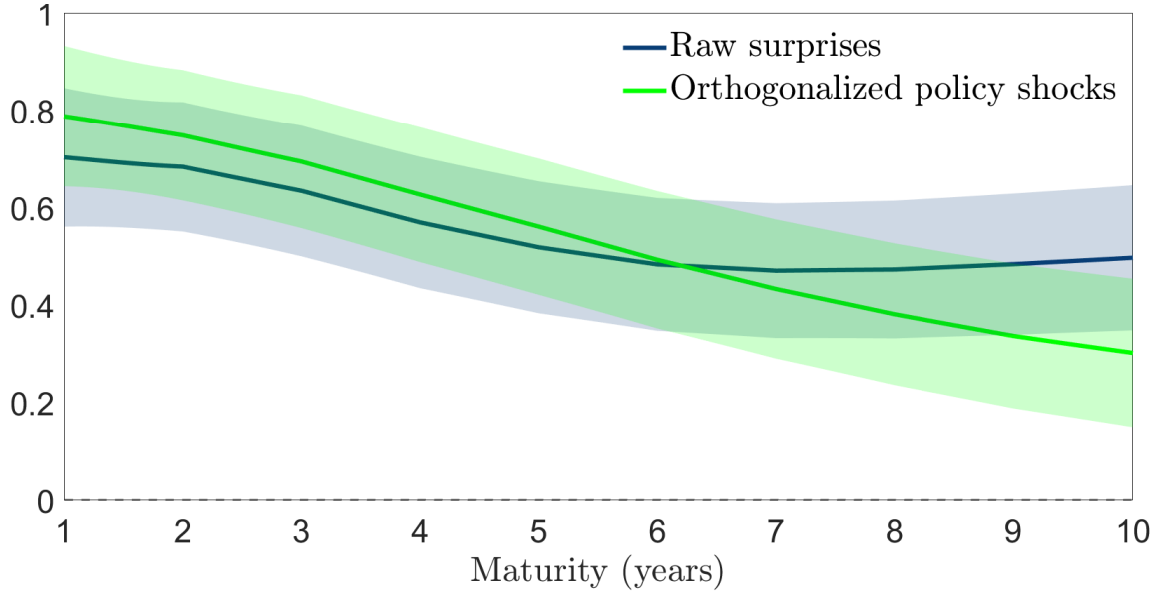
However, effective communication is not trivial (Winkler, 2000). Guiding market expectations requires not only forward-looking statements, but also consistency between words and actions.[6] The empirical literature documents that higher-quality and more transparent communication improves predictability and reduces policy surprises. For example, Tarkka et al. (2000) find that publishing central bank inflation forecasts makes monetary policy more effective. Fracasso et al. (2003) show that countries with clearer and more con-

[4] See also Altavilla and Giannone (2017), Nakamura and Steinsson (2018), and Cieslak and Schrimpf (2019) for further evidence on communication with financial market participants and forecasters.

[5] Related research highlights the importance of broad-based communication strategies (Binder 2017, Coibion et al. 2022).

[6] As Candia et al. (2020) emphasize, careful attention must be paid to communication strategies, as households and firms may sometimes misinterpret information about inflation and output.

Figure A.8: Response of the Yield Curve to a 1 p.p. Tightening, Emerging Markets

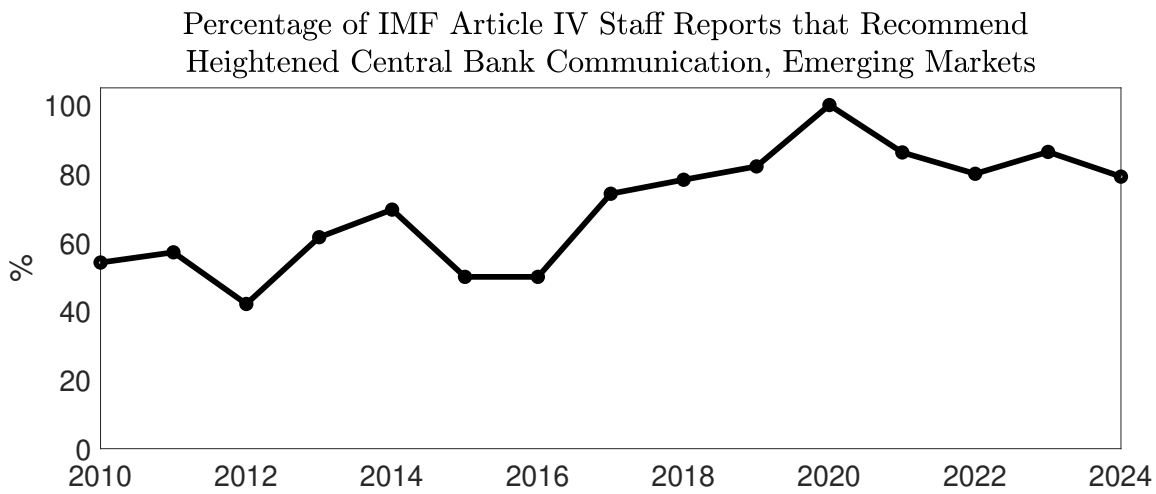


Note: This figure presents the estimated response of zero-coupon yields at different maturities in emerging markets to a 1 percentage point unanticipated increase in the domestic policy rate. The blue line uses raw forecast-based surprises as the instrument, while the green line uses orthogonalized monetary policy shocks that are purged of information effects. Shaded areas denote 90% confidence intervals. The sample is restricted to bonds with up to 10 years to maturity to minimize liquidity premia effects. The sample period begins in 2010. See Appendix A for details on sample construction, methodology, and the list of countries included.

vincing inflation reports experience fewer surprises in interest rate decisions. Meanwhile, Bholat et al. (2019) demonstrate that visual summaries can enhance public comprehension of central bank communications.

To assess the practical emphasis placed on central bank communication in emerging markets, I compile and review 333 IMF Article IV Staff Reports issued for 35 EM economies between 2010 and 2024. In each report, I check whether IMF authorities explicitly recommend that central banks enhance their communication strategies. Two main findings emerge. First, such recommendations are widespread: 234 of the 333 documents (70%) urge improvements in central bank communication. Second, and consistent with the growing recognition in the literature of the importance of clear communication, these recommendations have become more frequent over time. As shown in Figure A.9, the

Figure A.9: Central Bank Communication in Emerging Markets and the IMF



Note: This figure presents the annual percentage of IMF Article IV Staff Reports for emerging markets that include explicit recommendations for heightened central bank communication. It is based on a sample of 333 IMF documents issued for 35 emerging markets.

share of reports recommending heightened communication has trended upward over the past decade, frequently exceeding 80 percent in recent years. This persistent pattern underscores both the recognized importance and the continuing challenges of effective central bank communication in emerging markets.

In sum, both the academic literature and policy practice highlight the centrality of communication for the credibility and effectiveness of monetary policy. Yet, as evidenced by the widespread and rising frequency of IMF recommendations, effective communication remains an ongoing challenge for many emerging market central banks. Persistent calls for greater clarity and outreach reflect not only the recognized benefits of well-anchored expectations, but also the considerable institutional and practical hurdles that remain.

A.3 Model Appendix

A.3.1 Imperfect Information Block

This appendix discusses the results from the imperfect information block in Section 1.4.1.

A.3.2 Arbitrageurs' Expectations Formation

Let's start by putting this problem in state-space form. Let $z_t = \pi_t - \bar{\pi}$ be the deviation of inflation from its unconditional mean. The state equation is thus

$$dz_t = -\kappa_\pi z_t\, dt - \theta\, di_t + \sigma_\pi\, dB_{\pi,t}. \tag{A.26}$$

Using equation (1.7), the state dynamics can be rewritten as

$$dz_t = -(\kappa_\pi + \phi\,\theta)\, z_t\, dt + \sigma_\pi\, dB_{\pi,t} - \theta\,\sigma_i\, dB_{i,t}, \tag{A.27}$$

where the process noise variance is $\text{var}_t[dz_t] = \sigma_\pi^2 + \theta^2\sigma_i^2 \equiv Q$, and the state and measurement noises are correlated with covariance $\text{cov}_t(dz_t, di_t) = -\theta\,\sigma_i^2\, dt$. The two measurement equations, written here again for completeness are:

$$\text{(short rate increment)} \qquad di_t \;=\; \phi\, z_t\, dt \;+\; \sigma_i\, dB_{i,t}; \tag{A.28}$$

$$\text{(aggregate private signal)} \qquad ds_t \;=\; z_t\, dt \;+\; \sigma_s\, dB_{s,t}. \tag{A.29}$$

Denote $dm_t = [di_t \;\; ds_t]'$, I can stack the two measurement equations together as

$$dm_t = H\, z_t dt + R^{1/2}\, dV_t, \tag{A.30}$$

where

$$H = \begin{bmatrix} \phi \\ 1 \end{bmatrix} \qquad R = \begin{bmatrix} \sigma_i^2 & 0 \\ 0 & \sigma_s^2 \end{bmatrix} \qquad dV_t = \begin{bmatrix} dZ_{i,t} \\ dZ_{s,t} \end{bmatrix}. \tag{A.31}$$

Let $\hat{z}_t = \mathbb{E}[z_t]$ denote arbitrageurs' posterior expectation of inflation (relative to its mean),

and $F_t = \text{var}_t[z_t]$ the associated posterior variance. The key step is to express the measurement equations in terms of surprises. Define

$$\text{(policy rate surprise)} \qquad dI_t \;=\; di_t - \phi\,\hat{z}_t dt; \tag{A.32}$$

$$\text{(private signal surprise)} \qquad dS_t \;=\; ds_t - \hat{z}_t dt. \tag{A.33}$$

By construction, these satisfy $\mathbb{E}t[dI_t] = \mathbb{E}t[dS_t] = 0$, with variances $\text{var}_t[dI_t] = \sigma_i^2\,dt$ and $\text{var}_t[dS_t] = \sigma_s^2\,dt$. They are also mutually uncorrelated.

The Kalman-Bucy filter implies that the posterior mean evolves as

$$d\hat{z}_t = -(\kappa_\pi + \phi\,\theta)\,\hat{z}_t\,dt + K_i\,dI_t + K_s\,dS_t, \tag{A.34}$$

where, taking into account the correlation between state and measurement noise, the Kalman gains are

$$K_i = F_t\frac{\phi}{\sigma_i^2} - \theta, \qquad K_s = F_t\,\frac{1}{\sigma_s^2}. \tag{A.35}$$

In terms of inflation expectations, recall that $\hat{\pi}_t = \hat{z}_t + \bar{\pi}$, so the law of motion can be written as equation (1.11).

The posterior variance F_t evolves according to the Riccati equation

$$\dot{F}_t = -2\kappa_\pi F_t + \sigma_\pi^2 - \Lambda F_t^2, \tag{A.36}$$

where

$$H'R^{-1}H = \frac{\phi^2}{\sigma_i^2} + \frac{1}{\sigma_s^2} \equiv \Lambda. \tag{A.37}$$

In steady state ($\dot{F}_t = 0$), F_t converges to the positive solution F, yielding

$$F = \frac{-\kappa_\pi + \sqrt{\kappa_\pi^2 + \Lambda\sigma_\pi^2}}{\Lambda}. \tag{A.38}$$

A.3.3 Proof of Lemma 1

Proof of Lemma 1. The Kalman gains, $K = [K_i \;\; K_s]$, are obtained from

$$K = (FH' + S)R^{-1}, \tag{A.39}$$

where H and R are defined as in (A.31). The covariance between the state noise and the measurement noise is $S = [-\theta\,\sigma_i^2 \;\; 0]$. Simple matrix multiplication then yields

$$K_i = \frac{\phi F}{\sigma_i^2} - \theta, \qquad K_s = F\,\frac{1}{\sigma_s^2}. \tag{A.40}$$

Plugging equation (A.38) into (A.40) yields equations (1.14) and (1.15). □

A.3.4 Proof of Proposition 1

Proof of Proposition 1. The proof is simple and follows from computing the derivative of equation (1.11) with respect to dI_t. □

A.3.5 Portfolio-Balance Framework

This appendix discusses the results from the portfolio-balance block in Section 1.4.1. For further details and extensions, see Greenwood and Vayanos (2014a).

