Essays in Macroeconomics:

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Abstract

The dissertation studies the sources of business cycles taking both an open and a closed economy perspective. A common feature of the two chapters composing the dissertation is the use of simple, but powerful classifications and identifications of sources of business cycles. In particular, the first chapter, titled "What are the Sources of Boom-Bust Cycles?", concerns the distinction between economic fluctuations due to changes in beliefs, and fluctuations due to changes in fundamentals, showing results that challenge traditional approaches to modeling business cycles. The second chapter, titled "Shocks and Exchange Rates in Small Open Economies", takes the perspective of small open economies, and concerns the distinction between global and domestic shocks, showing results that are informative for a series of puzzling facts concerning the dynamics of the exchange rate.

In "What are the Sources of Boom-Bust Cycles?," joint with Marco Brianti, we provide a synthesis of two major views on economic fluctuations. One view maintains that expansions and recessions arise from the interchange of positive and negative persistent exogenous

shocks to fundamentals. This is the conventional view that gave rise to the profusion of shocks used in modern dynamic stochastic general equilibrium models. In contrast, a second view, which we call the endogenous cycles view, holds that business cycle fluctuations are due to forces that are internal to the economy and that endogenously favor recurrent periods of boom followed by a bust. In this environment, cycles can occur after small perturbations of the long run equilibrium. We find empirical evidence pointing at the coexistence of both views. In particular, we find that the cyclical behaviour of economic aggregates is due in part to strong internal mechanisms that generate boom-bust phenomena in response to small changes in expectations, and in part to the interchange of positive and negative persistent fundamental shocks. Motivated by our findings, we build a theory that unifies the dominant paradigm with the endogenous cycles approach. Our theory suggests that recessions and expansions are intimately related phenomena, and that understanding the nature of an expansion, whether it is driven by fundamentals or by beliefs, is a first order issue for policy makers whose mandate is to limit the occurrance of inefficient economic fluctuations.

In “Shocks and Exchange Rates in Small Open Economies,” joint with Pierre De Leo, we propose a novel approach to separately identify domestic and external shocks in small open economies. Our results provide guidance about the transmission mechanism of these shocks and revisit recent conclusions drawn on the exchange rate effects of monetary policy in small open economies. The identification method is based on the premise that shocks originating

from within a small economy should not influence world variables at any horizon, while external (or global) shocks should affect world variables at least at some horizon. We obtain three empirically related findings. First, external shocks feature large deviations from uncovered interest parity, while domestic shocks do not. Second, external shocks strongly comove with global risk aversion and U.S. macroeconomic variables. Third, recent puzzling estimates of the exchange rate effects of monetary policy stem from an identification of domestic shocks that fails to properly account for international spillovers. We show that a two-country small open economy model with international asset market imperfections is consistent with these facts. In our proposed model, global risk aversion shocks drive exchange rate dynamics, and a country's net foreign asset position governs their international transmission. We provide empirical evidence that a country's exposure to external shocks indeed depends on its net foreign asset position.

A zio Franco

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Contents

List of Tables

List of Figures

Chapter 1

What are the Sources of Boom-Bust Cycles?

1.1 Introduction

This paper provides a synthesis of two major views on economic fluctuations. One view maintains that expansions and recessions arise from the interchange of positive and negative persistent exogenous shocks to fundamentals. This is the conventional view that gave rise to the proliferation of shocks embedded in modern dynamic stochastic general equilibrium (DSGE) models. A second view, which we call the endogenous cycles view, holds that business cycle fluctuations are due to forces that are internal to the economy and that favor recurrent periods of boom followed by an *endogenous* bust. In this environment, cycles can occur even in absence of shocks to fundamentals. Conclusive evidence in favor of either view is hard to find. One reason may be that a complete representation of the economy is one in which both views coexist.