A.3.6 Proof of Lemma 2

Proof of Lemma 2. Recall that, to pin down equilibrium bond prices $P_t^{(\tau)}$, we conjecture they are affine in the state variables:

$$P_t^{(\tau)} = e^{-[A_i(\tau)i_t + A_\pi(\tau)\hat{\pi}_t + C(\tau)]}. \tag{A.41}$$

Using Itô's Lemma, the instantaneous return on a bond with maturity τ is

$$\begin{aligned}\frac{dP_t^{(\tau)}}{P_t^{(\tau)}} = &\left[A_i'(\tau)i_t + A_\pi'(\tau)\hat{\pi}_t + C'(\tau) - A_i(\tau)\phi(\hat{\pi}_t - \bar{\pi}) + A_\pi(\tau)(\kappa_\pi + \theta\,\phi)(\hat{\pi}_t - \bar{\pi})\right]dt \\ &+ \left[A_i(\tau)A_\pi(\tau)K_i\sigma_i^2\right]dt + \frac{1}{2}\left[A_i(\tau)^2\sigma_i^2 + A_\pi(\tau)^2\left(K_i^2\sigma_i^2 + K_s^2\sigma_s^2\right)\right]dt \\ &- A_i(\tau)\,dI_t - A_\pi(\tau)(K_i\,dI_t + K_s\,dS_t). \end{aligned} \tag{A.42}$$

where

$$\begin{aligned}\mu_t^{(\tau)} = &A_i'(\tau)i_t + A_\pi'(\tau)\hat{\pi}_t + C'(\tau) - A_i(\tau)\phi(\hat{\pi}_t - \bar{\pi}) + A_\pi(\tau)(\kappa_\pi + \theta\,\phi)(\hat{\pi}_t - \bar{\pi}) \\ &+ A_i(\tau)A_\pi(\tau)K_i\sigma_i^2 + \frac{1}{2}\left[A_i(\tau)^2\sigma_i^2 + A_\pi(\tau)^2\left(K_i\sigma_i^2 + K_s^2\sigma_s^2\right)\right], \end{aligned} \tag{A.43}$$

Substituting (A.42) into the arbitrageur's budget constraint (equation (1.21)) yields

$$\begin{aligned}\frac{dW_t}{W_t} = &(i_t - \hat{\pi}_t)dt + \left[\int_0^T x_t^{(\tau)}(\mu_t^{(\tau)} - i_t)d\tau\right]dt \\ &- \left[\int_0^T x_t^{(\tau)}\left[A_i(\tau) + A_\pi(\tau)K_i\right]d\tau\right]dI_t - \left[\int_0^T x_t^{(\tau)}A_\pi(\tau)K_s d\tau\right]dS_t. \end{aligned} \tag{A.44}$$

The mean and variance of the instantaneous change in real wealth are therefore

$$\mathbb{E}_t\left(\frac{dW_t}{W_t}\right) = i_t - \pi_t + \int_0^T x_t^{(\tau)}(\mu_t^{(\tau)} - i_t)d\tau; \tag{A.45}$$

$$\mathrm{var}_t\left(\frac{dW_t}{W_t}\right) = \sigma_i^2\left[\int_0^T x_t^{(\tau)}[A_i(\tau) + A_\pi(\tau)K_i]d\tau\right]^2 + K_s^2\sigma_s^2\left[\int_0^T x_t^{(\tau)}A_\pi(\tau)d\tau\right]^2. \tag{A.46}$$

Pointwise maximization of the mean-variance objective in equation (1.22), using (A.45) and (A.46), yields the first-order condition in (1.28). □

A.3.7 Proof of Theorem 1

Proof of Theorem 1. To show these results, I first present the following lemma, which is identical to Greenwood and Vayanos (2014a, Lemma A.1). Please refer to their paper for the proof, which uses Assumption 1.

Lemma 1. *If a function $g(\tau)$ is positive and increasing, then $\int_0^T g(\tau)\alpha(\tau) > 0$.*

Using the arbitrageurs' first-order condition (equation (1.28)), along with the instantaneous expected return (equation (1.29)), the market prices of risk (equations (1.30) and (1.31)), and imposing market clearing (equations (1.23) and (1.26)), obtains

$$\begin{aligned}
&A_i'(\tau)i_t + A_\pi'(\tau)\hat{\pi}_t + C'(\tau) - A_i(\tau)\phi(\hat{\pi}_t - \bar{\pi}) + A_\pi(\tau)(\kappa_\pi + \theta\phi)(\hat{\pi}_t - \bar{\pi}) \\
&+ \frac{1}{2}A_i(\tau)^2\sigma_i^2 + \frac{1}{2}A_\pi(\tau)^2\left(K_i^2\sigma_i^2 + K_s^2\sigma_s^2\right) + K_i\sigma_i^2 A_i(\tau)A_\pi(\tau) - i_t \\
&= A_i(\tau)\gamma\sigma_i^2\left[\int_0^T A_i(\tau)\left[s_t^{(\tau)} - \left(\bar{d}_t^{(\tau)} - \alpha(\tau)(\hat{\pi}_t - \bar{\pi})\right)\right]d\tau\right] \\
&+ A_i(\tau)\gamma\sigma_i^2 K_i\left[\int_0^T A_\pi(\tau)\left[s_t^{(\tau)} - \left(\bar{d}_t^{(\tau)} - \alpha(\tau)(\hat{\pi}_t - \bar{\pi})\right)\right]d\tau\right] \\
&+ A_\pi\gamma K_s^2\sigma_s^2\left[\int_0^T A_\pi(\tau)\left[s_t^{(\tau)} - \left(\bar{d}_t^{(\tau)} - \alpha(\tau)(\hat{\pi}_t - \bar{\pi})\right)\right]d\tau\right] \\
&+ A_\pi(\tau)\gamma\sigma_i^2 K_i\left[\int_0^T A_i(\tau)\left[s_t^{(\tau)} - \left(\bar{d}_t^{(\tau)} - \alpha(\tau)(\hat{\pi}_t - \bar{\pi})\right)\right]d\tau\right] \\
&+ A_\pi(\tau)\gamma\sigma_i^2 K_i^2\left[\int_0^T A_\pi(\tau)\left[s_t^{(\tau)} - \left(\bar{d}_t^{(\tau)} - \alpha(\tau)(\hat{\pi}_t - \bar{\pi})\right)\right]d\tau\right].
\end{aligned} \tag{A.47}$$

Identifying terms with i_t

$$A_i'(\tau) = 1; \tag{A.48}$$

identifying terms with $\hat{\pi}_t$

$$A_\pi'(\tau) + \kappa_\pi^\star A_\pi(\tau) = A_i(\tau)\left[\phi + \gamma\sigma_i^2\left(\int_0^T A_i(\tau)\alpha(\tau)d\tau + K_i\int_0^T A_\pi(\tau)\alpha(\tau)d\tau\right)\right], \tag{A.49}$$

where

$$\kappa_\pi^\star = \kappa_\pi + \phi\,\theta - \gamma\left[K_s^2\sigma_s^2\int_0^T A_\pi(\tau)\alpha(\tau)d\tau + \sigma_i^2 K_i\left(\int_0^T A_i(\tau)\alpha(\tau)d\tau + K_i\int_0^T A_\pi(\tau)\alpha(\tau)d\tau\right)\right]. \tag{A.50}$$

Lastly, identifying constant terms

$$\begin{aligned}
&C'(\tau) + A_i(\tau)\phi\,\bar{\pi} - A_\pi(\tau)(\kappa_\pi + \phi\,\theta)\,\bar{\pi} + K_i\sigma_i^2 A_i(\tau)A_\pi(\tau) \qquad \text{(A.51)}\\
&\quad + \frac{1}{2}\left[A_i(\tau)^2\sigma_i^2 + A_\pi(\tau)^2\left(K_i^2\sigma_i^2 + K_s^2\sigma_s^2\right)\right]\\
&\quad = A_i(\tau)\gamma\sigma_i^2\left[\int_0^T A_i(\tau)\left[s_t^{(\tau)} - \left(\overline{d}_t^{(\tau)} + \alpha(\tau)\bar{\pi}\right)\right]d\tau\right]\\
&\quad + A_i(\tau)\gamma\sigma_i^2 K_i\left[\int_0^T A_\pi(\tau)\left[s_t^{(\tau)} - \left(\overline{d}_t^{(\tau)} + \alpha(\tau)\bar{\pi}\right)\right]d\tau\right]\\
&\quad + A_\pi\gamma K_s^2\sigma_s^2\left[\int_0^T A_\pi(\tau)\left[s_t^{(\tau)} - \left(\overline{d}_t^{(\tau)} + \alpha(\tau)\bar{\pi}\right)\right]d\tau\right]\\
&\quad + A_\pi(\tau)\gamma\sigma_i^2 K_i\left[\int_0^T A_i(\tau)\left[s_t^{(\tau)} - \left(\overline{d}_t^{(\tau)} + \alpha(\tau)\bar{\pi}\right)\right]d\tau\right]\\
&\quad + A_\pi(\tau)\gamma\sigma_i^2 K_i^2\left[\int_0^T A_\pi(\tau)\left[s_t^{(\tau)} - \left(\overline{d}_t^{(\tau)} + \alpha(\tau)\bar{\pi}\right)\right]d\tau\right]. \qquad \text{(A.52)}
\end{aligned}$$

Equations (A.48) and (A.49) are ordinary differential equations (ODE) that must be solved with the initial condition $A_i(\tau) = A_\pi(\tau) = 0$. Equation (1.32) follows instantly. Plugging (1.32) into (A.49) and solving for $A_\pi(\tau)$ yields (1.33).

Next, I establish that equation (1.35) admits a solution whenever the risk-aversion parameter γ is sufficiently small. Since both $A_i(\tau)$ and $A_\pi(\tau)$ are strictly positive and increasing in maturity, the integrals in (1.35) are strictly positive as well. As a result, the final term on the right-hand side of (1.35) is positive, which in turn implies that $\kappa_\pi^\star$ must lie strictly below $\kappa_\pi + \phi\,,\theta$.

To guarantee existence, it is enough to verify that the right-hand side of (1.35) exceeds

the left-hand side when $\kappa_\pi^\star = 0$. This requirement can be written as

$$\gamma < \frac{\kappa_\pi + \phi, \theta}{\sigma_i^2 K_i \int_0^T \tau\alpha(\tau)d\tau + Z\left(K_s^2\sigma_s^2 + \sigma_i^2 K_i^2\right)\int_0^T \left(\frac{\tau}{\kappa_\pi^\star} + \frac{e^{-\kappa_\pi^\star}-1}{\kappa_\pi^{\star 2}}\right)\alpha(\tau)d\tau}. \tag{A.53}$$

Accordingly, a sufficient condition for existence is that γ falls below a critical threshold $\bar{\gamma} > 0$, defined by the right-hand side of (A.53). Under this condition, equation (1.35) admits at least one solution. □

A.3.8 Proof of Proposition 2

Proof to Proposition 2. First-order derivative of equation (1.36) with respect to the monetary policy rate surprise yields (1.37). □

A.3.9 Proof of Proposition 3

Proof to Proposition 3. From the market clearing condition (equation (1.26)), the residual supply to arbitrageurs changes by

$$\frac{\partial x_t^{(\tau)}}{\partial dI_t} = \alpha(\tau)K_i. \tag{A.54}$$

Define the exposure integrals

$$R_i \equiv \int_0^T x_t^{(\tau)}(A_i(\tau) + A_\pi(\tau)K_i)\,d\tau, \qquad R_\pi \equiv \int_0^T x_t^{(\tau)}A_\pi(\tau)\,d\tau. \tag{A.55}$$

Differentiating with respect to dI_t:

$$\frac{\partial R_i}{\partial dI_t} = K_i\int_0^T \alpha(\tau)(A_i(\tau) + A_\pi(\tau)K_i)\,d\tau, \qquad \frac{\partial R_\pi}{\partial dI_t} = K_i\int_0^T \alpha(\tau)A_\pi(\tau)\,d\tau. \tag{A.56}$$

Because $A_i(\tau)$ and $A_\pi(\tau)$ are positive, and given Assumption 1, then R_i and R_π inherit the sign of K_i. Thus, from equations (1.30) and (1.31)

$$\frac{\partial \lambda_{i,t}}{\partial dI_t} \propto K_i; \qquad \frac{\partial \lambda_{\hat{\pi},t}}{\partial dI_t} \propto K_i, \tag{A.57}$$

which completes the proof. □

A.3.10 Proof of Proposition 4

Proof of Proposition 4. The proof follows from Appendix A.3.9. Following a surprise tightening, market prices of risk move. Forward rates incorporate this change in risk compensation, while expected short rates do not. This yields the overreaction/underreaction results. □

Table A.10: Average Dispersion of Macroeconomic Forecasts, 2010–2019

A. Emerging Markets

	Annual inflation (bps)	Real GDP YoY% (bps)	Unemployment rate (bps)
Brazil	8.2	30.5	11.9
Chile	11.7	13.9	12.0
Colombia	8.3	20.1	34.2
Hungary	14.4	33.7	
India	22.7	27.5	
Indonesia	16.8	14.7	
Malaysia	16.9	42.9	
Mexico	8.5	29.3	12.4
Nigeria	24.2		
Peru			17.7
Philippines	16.3	40.6	
Romania	16.6	48.6	
Russia	16.5	39.3	12.6
South Africa	12.4	25.7	30.8
Thailand	16.1	62.3	
Median	**16.2**	**30.5**	**12.6**

B. Advanced Economies

	Annual inflation (bps)	Real GDP YoY% (bps)	Unemployment rate (bps)
Australia	10.9	18.4	6.3
Canada	10.8	10.1	5.7
Czech Republic	8.4	27.0	5.9
Euro Area	8.8	10.3	4.6
Japan	9.4	17.8	5.4
New Zealand	10.5	24.4	8.1
Norway	11.0	16.0	10.8
South Korea	10.2	14.8	6.0
Sweden	8.5	19.9	8.3
Taiwan	10.3	49.4	2.9
United Kingdom	9.5	10.5	5.0
United States	8.1	16.9	5.8
Median	**9.9**	**17.3**	**5.9**

Note: This table presents the average cross-sectional standard deviation of analysts' forecasts for annual inflation, real GDP growth, and the unemployment rate over the period 2010–2019. Dispersion is expressed in basis points (bps) for all variables. Months with fewer than five available forecasts are excluded. Data is from *Bloomberg*. See Section 1.5 for further information on the data and methodology.

Appendix B: Appendix to Chapter 2

B.1 Model Appendix

B.1.1 Return of a Perpetuity: Campbell-Shiller Approximation

We use the Campbell and Shiller (1988) log-linear approximation to model the before-tax return on default-free long-term bonds. We assume that agents take as given the exogenous tax described in the model section. These bonds are self-amortizing perpetuities whose payments decline geometrically, are free of default risk, and have a face value of 1 at time t. Let P_t^y be the price and Y_t the yield-to-maturity of these long-term bonds at time t. At $t+1$, this instrument will offer a coupon payment of C, a principal repayment of $1-\kappa$ for some $\kappa \in [0,1]$, and κ units of the asset. Here, κ is the continuation rate. The gross before-tax return on long-term bonds from t to $t+1$ is thus

$$1+R_{t+1}^y = \frac{C+1-\kappa+\kappa P_{t+1}^y}{P_t^y}, \tag{B.1}$$

where

$$P_t^y = \sum_{j=1}^{\infty} \frac{\kappa^{j-1}(1-\kappa+C)}{(1+Y_t)^j} = \frac{1+C-\kappa}{1+Y_t-\kappa}. \tag{B.2}$$

Taking a log-linear approximation of the long-term bond's return around the point where it is trading at par at $t+1$ obtains

$$r^y_{t+1} \approx \theta + \delta p^y_{t+1} - p^y_t, \tag{B.3}$$

where $\theta \equiv \ln(1+C)$ and $\delta \equiv \kappa/(1+C)$ are parameters. We can iterate this equation forward and apply this approximation to Y_t to get

$$p^y_t \approx \frac{1}{1-\delta}\theta - \frac{1}{1-\delta}y_t. \tag{B.4}$$

We plug equation (B.4) into (B.3) to get the approximate one-period log return on the long-term bond

$$r^y_{t+1} \approx \frac{1}{1-\delta}y_t - \frac{\delta}{1-\delta}y_{t+1}, \tag{B.5}$$

where $D = (1-\delta)^{-1}$ is the Macaulay duration when the instrument is trading at par.

Lastly, one can subtract the tax g_t from this return to get expression (2.3).

B.1.2 Spot and forward exchange markets and Covered Interest Parity

Let d^x_t denote the intermediary's net positions in the borrow-foreign and lend-domestic FX trade in the spot market. Let d^f_t denote a long-domestic forward position of the intermediary. Both d^x_t and d^f_t are expressed in domestic currency units. The realized returns from these positions are given by:

$$d^x_t[i_t - i^\star_t - (q_{t+1} - q_t)] + d^f_t[f_t - q_{t+1}] \tag{B.6}$$

where f_t is the time-t forward log nominal exchange rate, measured in units of domestic currency per unit of foreign currency. Defining the total FX market position as $d^q_t \equiv d^x_t + d^f_t$,

we rewrite the expression above as:

$$d_t^q[i_t - i_t^\star - (q_{t+1} - q_t)] - d_t^f[i_t - i_t^\star - (f_t - q_t)] \tag{B.7}$$

The first term is the realized return of the intermediary's net exposure to the UIP trade, combining long spot and forward FX positions. The second term is the return from a *covered* FX trade, which earns any covered interest parity (CIP) deviations CIP devs$_t \equiv i_t - i_t^\star - (f_t - q_t)$.

Defining $\mathbf{d}_t \equiv [d_t^y, d_t^q, d_t^f]'$ and $\mathbf{rx}_{t+1} \equiv [rx_{t+1}^y, rx_{t+1}^q, \text{CIP devs}_t]'$, intermediaries solve:

$$\max_{\mathbf{d}_t} \left\{ \mathbf{d}_t' \, \mathrm{E}_t[\mathbf{rx}_{t+1}] - \frac{1}{2\tau} \mathbf{d}_t' \, \mathrm{var}_t[\mathbf{rx}_{t+1}] \mathbf{d}_t \right\}, \tag{B.8}$$

Taking first-order condition yields the optimality condition faced by intermediaries:

$$\mathrm{E}_t[\mathbf{rx}_{t+1}] = \tau^{-1} \, \mathrm{var}_t[\mathbf{rx}_{t+1}] \mathbf{d}_t. \tag{B.9}$$

or:

$$\begin{bmatrix} \mathrm{E}_t[rx_{t+1}^y] \\ \mathrm{E}_t[rx_{t+1}^q] \\ \mathrm{E}_t[\text{CIP devs}_t] \end{bmatrix} = \tau^{-1} \begin{bmatrix} \mathrm{var}_t(rx_{t+1}^y) & \mathrm{cov}_t(rx_{t+1}^y, rx_{t+1}^q) & 0 \\ \mathrm{cov}_t(rx_{t+1}^q, rx_{t+1}^y) & \mathrm{var}_t(rx_{t+1}^q) & 0 \\ 0 & 0 & 0 \end{bmatrix} \begin{bmatrix} d_t^y \\ d_t^q \\ d_t^f \end{bmatrix}$$

and thus:

$$\text{CIP devs}_t = 0$$

CIP holds in this model, because any CIP deviations would create *riskless* arbitrage opportunities. Regardless of their risk tolerance, intermediaries would thus eliminate away any CIP deviations.

As a result, we can simplify the intermediary's optimization problem by recasting it in terms of their positions in the long-term domestic bond market (d_t^y) and in the combined

FX market ($d_t^q \equiv d_t^x + d_t^f$), which consolidates spot and forward FX exposures.

B.1.3 Solution of baseline model

In this subsection, we derive the system of equations necessary to solve the baseline model presented in Section 2.3.1. We follow Greenwood et al. (2023a) by conjecturing that equilibrium prices are a linear function of a state vector of shocks and arrive at a system of two equations with two unknowns. The resulting equations can be studied to derive qualitative implications about the model. We close out Appendix B.1 by proving the main propositions of the paper.

Equilibrium Conjecture and Properties

Equilibrium Conjecture We conjecture that the two prices that we need to pin down in equilibrium, y_t and q_t, are a linear function of a state vector of z_t

$$y_t = \alpha_0^y + \alpha_1^{y'} z_t;$$
$$q_t = \alpha_0^q + \alpha_1^{q'} z_t,$$

where the 6×1 state vector $\mathbf{z}_t = [i_t - \bar{\imath}, i_t^* - \bar{\imath}, g_t, \bar{q}_t, s_t^y, s_t^q]'$ follows a VAR(1) process $\mathbf{z}_{t+1} = \Phi \mathbf{z}_t + \varepsilon_{t+1}$, with $\text{var}_t[\varepsilon_{t+1}] = \Sigma$ and $\Phi = \text{diag}(\phi_i, \phi_i, \phi_g, 1, \phi_{s^y}, \phi_{s^q})$. In vector form the two prices yield $\mathbf{y}_t + \mathbf{a} + \mathbf{A}\mathbf{z}_t$, where $\mathbf{y}_t = [y_t, q_t]'$, $\mathbf{a} = [\alpha_0^y, \alpha_0^q]'$, and $\mathbf{A} = [\alpha_1^y, \alpha_1^q]'$.

Rational Expectations Equilibrium Let $f(\text{ff}_0)$ be an operator that gives the price-impact coefficients that clear the markets for long-term bonds and FX when agents conjecture that $\text{ff} = \text{ff}_0$, where $\text{ff} = vec(\text{A})$. We say that a rational expectations equilibrium in this model is a fixed point

$$\alpha^* = f(\alpha^*). \tag{B.11}$$

Within this context, intermediaries must form beliefs—specifically, price-impact coefficients—about how the net supplies of long-term bonds and foreign exchange, s_t^y and s_t^q, influence equilibrium asset prices, y_t and q_t. A rational expectations equilibrium is therefore a fixed point of a specific operator involving these price-impact coefficients.

Equilibrium Properties The presence of supply risk in this model makes the risk tolerance of investors τ a key variable. If agents are not risk-tolerant enough, then an equilibrium does not exist. However, for high levels of risk tolerance, stochastic supply shocks generate multiple equilibria. The different equilibria correspond to different self-fulfilling beliefs (price-impact coefficients) that investors might have about the risk of holding the different assets. For example, if investors believe that supply shocks barely affect prices, they will perceive these assets as less risky. Consequently, investors will absorb large supply shocks and will not require large compensations through a fall in asset prices. However, if investors believe supply shocks will have greater impact on prices, they demand a large decline in prices as compensation for absorbing these shocks.