We make three contributions. First, we document a data conundrum that stems from contrasting unconditional and conditional evidence on the presence of endogenous cycles. We build on Beaudry et al. (2019, BGP henceforth) who provide compelling evidence that U.S. macroeconomic aggregates tend to move in regular cycles. We ask whether fundamental sources of fluctuations, such as technology shocks, can explain the regular cyclicality present in the data. We find that fundamental-driven expansions do not feature predictable future recessions and therefore fundamental shocks cannot account for the unconditional moments documented by BGP. Second, we build a theory that rationalizes the conundrum and proposes shocks to expectations as the key source of boom-bust cycles. According to our theory, positive shocks to expectations, such as waves of optimism, generate periods of boom that are *endogenously* followed by a recession. In contrast, expansionary fundamental shocks do not generate predictable busts. Thus, our theory provides new discipline for both the conventional view of exogenous cycles *and* the more heterodox view of endogenous cycles, by restricting their domain of application to fundamental shocks or to expectation shocks, respectively. Third, we identify expectation shocks using survey data from the U.S. and verify that, indeed, expectations shocks (*i*) generate predictable boom-bust episodes, (*ii*) bring about economic dynamics quantitatively consistent with our model, and (*iii*) account for a sizeable fraction of business cycle fluctuations.

In the first part of the paper, we begin by documenting the presence of a systemic cyclical behaviour in economic aggregates. We do so in two ways. First, we show that the spectral densities of a number of U.S. macroeconomic and financial variables display a peak at periodicities of around 8 to 10 years. A hump-shaped spectral density signals the presence of periodic motions that repeat themselves in a regular cycle. Second, we show that the probability of a recession peaks about two years after an expansion – findings that are *inconsistent* with the predictions of standard DSGE models. Next, we argue that the responses to identified fundamental shocks almost always deliver mean-reverting responses more aligned with the conventional view. We take a temporary shock to utilization-adjusted TFP as the leading case. A positive TFP shock leads to a temporary expansion that is not systematically followed by a recession. By comparing the conditional spectral densities implied by a TFP shock with their unconditional counterparts, we show that these shocks cannot be responsible for the cyclical properties of the data.

The presence of systemic cyclicality that is not due to fundamental shocks poses a conundrum. In the second part of the paper, we propose a general equilibrium model that rationalizes such conundrum. Given the particularly pronounced evidence of cyclical behaviour among financial variables, we place financial frictions at the hearth of our theory. The structure of the model echoes Jermann and Quadrini (2012) in that there are firms who borrow from households by issuing short- and long-term debt. Short-term debt is in the form of an intra-period working capital loan and therefore it is used to finance production inputs. For simplicity, we assume that the long-term debt is in the form of a one period bond that firms issue to smooth out dividends. The central innovation of the model is a default-deterring borrowing constraint that depends positively on firms' market value.

The endogenous borrowing limit has two important features. First, it introduces a pecuniary externality as firm's market value, defined as the discounted cumsum of its future cash flows, depends upon two components: firm's future profits which are under the direct control of the firm, and households' stochastic discount factor which the single firm takes as given. Second, it generates strong financial amplification due to a positive feedback loop between firms' market value and households' income. These two features combined make the model economy display boom-bust episodes. Crucially, busts arise endogenously after expansions led by positive shifts in agents' expectations, but not after expansions due to positive shocks to technology.

The intuition is as follows. Suppose that households become more optimistic regarding firms' future value so that equity prices increase. Increased equity prices relax borrowing constraints and allow firms to issue more short- and long-term debt. Since short-term debt is useful to finance production, looser borrowing constraints raise firms' demand for labor. The resulting higher wages increase households' labor income, their willingness to save, thereby leading to a further increase equity prices, and a relaxation of borrowing constraints. Therefore, an expectation-driven expansion features increasing equity prices, wages, debt, and output, due to a positive feedback loop between firms' market value and households' income. Crucially, the increase in equity prices is due to higher households' stochastic discount factor, and *not* to a change in firms' future profits. In fact, increased borrowing capacity increase wages which reduce firms' marginal profits and may lead to a profits *decline* over the expectation-driven expansion. As the economy evolves, lower firms' profitability distorts firms' incentives in such a way to trigger a recession. From the perspective of a single firm, lower profitability reduces its incentives to allocate borrowing capacity into working capital. However, firms fail to internalize that by hiring less input, households will receive less labor income, commanding a fall in equity prices and tighter borrowing constraints.

Intuitively, the amplification channel should deliver similar boom-bust responses after shocks to technology, but it does not. The reason is that equity prices increase primarily because of the increase in firms' profits thanks to higher productivity. Because of the increased profitability, firms will allocate funds predominantly to hire more inputs until the shock is absorbed, and, as a consequence, the distortion coming from the pecuniary externality will be less important.