Despite multiple equilibria, there is always a unique stable equilibrium. Denoting by $\{\lambda_i\}$ the eigenvalues of the Jacobian $D_\alpha f(\alpha^*)$, we see that α^* is stable if $|\lambda_i| < 1$. The importance of a unique stable equilibrium lies in the fact that we can infer qualitative implications from our model. It also implies that comparative statics on our equilibrium price-impact coefficients α^* offer an easy and informative interpretation.

Equilibrium Solution

Recall the first-order condition of intermediaries:

$$\mathrm{E}_t[\mathbf{rx}_{t+1}] = \tau^{-1} \operatorname{var}_t[\mathbf{rx}_{t+1}]\mathbf{d}_t. \tag{B.12}$$

This equation, along with the market clearing condition ($\mathbf{d}_t = \mathbf{s}_t$), implies

$$\mathrm{E}_t[\mathbf{rx}_{t+1}] = \tau^{-1}\mathbf{V}\mathbf{s}_t, \tag{B.13}$$

where

$$\mathbf{V} = \mathrm{var}_t[\mathbf{rx}_{t+1}] = \begin{bmatrix} V_y & C_{y,q} \\ C_{y,q} & V_q \end{bmatrix}.$$

We can write out equation (B.12), yielding individual excess return equations:

$$\mathrm{E}_t\left[rx^y_{t+1}\right] = \frac{1}{\tau}\left[V_y \cdot s^y_t + C_{y,q} \cdot s^q_t\right]; \tag{B.14a}$$

$$\mathrm{E}_t\left[rx^q_{t+1}\right] = \frac{1}{\tau}\left[C_{y,q} \cdot s^y_t + V_q \cdot s^q_t\right], \tag{B.14b}$$

where $V_y \equiv \mathrm{var}_t[rx^y_{t+1}]$, $V_q \equiv \mathrm{var}_t[rx^q_{t+1}]$, and $C_{y,q} \equiv \mathrm{cov}_t[rx^y_{t+1}, rx^q_{t+1}]$. Note that $\mathbf{V}$ is constant in equilibrium and we hereafter drop the time subscripts.

Using these equilibrium excess return equations, along with asset prices equations (2.5) and (3.11), as well as the exogenous processes, we can characterize equilibrium bond yields and foreign exchange prices:

$$y_t = \left\{\bar{\imath} + \frac{1-\delta}{1-\delta\phi_i} \cdot (i_t - \bar{\imath})\right\} + \frac{1-\delta}{1-\delta\phi_g} g_t + \tau^{-1}\left\{\frac{1-\delta}{1-\delta\phi_{s^y}} V_y \cdot s^y_t + \frac{1-\delta}{1-\delta\phi_{s^q}} C_{y,q} \cdot s^q_t\right\}; \tag{B.15}$$

$$q_t = \left\{\frac{1}{1-\phi_i} \cdot (i^\star_t - i_t)\right\} + \bar{q}_t + \tau^{-1}\left\{\frac{1}{1-\phi_{s^y}} C_{y,q} \cdot s^y_t + \frac{1}{1-\phi_{s^q}} V_q \cdot s^q_t\right\}. \tag{B.16}$$

We now focus on the fixed-point problem. The vector of excess returns is

$$\mathbf{rx}_{t+1} \equiv \begin{bmatrix} rx^y_{t+1} \\ rx^q_{t+1} \end{bmatrix} = \begin{bmatrix} \frac{1}{1-\delta} y_t - \frac{\delta}{1-\delta} y_{t+1} - i_t - g_t \\ i_t - i^\star_t - (q_{t+1} - q_t) \end{bmatrix} = \mathbf{B}_0\mathbf{y}_t + \mathbf{B}_1\mathbf{y}_{t+1} + \mathbf{R}_1\mathbf{z}_t + \mathbf{r}_0.$$

where we have used equations (2.1), (2.2), (2.3), (2.6), and the fact that $rx^y_{t+1} \equiv r^y_{t+1} - i_t$. Additionally,

$$\mathbf{B}_0 = \begin{bmatrix} \frac{1}{1-\delta} & 0 \\ 0 & 1 \end{bmatrix}, \ \mathbf{B}_1 = \begin{bmatrix} -\frac{\delta}{1-\delta} & 0 \\ 0 & -1 \end{bmatrix}, \ \mathbf{R}_1 = \begin{bmatrix} -1 & 0 & -1 & 0 & 0 & 0 \\ 1 & -1 & 0 & 0 & 0 & 0 \end{bmatrix}, \mathbf{r}_0 = \begin{bmatrix} -\bar{\iota} \\ 0 \end{bmatrix}.$$

Note that when conjecturing the equilibrium, we defined $\mathbf{y}_t = \mathbf{a} + \mathbf{A}\mathbf{z}_t$. Iterating this expression one period forward and using that $\mathbf{z}_{t+1} = \Phi \mathbf{z}_t + \varepsilon_{t+1}$, one can obtain that

$$\mathbf{y}_{t+1} = \mathbf{a} + \mathbf{A}\mathbf{z}_{t+1} = \mathbf{a} + \mathbf{A}\Phi\mathbf{z}_t + \mathbf{A}\varepsilon_{t+1}. \tag{B.17}$$

Recall that Φ is a diagonal matrix with the AR(1) coefficients of the 6 different exogenous processes. Going back to the equation for $\mathbf{rx}_{t+1}$ we just derived yields

$$\mathbf{rx}_{t+1} = [\mathbf{B}_0\mathbf{a} + \mathbf{B}_1\mathbf{a} + \mathbf{r}_0] + [\mathbf{B}_0\mathbf{A} + \mathbf{B}_1\mathbf{A}\Phi + \mathbf{R}_1]\mathbf{z}_t + [\mathbf{B}_1\mathbf{A}]\varepsilon_{t+1}, \tag{B.18}$$

which implies that

$$E_t[\mathbf{rx}_{t+1}] = [\mathbf{B}_0\mathbf{a} + \mathbf{B}_1\mathbf{a} + \mathbf{r}_0] + [\mathbf{B}_0\mathbf{A} + \mathbf{B}_1\mathbf{A}\Phi + \mathbf{R}_1]\mathbf{z}_t; \tag{B.19}$$

$$\mathbf{V} \equiv \text{var}_t[\mathbf{rx}_{t+1}] = \mathbf{B}_1\mathbf{A}\Sigma\mathbf{A}'\mathbf{B}_1'. \tag{B.20}$$

Going back to the market-clearing condition in equation (B.13), $\mathbf{s}_t = [s^y_t, s^q_t]'$ can be written as $\mathbf{s}_t = \mathbf{S}_1\mathbf{z}_t$, where

$$\mathbf{S}_1 \equiv \begin{bmatrix} 0 & 0 & 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 0 & 0 & 1 \end{bmatrix}$$

which allows us to write equation (B.13) as

$$[\mathbf{B}_0\mathbf{a} + \mathbf{B}_1\mathbf{a} + \mathbf{r}_0] + [\mathbf{B}_0\mathbf{A} + \mathbf{B}_1\mathbf{A}\Phi + \mathbf{R}_1]\mathbf{z}_t = \tau^{-1}(\mathbf{B}_1\mathbf{A}\Sigma\mathbf{A}'\mathbf{B}_1')(\mathbf{S}_1\mathbf{z}_t). \tag{B.21}$$

Equation (B.21) is the main equation which will be used to solve the fixed-point problem. We first separate the terms that contain $\mathbf{z}_t$ from the terms that do not. For the constant terms we find that

$$(\mathbf{B}_0 + \mathbf{B}_1)\mathbf{a} = -\mathbf{r}_0. \tag{B.22}$$

Based on the structure of $\mathbf{B}_0$ and $\mathbf{B}_1$, the last row of the $\mathbf{B}_0 + \mathbf{B}_1$ only contains zeros. Therefore, the domestic long-term bond yield is pinned down in equilibrium—but the constant for the exchange rate is not.

For the terms containing $\mathbf{z}_t$, note that $\mathbf{B}_0, \mathbf{B}_1$, and Φ are diagonal, and thus it follows that

$$[\mathbf{B}_0 \mathrm{A} + \mathbf{B}_1 \mathrm{A}\Phi] = \mathbf{A} \circ [\mathbf{B}_0 \mathbf{E} + \mathbf{B}_1 \mathbf{E}\Phi], \tag{B.23}$$

where $\circ$ is the element-wise matrix multiplication and $\mathbf{E}$ is a 2×6 matrix of 1s. Thus, we get

$$[\mathbf{B}_0 \mathbf{E} + \mathbf{B}_1 \mathbf{E}\Phi] = \begin{bmatrix} \frac{1-\delta\phi_i}{1-\delta} & \frac{1-\delta\phi_i}{1-\delta} & \frac{1-\delta\phi_g}{1-\delta} & \frac{1}{1-\delta} & \frac{1-\delta\phi_{s^y}}{1-\delta} & \frac{1-\delta\phi_{s^q}}{1-\delta} \\ 1-\phi_i & 1-\phi_i & 1-\phi_g & 1 & 1-\phi_{s^y} & 1-\phi_{s^q} \end{bmatrix}.$$

Using this, the terms containing $\mathbf{z}_t$ have to equate on both sides. That is,

$$[\mathbf{A} \circ (\mathbf{B}_0 \mathbf{E} + \mathbf{B}_1 \mathbf{E}\Phi) + \mathbf{R}_1]\mathbf{z}_t = \tau^{-1}(\mathbf{B}_1 \mathbf{A}\Sigma\mathbf{A}'\mathbf{B}_1')\mathbf{S}_1 \mathbf{z}_t. \tag{B.24}$$

Solving for the $\mathbf{A}$ in the LHS yields

$$\mathbf{A} = [\tau^{-1}\mathbf{B}_1 \mathbf{A}\Sigma\mathbf{A}'\mathbf{B}_1'\mathbf{S}_1 - \mathbf{R}_1] \oslash [\mathbf{B}_0 \mathbf{E} + \mathbf{B}_1 \mathbf{E}\Phi], \tag{B.25}$$

where $\oslash$ is element-wise matrix division. To further characterize the solution to the problem in (B.25), we can partition $\mathbf{z}_t$ as $\mathbf{z}_t = [\mathbf{z}_{1,t}', \mathbf{z}_{2,t}', \mathbf{z}_{3,t}']'$, where $\mathbf{z}_{1,t} = [i_t - \bar{\imath}, i_t^* - \bar{\imath}, g_t]'$, $\mathbf{z}_{2,t} = \bar{q}_t$, and $\mathbf{z}_{3,t} = [s_t^y, s_t^q]'$. Similarly, we partition $\mathbf{A} = [\mathbf{A}_1, \mathbf{A}_2, \mathbf{A}_3]$, where $\mathbf{A}_1$ is the 2×3 matrix of loadings on $\mathbf{z}_{1,t}$, $\mathbf{A}_2$ is the 2×1 matrix of loadings on $\mathbf{z}_{2,t}$, and $\mathbf{A}_3$ is the 2×2

matrix of loadings on $\mathbf{z}_{3,t}$.

For an arbitrary matrix $\mathbf{X}$, denote $\mathbf{X}^{[n-m]}$ for $n < m$ be the submatrix of $\mathbf{X}$ consisting of columns $n, n+1, ..., m-1, m$. Therefore, given the form of $\mathbf{R}_1$ and $\mathbf{S}_1$ ($\mathbf{S}_1^{[1-3]} = \mathbf{0}_{2\times 3}$) we can define submatrix $\mathbf{A}_1$ as

$$\mathbf{A}_1 = -\mathbf{R}_1^{[1-3]} \oslash [\mathbf{B}_0\mathbf{E} + \mathbf{B}_1\mathbf{E}\Phi]^{[1-3]} = \begin{bmatrix} \frac{1-\delta}{1-\delta\phi_i} & 0 & \frac{1-\delta}{1-\delta\phi_g} \\ -\frac{1}{1-\phi_i} & \frac{1}{1-\phi_i} & 0 \end{bmatrix}.$$

This matrix displays the price-impact coefficients of the domestic short-term rate (first column), the foreign short-term rate (second column), and the bond-specific shock (third column), on the domestic long-term yield (first row), and on FX (second row).

For $\mathbf{A}_2$, which contains the FX-specific shock, we can already see from the equilibrium price in equation (B.16) that $\mathbf{A}_2 = [0, 1]'$. In other words, the specific shock on the price of the exchange rate has no impact on the long-term bond price, while affecting the FX rate one-for-one.

Lastly, we now move to the supply shocks; that is, pinning down $\mathbf{A}_3$. Due to the orthogonality of the different shocks, the variance-covariance matrix Σ can be partitioned as

$$\Sigma = \begin{bmatrix} \Sigma_1 & 0_{3\times 1} & 0_{3\times 2} \\ 0_{1\times 3} & \Sigma_2 & 0_{1\times 2} \\ 0_{2\times 3} & 0_{2\times 1} & \Sigma_3 \end{bmatrix} \quad \text{where} \quad \Sigma_1 = \begin{bmatrix} \sigma_i^2 & \rho\sigma_i^2 & 0 \\ \rho\sigma_i^2 & \sigma_i^2 & 0 \\ 0 & 0 & \sigma_g^2 \end{bmatrix}, \quad \Sigma_3 = \begin{bmatrix} \sigma_{s^y}^2 & \sigma_{\eta^*}\sigma_{\gamma^*} \\ \sigma_{\eta^*}\sigma_{\gamma^*} & \sigma_{s^q}^2 \end{bmatrix},$$

and $\Sigma_2 = \sigma_q^2$. The variance-covariance matrix of excess returns ($\mathbf{V} \equiv \text{var}_t[\mathbf{rx}_{t+1}]$) becomes

$$\mathbf{V} = (\mathbf{B}_1\mathbf{A}_1\Sigma_1\mathbf{A}_1'\mathbf{B}_1') + (\mathbf{B}_1\mathbf{A}_2\Sigma_2\mathbf{A}_2'\mathbf{B}_1') + (\mathbf{B}_1\mathbf{A}_3\Sigma_3\mathbf{A}_3'\mathbf{B}_1'). \tag{B.26}$$

Making use of the form $\mathbf{S}_1$ and $\mathbf{R}_1$ ($\mathbf{R}_1^{[5-6]} = \mathbf{0}_{2\times 2}$), the following fixed-point problem

involving $\mathbf{A}_3$ is obtained

$$\mathbf{A}_3 = \mathbf{F}_3(\mathbf{A}_3) \equiv \tau^{-1}[(\mathbf{B}_1\mathbf{A}_1\Sigma_1\mathbf{A}_1'\mathbf{B}_1') + (\mathbf{B}_1\mathbf{A}_2\Sigma_2\mathbf{A}_2'\mathbf{B}_1') + (\mathbf{B}_1\mathbf{A}_3\Sigma_3\mathbf{A}_3'\mathbf{B}_1')] \oslash [\mathbf{B}_0\mathbf{E} + \mathbf{B}_1\mathbf{E}\Phi]^{[5-6]}. \tag{B.27}$$

The operator $\mathbf{F}_3(\mathbf{A}_3)$ gives the price function $\mathbf{y}_t = \mathrm{g}(\mathbf{A}_3) + \mathbf{A}_1\mathbf{z}_{1,t} + \mathbf{A}_2\mathbf{z}_{2,t} + \mathbf{F}_3(\mathbf{A}_3)\mathbf{z}_{3,t}$ that will clear the markets for long-term bonds and FX when agents conjecture that the risk of holding of assets is determined by the price function $\mathbf{y}_{t+1} = \mathbf{a}_0 + \mathbf{A}_1\mathbf{z}_{1,t+1} + \mathbf{A}_2\mathbf{z}_{2,t+1} + \mathbf{A}_3\mathbf{z}_{3,t+1}$.

From equation (B.27), and using $\mathbf{V}$ in its matrix form, the equilibrium price-impact coefficients must satisfy

$$\begin{bmatrix} \alpha_{s^y}^{y} & \alpha_{s^q}^{y} \\ \alpha_{s^y}^{q} & \alpha_{s^q}^{q} \end{bmatrix} = \tau^{-1} \begin{bmatrix} \frac{1-\delta}{1-\delta\phi_{s^y}} V_y & \frac{1-\delta}{1-\delta\phi_{s^q}} C_{y,q} \\ \frac{1}{1-\phi_{s^y}} C_{q,y} & \frac{1}{1-\phi_{s^q}} V_q \end{bmatrix}. \tag{B.28}$$

The var.-cov. matrix in the absence of supply risk is $[(\mathbf{B}_1\mathbf{A}_1\Sigma_1\mathbf{A}_1'\mathbf{B}_1') + (\mathbf{B}_1\mathbf{A}_2\Sigma_2\mathbf{A}_2'\mathbf{B}_1')] =$

$$\begin{bmatrix} \left(\frac{\delta}{1-\delta\phi_i}\right)^2 \sigma_i^2 + \left(\frac{\delta}{1-\delta\phi_g}\right)^2 \sigma_g^2 & -\frac{\delta}{1-\delta\phi_i}\frac{1}{1-\phi_i}\sigma_i^2(1-\rho) \\ -\frac{\delta}{1-\delta\phi_i}\frac{1}{1-\phi_i}\sigma_i^2(1-\rho) & \left(\frac{1}{1-\phi_i}\right)^2 2\sigma_i^2(1-\rho) + \sigma_q^2 \end{bmatrix}. \tag{B.29}$$

Solving for the contribution of supply risk to the variance-covariance matrix, one can additionally find $\mathbf{B}_1\mathbf{A}_3\Sigma_3\mathbf{A}_3'\mathbf{B}_1'$. Note that one can recast the fixed-point problem in terms of the variance-covariance matrix, instead of using the 2×2 matrix $\mathbf{A}_3$. This is convenient because $\mathbf{V}$ is symmetric, effectively reducing the fixed-point problem to one involving three unknowns instead of four. One needs to find a fixed point in the form $\mathbf{V} = \mathbf{G}(\mathbf{V})$. Plugging the α's of equation (B.28) in the contribution of supply risk to the variance-covariance

matrix, and using this and (B.29) in (B.26), along with denoting the constants

$$g_y \equiv \tau^{-1}\frac{\delta}{1-\delta\phi_{s^y}}\sigma_\eta,\;\; g_y^* \equiv \tau^{-1}\frac{\delta}{1-\delta\phi_{s^y}}\sigma_{\eta^*},\;\; g_q \equiv \tau^{-1}\frac{\delta}{1-\delta\phi_{s^q}}\sigma_\gamma,\;\; g_q^* \equiv \tau^{-1}\frac{\delta}{1-\delta\phi_{s^q}}\sigma_{\gamma^*},$$

$$h_y \equiv \tau^{-1}\frac{1}{1-\phi_{s^y}}\sigma_\eta,\;\; h_y^* \equiv \tau^{-1}\frac{1}{1-\phi_{s^y}}\sigma_{\eta^*},\;\; h_q \equiv \tau^{-1}\frac{1}{1-\phi_{s^q}}\sigma_\gamma,\;\; h_q^* \equiv \tau^{-1}\frac{1}{1-\phi_{s^q}}\sigma_{\gamma^*}. \tag{B.30}$$

we get that $\mathbf{V}$ must satisfy the following system of three equations in three unknowns:

$$V_y = \left(\frac{\delta}{1-\delta\phi_i}\right)^2\sigma_i^2 + \left(\frac{\delta}{1-\delta\phi_g}\right)^2\sigma_g^2 + (V_y)^2(g_y^2+g_{y^*}^2) + (C_{y,q})^2(g_q^2+g_{q^*}^2) + 2g_{y^*}g_{q^*}V_yC_{y,q}; \tag{B.31a}$$

$$V_q = \left(\frac{1}{1-\phi_i}\right)^2 2\sigma_i^2(1-\rho) + \sigma_q^2 + (C_{y,q})^2(h_y^2+h_{y^*}^2) + (V_q)^2(h_q^2+h_{q^*}^2) + 2h_{y^*}h_{q^*}V_qC_{y,q}; \tag{B.31b}$$

$$\begin{aligned} C_{y,q} = &-\frac{\delta}{1-\delta\phi_i}\frac{1}{1-\phi_i}\sigma_i^2(1-\rho) + V_yC_{y,q}(g_yh_y+g_{y^*}h_{y^*}) + V_qC_{y,q}(g_qh_q+g_{q^*}h_{q^*}) \\ &+ (C_{y,q})^2g_{q^*}h_{y^*} + V_yV_qg_{y^*}h_{q^*}. \end{aligned} \tag{B.31c}$$

where $C_{y,q} = C_{q,y}$ and

$$\mathbf{V} = \mathrm{var}_t[\mathbf{rx}_{t+1}] = \begin{bmatrix} V_y & C_{y,q} \\ C_{y,q} & V_q \end{bmatrix}.$$

which completes the full write-down of the solution method. One must now combine these three equations to find qualitative properties of the three unknowns. Instead, in the next subsection we prove Proposition 3, which effectively reduces this system of equations into a more manageable set that yields qualitatively similar implications.

B.1.4 Proofs

Proof of Proposition 1. Consider a purchase of local-currency denominated long-term bonds by foreign investors, *i.e.*, a decline in $\eta^\star$. From equation (2.14), foreign demand for home bonds is positively associated with foreign demand for the home currency. Hence, a fall in $\eta^\star$ implies a simultaneous fall in $\gamma^\star$, corresponding to a purchase of domestic currency by foreign investors. Using equations (2.11) and (2.12), the joint decline in $\eta^\star$ and $\gamma^\star$ reduces the amount of bond and currency risk intermediaries must absorb in equilibrium. Finally, the equilibrium pricing conditions (B.15) and (B.16) imply that lower intermediary risk exposure compresses the required risk premia. As a result, the local-currency denominated long-term bond yield declines and the home currency appreciates. □

Proof of Proposition 2. Let's begin by observing that

$$\mathbb{E}(\Delta d_t^y \, rx_{t+1}^y) = \mathbb{E}(d_t^y \, rx_{t+1}^y) - \mathbb{E}(d_{t-1}^y rx_{t+1}^y). \tag{B.32}$$

We can use the law of interated expectations on the second term in the right-hand side of (B.32) to obtain:

$$\mathbb{E}(d_{t-1}^y rx_{t+1}^y) = \mathbb{E}\left(\mathrm{E}_t(d_{t-1}^y rx_{t+1}^y)\right) = \mathbb{E}(d_{t-1}^y \, \mathrm{E}_t \, rx_{t+1}^y), \tag{B.33}$$

Using the intermediaries' optimality condition, (2.10), this can be written as:

$$\mathbb{E}(d_{t-1}^y rx_{t+1}^y) = \tau^{-1} \operatorname{var}_t(rx_{t+1}^y)\mathbb{E}(d_{t-1}^y d_t^y) + \tau^{-1} \operatorname{cov}_t(rx_{t+1}^y, rx_{t+1}^q)\mathbb{E}(d_{t-1}^y d_t^q). \tag{B.34}$$