We argue that changes in expectations distinct from changes of technology can rationalize the boom-bust features of the data, but what triggers such changes? The model's answer is that equilibrium outcomes are the product of self-fulfilling shifts in agents' expectations, and when these changes are unrelated to fundamentals they generate boom-bust dynamics. The intuition is that boom-bust dynamics obtain when the internal financial amplification channel is sufficiently strong, but this happens *only* in the case in which the dynamic equilibrium is indeterminate, that is, the economy is subject to self-fulfilling shifts in expectations (a.k.a. sunspots).

In the third part of the paper, we empirically identify expectation shocks and test the predictions of the model. Specifically, we construct an indicator that summarizes the revisions of expectations on the future economic outlook using quarterly data on expectations from the Survey of Professional Forecasters and the Survey of Consumers. We use the indicator to identify exogenous shifts in expectations that are uncorrelated with past, present and future realizations of TFP. In addition, we control for a number of leads and lags of shocks to expectations of TFP in order to isolate shifts in expectations that are pure sentiments from those originating from beliefs on future TFP. Using local projections, we find that expectation shocks generate significant boom-bust dynamics in all the aggregate variables that we examine, and explain up to 40% of real GDP at business cycle frequencies, consistent with the findings of Angeletos et al. (2018) and Chahrour and Ulbricht (2019).

Finally, we show that the mechanism of the model is consistent with many features of the data. First, we find that the model is able to reproduce the empirical impulse responses to both expectation and TFP shocks. As in the model, expectation shocks bring about a countercyclical movement of the labor wedge, while the labor wedge *increases* after TFP improvements. Second, we show that the model can replicate the reduced-form evidence on boom-bust cycles that motivated our analyses. Unlike standard business cycle models, our theory can explain both the hump in the spectral densities of macroeconomic and financial variables, and the rising probability of a recession during an expansion.

Related literature. This paper lies at the intersection between the strand of the finance literature that focuses on credit cycles and the broad macroeconomic literature that aims at understanding the sources of business cycles.

The idea that the financial system is prone to generate economic instability through endogenous credit booms traces back at least to Kindleberger (1978) and Minsky (1975,1986). Minsky (1986) provides groundbreaking insights on the relation between the economic and the financial system. Of particular interest for this paper is his distinction between "periods of tranquility," defined as situations during which the economy is not subject to disruptive changes, and "unstable times" during which market forces lead to a rise of financial instability which culminates in "speculative frenzies". Through the lenses of our model and empirical evidence, we view such "periods of tranquility" as moments during which technological changes are the major contributor to economics fluctuations, whereas "unstable times" are characterized by economic fluctuations primarily driven by changes in market expectations.

More recently, the idea that an increase in credit associated with a decrease in borrowing costs can be a powerful predictor of future economic crises has been empirically tested and verified using both macro and micro level data. For example, Schularick and Taylor (2012) and Jordà et al. (2013), using data on 14 developed countries from 1870 to 2008, demonstrate that rapid credit expansions forecast declines in real activity.[1] Using data on the credit quality of corporate debt issuers, Greenwood and Hanson (2013) find that a

[1] Other examples include Demirgüç-Kunt and Detragiache (1998), Hardy and Pazarbasioglu (1998), Kaminsky and Reinhart (1999), Gourinchas et al. (2001), Goldfajn and Valdes (2006), Borio and Drehmann (2009), Reinhart and Rogoff (2009), Claessens et al. (2011), Gourinchas and Obstfeld (2012), and Laeven and Valencia (2013).

high share of risky loans tends to forecast low corporate bond returns. Krishnamurthy and Muir (2017) show that crises are preceded by a period of high credit to GDP growth and leverage, and low spread and risk premium. We complement this literature by providing conditional evidence on the link between a credit boom and the ensuing recession. We show that positive expectation shocks - but not TFP shocks - are systematically followed by a recession. Our evidence on expectation shocks also relates to López-Salido et al. (2017) who focus on credit market sentiment identified using credit spreads and find that high credit market sentiments are a predictor of future negative output growth. We complement their analysis by showing that sentiment shocks not only predict a negative output growth but also prolonged periods during which the *level* of output is below trend.