Assumption 1 implies that $\phi_{s^y} = \phi_{s^q} = 0$, therefore, by equations (2.11) and (2.17):

$$\mathbb{E}(d_{t-1}^y d_t^y) = \mathbb{E}(d_{t-1}^y d_t^q) = 0. \tag{B.35}$$

This equation, along with (B.32), implies that:

$$\mathbb{E}(\Delta d_t^y rx_{t+1}^y) = \mathbb{E}(d_t^y rx_{t+1}^y). \tag{B.36}$$

Analogously, we can show that $\mathbb{E}(\Delta d_t^q rx_{t+1}^q) = \mathbb{E}(d_t^q rx_{t+1}^q)$. Next, using equation (2.10), one can show that

$$\mathbb{E}(d_t^y rx_{t+1}^y) = \tau^{-1}\left[\text{var}_t(rx_{t+1}^y)\mathbb{E}\left((d_t^y)^2\right) + \text{cov}_t(rx_{t+1}^y, rx_{t+1}^q)\mathbb{E}\left(d_t^q d_t^y\right)\right]; \tag{B.37}$$

$$\mathbb{E}\left(d_t^q rx_{t+1}^q\right) = \tau^{-1}\left[\text{var}_t(rx_{t+1}^q)\mathbb{E}\left((d_t^q)^2\right) + \text{cov}_t(rx_{t+1}^y, rx_{t+1}^q)\mathbb{E}\left(d_t^y d_t^q\right)\right]. \tag{B.38}$$

Using the market clearing condition (2.18), the processes of net supplies, (2.11) and (3.11), the properties of foreign and home investors' flows, (2.14) and (2.15), as well as Proposition 3, one can show that $\mathbb{E}(d_t^y rx_{t+1}^y) \geq 0$ and $\mathbb{E}(d_t^q rx_{t+1}^q) \geq 0$. □

Proof of Proposition 3. Assumption 1 implies that $V_y = V_q$, which simplifies our system of equations (B.31a)–(B.31c) as follows:

$$V_y = \sigma_g^2 + (V_y)^2\tau^{-2}(\sigma_\eta^2 + \sigma_{\eta^*}^2) + (C_{y,q})^2\tau^{-2}(\sigma_\eta^2 + \sigma_{\eta^*}^2) + 2V_yC_{y,q}\tau^{-2}\sigma_{\eta^*}^2; \tag{B.39a}$$

$$C_{y,q} = 2V_yC_{y,q}\tau^{-2}(\sigma_\eta^2 + \sigma_{\eta^*}^2) + (C_{y,q})^2\tau^{-2}\sigma_{\eta^*}^2 + (V_y)^2\tau^{-2}\sigma_{\eta^*}^2. \tag{B.39b}$$

We can use equation (2.19) to rewrite this system as

$$V_y = \sigma_g^2 + \tau^{-2}(\sigma_\eta^2 + \sigma_{\eta^*}^2)\left[(V_y)^2 + (C_{y,q})^2 + 2\Omega V_yC_{y,q}\right]; \tag{B.40a}$$

$$C_{y,q} = \tau^{-2}(\sigma_\eta^2 + \sigma_{\eta^*}^2)\left[2V_yC_{y,q} + (C_{y,q})^2\Omega + (V_y)^2\Omega\right]. \tag{B.40b}$$

First, setting $\Omega = 0$, it can be readily seen from equations (B.40a)-(B.40b) that the only real solution implies $C_{y,q} = 0$. Second, dividing both sides of (B.40b) by V_y and rearranging, we

obtain

$$\frac{C_{y,q}}{V_y}\left(1-2V_y\tau^{-2}(\sigma_\eta^2+\sigma_{\eta^\star}^2)\right)=\Omega\, V_y\, \tau^{-2}(\sigma_\eta^2+\sigma_{\eta^\star}^2)\left(1+\left(\frac{C_{y,q}}{V_y}\right)^2\right). \tag{B.41}$$

Since $V_y > 0$ and $\tau > 0$, we therefore get that

$$\text{sign}(C_{y,q}) = \text{sign}(\Omega), \tag{B.42}$$

as long as agents are risk-tolerant enough. That is, for τ sufficiently large, we get that

$$1-2V_y\tau^{-2}(\sigma_\eta^2+\sigma_{\eta^\star}^2) > 0. \tag{B.43}$$

□

Proof of Proposition 4. Consider a unit decrease in η_t in (2.11), which implies a reduction in bond net supply s_y without a corresponding movement in currency net supply (i.e. $s_q = 0$). By the equilibrium solution of the exchange rate, eq. (B.16) derived in Appendix B.1.3, such impulse causes a currency appreciation if and only if $\text{cov}_t(rx^y_{t+1}, rx^q_{t+1}) > 0$. This proves part (a). The proof of part (b) is analogous, and follows from considering a unit increase in γ_t in (2.17). □

B.2 Foreign investor type and hedging in fixed income

B.2.1 Foreign investor type in Colombia

Our evidence from Colombia suggests that foreigners tend to produce correlated demand in bond and currency markets. Additionally, the extent of currency hedging from

Table B.1: Baseline calibration

Parameter	Interpretation	Value
$1/\tau$	Risk-tolerance	0.03
δ	Maturity of long-term bonds	0.90
ϕ_i	Persistence of short-term rates	0.965
ρ	Correlation of short-term rates	0.69
ϕ_g	Persistence of bond-specific shocks	0
ϕ_{s^y}; ϕ_{s^q}	Persistence of net supply shocks	0.85
$\sigma_\eta; \sigma_\gamma$	Std. dev. of domestic investors demand	0.0248
$\sigma_{\eta^\star}; \sigma_{\gamma^\star}$	Std. dev. of foreign investors demand	0.0233
σ_i	Std. dev. of interest rate shocks	0.0015
σ_g	Std. dev. of bond-specific shocks	0.05
σ_q	Std. dev. of long-term q shocks	0.05
α	Currency hedge ratio	0

foreigners that purchase local-currency government bonds seems negligible. In this section, we explore the type of foreign investor present in Colombia's sovereign bond market.

As shown in Figure B.1a, foreign participation in Colombia's bond market was limited before 2012 but increased significantly thereafter. This growth was partly driven by regulatory reforms implemented between 2010 and 2013, which facilitated investment through local managers and simplified tax reporting for fixed-income securities. Furthermore, a 2012 tax reform reduced income tax on TES for non-resident investors from 33% to 14%. The country's investment-grade credit rating, awarded in 2011 by credit rating agencies such as *S&P*, *Moody's*, and *Fitch Ratings*, and its increased weight in *J.P. Morgan*'s emerging market debt indices in 2014, also played a key role in attracting international capital.

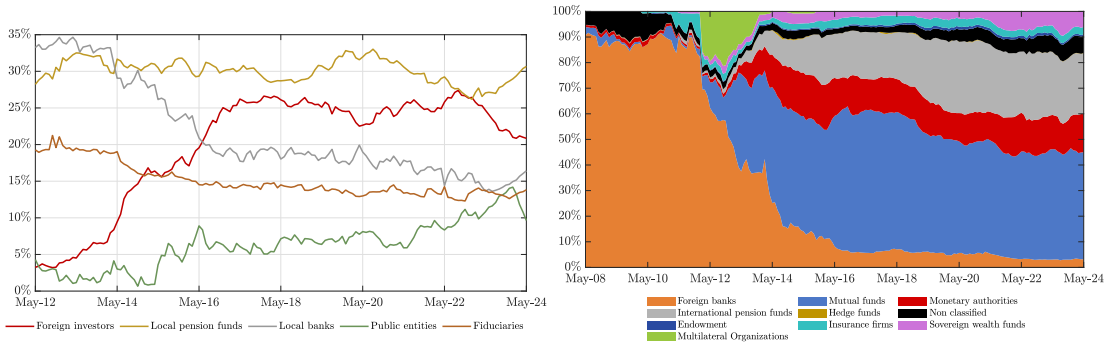
Figure B.1b shows the share of bond holding across foreign investors. Before 2014 foreign banks held the largest share of sovereign bonds among foreign investors, although their overall market participation was relatively low. Since 2014, mutual funds have become the dominant investors, followed by international pension funds and monetary authorities, collectively holding about 80% of the foreign investor base, which is now more significant relative to the total outstanding debt.

B.2.2 Additional Cross-Country Evidence on Comovement and Ω

In Figure B.2 we reproduce Figure 2.5, but replacing $\alpha = 0$ for emerging economies that we estimated have $\alpha < 0$ using the data from Chen and Zhou (2025).

In Figure B.3 we re-compute Ω under two different assumptions. First, we assign the median hedge ratio across countries to all countries (Panel a). Second, we assign the median standard deviation for foreign investors and other domestic investors to all countries (Panel b). In this way, we can keep fixed one of the two reasons why Ω might vary. We observe that both the hedge ratios and the relative volatility of foreign to non-bank domestic investors are relevant to explain the comovement between bond and currency returns across countries.

Figure B.1: Participation in the Colombian TES market

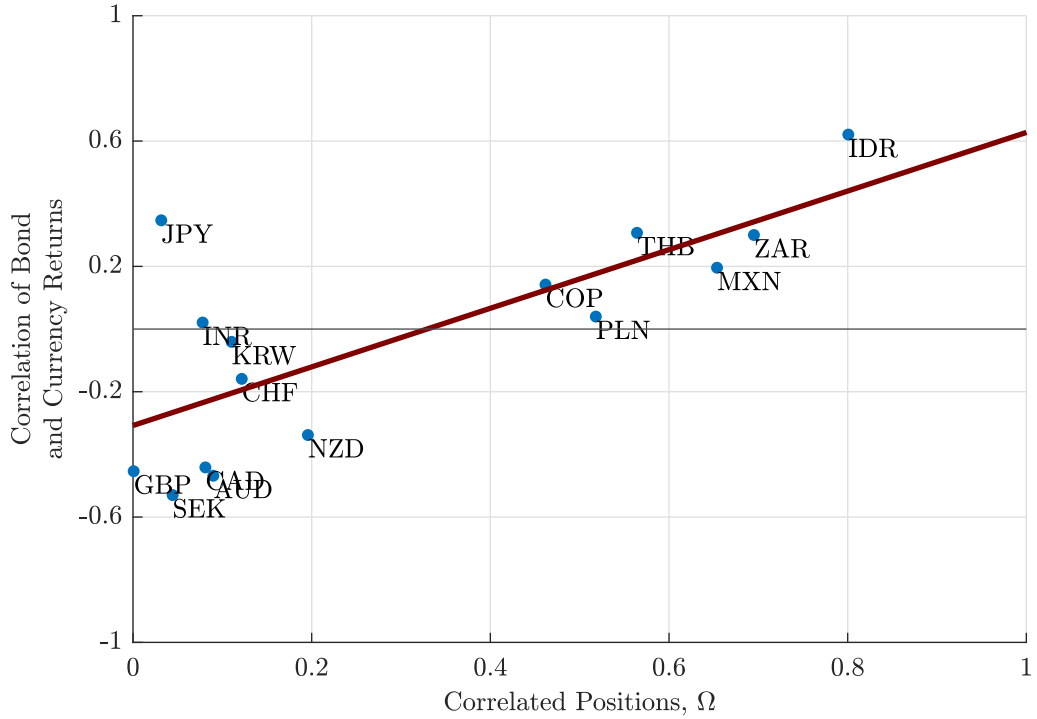


(a) Total bond holdings by entity

(b) Bond holdings across foreign investors

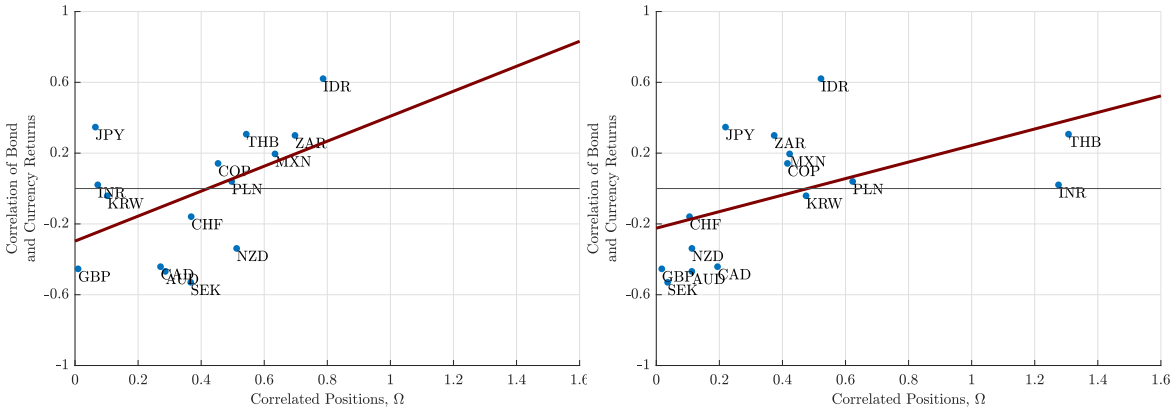
Note: The data presented correspond to all trades registered with the Central Securities Depository (DCV), which operates under the supervision of the Central Bank of Colombia. Panel A shows the TES participation (percentage of total outstanding TES volume) across major entities, while Panel B focuses specifically on the share held by foreign investors, detailing the distribution among different types of foreign investors within the total foreign share.

Figure B.2: Comovement in Excess Returns and Ω



Note: This figure displays the correlation of bond and currency excess returns (y-axis) and the Estimated Ω from equation (2.19) (x-axis). In the computation for Estimated Ω we restrict $\alpha \geq 0$.

Figure B.3: Comovement in Excess Returns and Ω

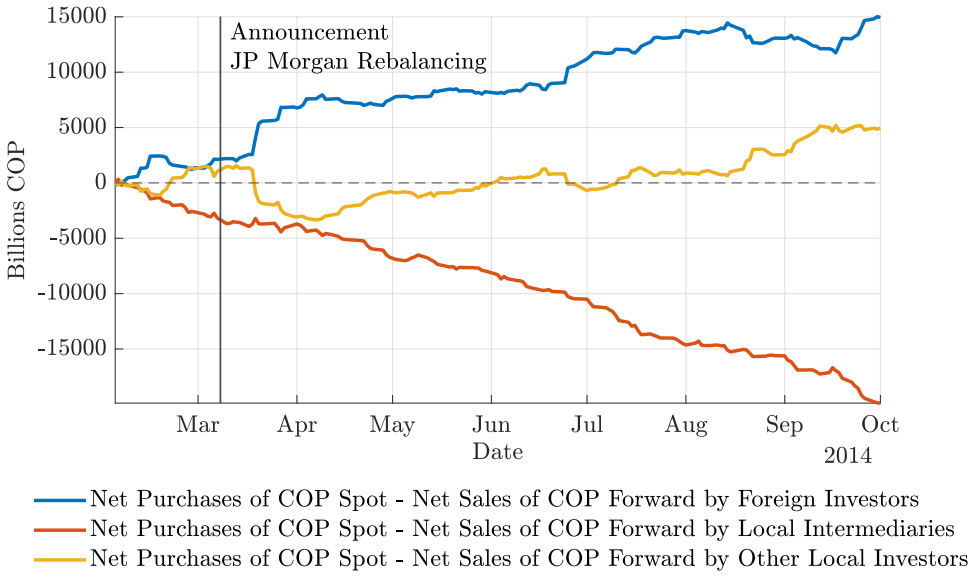


(a) Ω with median hedge ratios (b) Ω with median volatilities

Note: This figure displays the correlation of bond and currency excess returns (y-axis) and the Estimated Ω from eq. (2.19) (x-axis). In the computation for Estimated Ω in Panel (a) we restrict α to be the median hedge ratio for all countries. In Panel (b) we restrict the standard deviation for foreign and other domestics investors to the median across countries. Countries in red (blue) are part of the emerging (advanced) economies sample.

B.3 Additional Tables and Figures

Figure B.4: Flows in COP Spot and Forward markets during J.P. Morgan Index Rebalancing



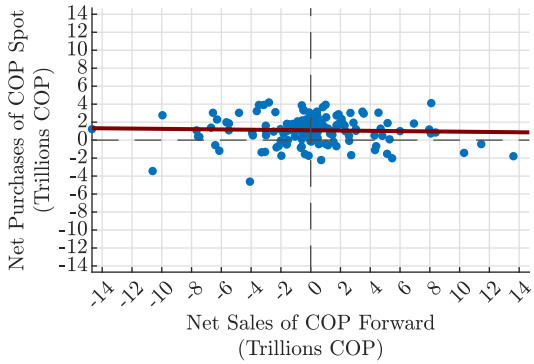
Note: The figure displays the evolution of the daily accumulated purchases of COP in the spot market net of sales of COP in the forward market during 2014 for different investor groups (described in Section 2.2 and Section 2.4). The black vertical line denotes 19th of March 2014, the day the rebalancing was announced.

Table B.2: Summary of FX and bond excess returns (1-month excess returns)

			Residualized excess returns	
Country	Regression coefficient	Correlation	Regression coefficient	Correlation
Australia	-0.70^{***}	-0.35	-0.78^{***}	-0.39
	(0.15)	$[-0.47, -0.22]$	(0.17)	$[-0.52, -0.24]$
Canada	-0.65^{***}	-0.39	-0.81^{***}	-0.47
	(0.12)	$[-0.50, -0.26]$	(0.13)	$[-0.59, -0.33]$
Colombia	0.25^{***}	0.27	0.34^{**}	0.34
	(0.09)	[0.13, 0.40]	(0.15)	[0.19, 0.48]
India	0.10	0.09	-0.06	-0.05
	(0.10)	$[-0.06, 0.23]$	(0.12)	$[-0.22, 0.12]$
Indonesia	0.57^{***}	0.59	0.64^{***}	0.60
	(0.11)	[0.46, 0.70]	(0.13)	[0.46, 0.71]
Japan	1.32^{***}	0.37	1.43^{***}	0.39
	(0.26)	[0.24, 0.49]	(0.34)	[0.24, 0.52]
Korea	-0.12	-0.07	-0.22	-0.12
	(0.15)	$[-0.25, 0.11]$	(0.18)	$[-0.31, 0.07]$
Mexico	0.65^{***}	0.40	0.69^{***}	0.39
	(0.17)	[0.25, 0.53]	(0.21)	[0.22, 0.54]
New Zealand	-0.55^{***}	-0.25	-0.52^{**}	-0.24
	(0.19)	$[-0.38, -0.11]$	(0.22)	$[-0.39, -0.08]$
Poland	0.43^{**}	0.21	0.50^{**}	0.25
	(0.17)	[0.07, 0.34]	(0.21)	[0.09, 0.40]
South Africa	0.96^{***}	0.53	1.11^{***}	0.58
	(0.11)	[0.42, 0.62]	(0.13)	[0.46, 0.68]
Sweden	-0.79^{***}	-0.41	-0.95^{***}	-0.47
	(0.15)	$[-0.52, -0.29]$	(0.17)	$[-0.59, -0.33]$
Switzerland	-0.17	-0.08	-0.07	-0.03
	(0.18)	$[-0.22, 0.06]$	(0.21)	$[-0.20, 0.13]$
Thailand	0.39^{*}	0.25	0.39^{*}	0.25
	(0.23)	$[-0.05, 0.51]$	(0.23)	$[-0.05, 0.51]$
United Kingdom	-0.57^{***}	-0.40	-0.58^{***}	-0.41
	(0.12)	$[-0.51, -0.27]$	(0.15)	$[-0.54, -0.26]$

Note: Column (1) of the table reports the estimated slope coefficient for the following baseline regression $rx^{q}_{j,t+1} = \alpha + \beta rx^{(10y)}_{j,t+1} + \varepsilon_{j,t+1}$. Column (2) reports the correlation between $rx^{q}_{j,t+1}$ and $rx^{(10y)}_{j,t+1}$. Column (3) reports the estimated slope coefficient of a regression of the residualized component of $rx^{q}_{j,t+1}$ on the residualized component of $rx^{(10y)}_{j,t+1}$. We residualize local-currency bond and FX excess returns according to the following regression $rx^{m}_{j,t+1} = \sum_{k=0}^{1} \phi_k (\text{Prob. default})_{j,t-k} + \varepsilon^{rx^m}_{j,t+1}$, for $m = \{10y, q\}$. Probability of default is the 10-year sovereign default probability, retrieved from *Refinitiv*. Column (4) reports the correlation between the residualized component of $rx^{q}_{j,t+1}$ and the residualized component of $rx^{(10y)}_{j,t+1}$. Stars denote significance: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. Values in parenthesis are standard errors, and values in square brackets represent the 95 percent confidence interval.

Figure B.5: Purchases of COP spot and sales of COP forward by foreign investors



Note: This figure displays the monthly-level purchases of COP in the spot market (x-axis) and the monthly-level sales of COP in the forward market (y-axis) by foreign investors. The fitted OLS regression line is also shown.

Table B.3: Participation and Bidding Patterns at TES bond and U.S. Dollar Auctions

	TES Auctions				U.S. Dollar Auctions		
Dealer	Participation	Barely Won	Barely Lost	Dealer	Participation	Barely Won	Barely Lost
H	97%	29%	2%	A	97%	75%	14%
O	96%	31%	2%	B	87%	54%	29%
A	94%	29%	2%	C	75%	39%	40%
I	86%	29%	1%	G	74%	40%	34%
P	81%	38%	3%	D	68%	53%	23%
J	81%	25%	8%	E	64%	55%	7%
Y	80%	29%	3%	F	62%	59%	9%
G	80%	33%	0%	M	48%	43%	27%
C	80%	30%	0%	H	47%	49%	30%
D	63%	13%	5%	K	45%	33%	35%
R	56%	37%	2%	J	45%	52%	32%
M	45%	29%	4%	I	39%	29%	33%
Z	43%	26%	5%	L	29%	15%	46%
L	27%	41%	2%	N	17%	38%	31%
W	27%	18%	5%	Q	8%	30%	54%
X	20%	17%	2%	O	6%	63%	20%
V	19%	10%	8%	P	2%	15%	62%

Note: Dealers (anonymized) are sorted by participation. Barely Winner and Barely Loser denote winners and losers within a bandwidth rank of $[-2, 2]$ around the cutoff price.

Table B.4: Transition Probability Matrix at TES bond and U.S. Dollar Auctions

	State at $t+1$			
State at t	Partially Winner	Winner	Partially Loser	Loser
Panel A: TES Auctions				
Partially Winner	16.4%	73.4%	2.3%	8.0%
Winner	32.8%	54.9%	3.0%	9.3%
Partially Loser	30.7%	50.7%	5.3%	13.3%
Loser	26.8%	38.8%	4.3%	30.1%
Panel B: USD Auctions				
Partially Winner	55.3%	11.8%	23.6%	9.4%
Winner	51.7%	28.2%	14.4%	5.6%
Partially Loser	43.9%	6.8%	32.6%	16.8%
Loser	33.4%	4.9%	35.6%	26.0%

Note: Each cell reports the probability (in %) of transitioning from state i at time t to state j at time $t+1$. States are mutually exclusive. Partially Winner and Partially Loser denote dealers who won or lost within a rank bandwidth of $[-2,2]$ around the cutoff price. Winner and Loser denote dealers who won or lost outside that bandwidth.

Table B.5: Estimated effects of asset purchases/sales

	TES Auctions		USD Auctions	
	TES Bond Prices	FX Prices	TES Bond Prices	FX Prices
Rank	2.13**	−0.05**	0.38	−0.01
	(0.94)	(0.023)	(0.28)	(0.01)
Dummy	−2.90*	0.06*	−0.61*	0.02**
	(1.72)	(0.03)	(0.36)	(0.01)
Dummy × Rank	−2.16***	0.04*	−1.34***	0.01
	(0.82)	(0.02)	(0.28)	(0.01)
Dummy × Exposure	0.001	−0.00002	0.0001	−0.00002***
	(0.001)	(0.00002)	(0.001)	(0.00001)
Constant	2.06	−0.02	−1.28***	−0.01
	(1.90)	(0.05)	(0.48)	(0.01)
Observations	238	181	615	603

Note: This table reports the estimates of equation (2.24), along with robust standard errors in parentheses. All specifications use a bandwidth of two ranks, consistent with Calonico et al. (2014).

Appendix C: Appendix to Chapter 3

C.1 Data Appendix

All macroeconomic variables are extracted from IMF (WEO and IFS), The World Bank, and national central banks. The index of capital controls is from Chinn and Ito (2006). Data spans from 1980-2022, when available. Data of residents' share of deposits denominated in foreign currency is from Honohan (2008). A total of 75 countries are included in the sample.

List of countries in the sample

Afghanistan	Guatemala	Nigeria
Albania	Hungary	Pakistan
Angola	India	Paraguay
Argentina	Indonesia	Peru
Armenia	Iraq	Philippines
Azerbaijan	Jamaica	Poland
Bangladesh	Kazakhstan	Romania
Belarus	Kenya	Russia
Bolivia	Korea, Rep.	Rwanda
Brazil	Kosovo	Serbia
Bulgaria	Kuwait	Seychelles
Cambodia	Kyrgyz Rep.	Sierra Leone
Chile	Latvia	Slovak Rep.
China	Libya	Slovenia
Colombia	Malaysia	South Africa
Congo, DR	Mauritania	Sri Lanka
Costa Rica	Mauritius	Tanzania
Croatia	Mexico	Thailand
Czech Rep.	Moldova	Tunisia
Dominican Rep.	Mongolia	Turkey
Ecuador	Morocco	Uganda
Egypt	Mozambique	Ukraine
Gambia	Myanmar	Uruguay
Georgia	Nepal	Vietnam
Ghana	Nicaragua	Zambia

Table C.1: Sample countries

C.2 Model Appendix

C.2.1 Solving the Model With Exchange Controls ($\phi \to \infty$)

In this section I will show the equivalence of the open and closed economy models whenever $I^* = 0$. Consider the case when $\phi_t \to \infty$ for all t. The shopping-time technology, V, is now

$$s_t = V\left[\frac{m_t}{c_t}\right] c_t, \tag{C.1}$$

where

$$V \geq 0, \quad V' \leq 0, \quad V'' > 0. \tag{C.2}$$

Like before, the agent chooses $\{c_t, h_t, m_t\}_{t=0}^{\infty}$ to maximize its utility subject to

$$\sum_{t=0}^{\infty}\left[\frac{1}{1+r}\right]^t(1-\tau_t)(1-h_t-s_t)=\sum_{t=0}^{\infty}\left[\frac{1}{1+r}\right]^t[c_t+I_t m_t]. \tag{C.3}$$

Again, one can solve this in two stages by noting that monetary holdings do not enter the utility function. Thus, in the first stage, the agent chooses $\{m_t\}_{t=0}^{\infty}$ to maximize its wealth (equation C.3), yielding

$$m_t: \qquad -V'(X_t^m)=I_t(1+\tilde{\tau}_t) \qquad\qquad t=0,1,2,... \tag{C.4}$$

Subsequently, and given the strict convexity of V, the implicit money demand function can be obtained

$$X_t^m=X^m(R_t) \qquad t=0,1,2,... \quad X_R^m<0. \tag{C.5}$$

Plugging this back in the lifetime budget constraint, I can rewrite the constraint for the second stage of the problem as

$$\sum_{t=0}^{\infty}\left[\frac{1}{1+r}\right]^t\left[\frac{(1-h_t)}{1+\tilde{\tau}_t}\right]=\sum_{t=0}^{\infty}\left[\frac{1}{1+r}\right]^t\left[\frac{q_t c_t}{1+\tilde{\tau}}\right], \tag{C.6}$$

where $q_t=q(\tilde{\tau},R_t)$ is the distortion generated by distortionary taxes:

$$q_t=1+\tilde{\tau}+V\left[X^m(R_t)\right]+R_t X^m(R_t). \tag{C.7}$$

The representative household chooses $\{c_t, h_t\}_{t=0}^{\infty}$ in the second stage to maximize its utility subject to C.6. Doing so, one can arrive at implicit functions for c and h which are functions of the distortion, q. Intuitively, welfare is decreasing in the level of the distortion, and thus the planner will desire to minimize it.

The government has to finance a constant path of spending $g_t=g$ for all t, and has

access to two taxes. Namely, the inflation tax and labor income tax. Conventional taxes carry collection costs $\kappa(\tau y)$ such that $\kappa'(\tau y) > 0$, and $\kappa''(\tau y) \geq 0$. Since g is constant, assuming that $\beta = \frac{1}{1+r}$ has the implication that this model is always in its steady state, and hence I drop the time subscripts.

$$g = \tau y + Im - \kappa(\tau y), \tag{C.8}$$

which can be rewritten, using the consumer's budget constraint and the fact that $y = 1 - h - s$, as

$$g = c\left[q(\tilde{\tau}, R)\right] \Omega(\tilde{\tau}, R), \tag{C.9}$$

where

$$\Omega(\tilde{\tau}, R) = \tilde{\tau} + RX^m - \frac{\kappa(\tilde{\tau} y)}{c}. \tag{C.10}$$

Since the distortion reduces household welfare, this optimal taxation problem can be re-cast into one where the government $(\tilde{\tau}, R)$ to minimize q, subject to its budget constraint. Joining the two first-order conditions yield the following optimality condition:

$$X^m = \frac{X^m + RX^m_R - \frac{1}{c}\frac{\partial \kappa(\tilde{\tau} y)}{\partial R}}{1 - \frac{1}{c}\frac{\partial \kappa(\tilde{\tau} y)}{\partial \tilde{\tau}}}. \tag{C.11}$$

This equation, along with C.9 are all I need to solve for the optimal levels of R and $\tilde{\tau}$ as a function of g. Once obtained, it would prove very easy to derive I and τ.

The equivalence between the economies with and without exchange controls can be by comparing 3.17 with C.9, and 3.18 with C.11 when $I^* = 0$. This proves the equivalence claim in the paper.

C.2.2 First- and Second-Order Derivatives of $\kappa(\cdot)$

Denoting that

$$\tilde{\tau y} = c\left[\tilde{\tau} + \frac{\tilde{\tau}}{1+\tilde{\tau}} RX^m + \tilde{\tau} I^* X^f\right]. \tag{C.12}$$

One can derive the first-order partial derivatives of $\kappa(\tilde{\tau y})$ with respect to the different taxes:

$$\frac{\partial \kappa(\tilde{\tau y})}{\partial R} = \kappa'(\tilde{\tau y}) c\left[\frac{\tilde{\tau}}{1+\tilde{\tau}}(X^m + RX^m_R) + \tilde{\tau} I^* X^f_R\right]; \tag{C.13}$$

$$\frac{\partial \kappa(\tilde{\tau y})}{\partial \tilde{\tau}} = \kappa'(\tilde{\tau y}) c\left[1 + \frac{RX^m}{(1+\tilde{\tau})^2} + \frac{\tilde{\tau}}{1+\tilde{\tau}} RI^* X^m_{R^*} + I^* X^f + \tilde{\tau} {I^*}^2 X^f_{R^*}\right]; \tag{C.14}$$

$$\frac{\partial \kappa(\tilde{\tau y})}{\partial I^*} = \kappa'(\tilde{\tau y}) c\left[\frac{\tilde{\tau}}{1+\tilde{\tau}} R\tilde{\tau} X^m_{R^*} + \tilde{\tau} X^f + \tilde{\tau}^2 I^* X^f_{R^*}\right]. \tag{C.15}$$

As well as the relevant second-order derivatives:

$$\begin{aligned}\frac{\partial^2 \kappa(\tilde{\tau y})}{\partial RR} = {} & \kappa''(\tilde{\tau y})\left[c\left(\frac{\tilde{\tau}}{1+\tilde{\tau}}(X^m + RX^m_R) + \tilde{\tau} I^* X^f_R\right)\right]^2 \\ & + \kappa'(\tilde{\tau y}) c\left[\frac{\tilde{\tau}}{1+\tilde{\tau}}(X^m_R + X^m_R + RX^m_{RR}) + \tilde{\tau} I^* X^f_{RR}\right];\end{aligned} \tag{C.16}$$

$$\begin{aligned}\frac{\partial^2 \kappa(\tilde{\tau y})}{\partial \tilde{\tau}\tilde{\tau}} = {} & \kappa''(\tilde{\tau y})\left[c\left(1 + \frac{RX^m}{(1+\tilde{\tau})^2} + \frac{\tilde{\tau}}{1+\tilde{\tau}} RI^* X^m_{R^*} + I^* X^f + \tilde{\tau} {I^*}^2 X^f_{R^*}\right)\right]^2 \\ & + \kappa'(\tilde{\tau y}) c\left[\frac{2RI^* X^m_{R^*} - 2RX^m}{(1+\tilde{\tau})^2} + \frac{\tilde{\tau}}{1+\tilde{\tau}} RI^{*^2} X^m_{R^*R^*} + 2I^* X^f_{R^*} + \tilde{\tau} I^{*^3} X^f_{R^*R^*}\right];\end{aligned} \tag{C.17}$$

$$\begin{aligned}\frac{\partial^2 \kappa(\tilde{\tau y})}{\partial R\tilde{\tau}} = {} & \kappa''(\tilde{\tau y}) c^2\left[\frac{\tilde{\tau}(X^m + RX^m_R)}{1+\tilde{\tau}} + \tilde{\tau} I^* X^f_R\right]\left[1 + \frac{RX^m}{(1+\tilde{\tau})^2} + \frac{\tilde{\tau} RI^* X^m_{R^*}}{1+\tilde{\tau}} + I^* X^f + \tilde{\tau} {I^*}^2 X^f_{R^*}\right] \\ & + \kappa'(\tilde{\tau y}) c\left[\frac{X^m + RX^m_R}{(1+\tilde{\tau})^2} + \frac{\tilde{\tau}(I^* X^m_{R^*} + RI^* X^m_{RR^*})}{1+\tilde{\tau}} + I^* X^f_R + \tilde{\tau} {I^*}^2 X^f_{RR^*})\right];\end{aligned} \tag{C.18}$$

$$\frac{\partial^2 \kappa(\tilde{\tau}y)}{\partial \tilde{\tau} R} = \kappa''(\tilde{\tau}y)c^2 \left[\frac{\tilde{\tau}(X^m + RX_R^m)}{1+\tilde{\tau}} + \tilde{\tau} I^* X_R^f \right] \left[1 + \frac{RX^m}{(1+\tilde{\tau})^2} + \frac{\tilde{\tau} R I^* X_{R^*}^m}{1+\tilde{\tau}} + I^* X^f + \tilde{\tau} I^{*^2} X_{R^*}^f \right]$$
$$+ \kappa'(\tilde{\tau}y)c \left[\frac{X^m + RX_R^m}{(1+\tilde{\tau})^2} + \frac{\tilde{\tau}(I^* X_{R^*}^m + R I^* X_{R^* R}^m)}{1+\tilde{\tau}} + I^* X_R^f + \tilde{\tau} I^{*^2} X_{R^* R}^f) \right]. \tag{C.19}$$

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