We relate to the literature that aims at rationalizing boom-bust phenomena. For example, Boissay et al. (2016) rationalize boom-bust episodes in a model where the increase in households' savings during a boom exacerbates adverse selection problems in the interbank market. In our model, the increase in savings brings about a recession because it reflects an increase in firms' debt which tightens financial markets. A subset of this literature builds model of chaos and limit cycles. Boldrin and Woodford (1990) survey the literature and analyze the conditions under which limit cycles can emerge. In a recent paper, Beaudry et al. (2019) revisit the reduced-form evidence on the spectral densities of a series of economic variables. They build a model of limit cycles where small exogenous shocks give rise to perpetual economic cycles. While our model can also exhibit limit cycles for regions of the parameter space that imply a sufficiently tight financial constraint, our aim is rather to rationalize the fact that only a subset of shocks trigger oscillatory dynamics while other shocks do not. Gorton and Ordonez (2016) distinguish between “good” and “bad” credit booms depending whether or not they end up in a crisis. They find that shocks in the trend of productivity are associated with “good” credit booms, whereas “bad” booms are typically associated with a decline in productivity. We differ from them in at least two aspects. First, we look at cycles at short and medium-run frequencies while their focus is on booms that last ten years on average. Second, we emphasize that the shocks responsible for boom-bust episodes are orthogonal to movements of TFP.

Furthermore, we relate to the class of models that generate self-fulfilling rational expectations equilibria due to credit market amplification. Examples of this class are Benhabib and Wen (2004), Benhabib and Wang (2013), Liu and Wang (2014), and Azariadis et al. (2015). While their emphasis is on a single shock, our model is built to capture the important different responses to fundamental and sunspot shocks.

Lastly, our theoretical framework shares some similarities with models of stock market bubbles as in Miao and Wang (2018), in that, debt limits depend upon firms’ market value and sentiment shocks can be interpreted as bubbles. However, models of stock market bubbles formalize the burst of a bubble as an exogenous event. In contrast, in our model sentiment shocks rationalize both the formation of a bubble and its subsequent burst.

1.2 The *cyclicality* conundrum

Boom-bust cycles are a recurrent feature of the data. Yet, there is virtually no evidence of boom-bust dynamics conditional on shocks. We refer to such incoherence between unconditional and conditional evidence as the *cyclicality* conundrum. This section documents the conundrum by showing that (i) there is a systemic cyclical component in the data and (ii) shocks to fundamentals do not impart economic dynamics that can account for such systemic cyclicality.

1.2.1 Unconditional evidence of cycles

In a recent article, Beaudry et al. (2019) find that U.S. business cycles are characterized by cyclical forces. In particular, they show that the spectral densities of a number of economic aggregates exhibit a common local peak at periodicities of 32 to 50 quarters. The spectral density is a useful diagnostic tool of cyclicality for two reasons.[2] First, a peak in the spectral density signals the presence of oscillatory dynamics in the autocovariance function of the data. Second, it tells us whether these oscillatory dynamics happen at business cycle frequencies or they reflect lower frequency forces unrelated to business cycles.

Figure 1.1 reports the spectral density of a series of macroeconomic and financial variables.[3] We use quarterly data from 1967:q1 to 2018:q4 and detrend variables using a band

[2]The notion of cyclicality that we use is analogous to Beaudry et al. (2019), that is a series is cyclical if its autocovariance function displays oscillations.

[3]The spectral density is computed using the Schuster's periodogram.

pass filter that removes fluctuations with periodicities longer than 100 quarters.[4,5] Two patterns emerge. First, results point at the presence of a strong common cyclical component. With the exception of utilization-adjusted TFP, all variables exhibit a peak in the spectral density in the interval between 32 and 50 quarters. Furthermore, the fact that there are no notable differences in the shape of the spectral density across variables, suggests the presence of an underlying mechanism responsible for the cyclical patterns rather than idiosyncrasies in the variables examined. Second, financial variables exhibit a more pronounced peak relative to macroeconomic variables suggesting that the cyclical features of the data might originate from shocks propagating through the financial sector, whereas shocks that primarily hit the real sector of the economy generate less oscillatory dynamics.

Importantly, a hump-shaped spectral density is a finding inconsistent with the predictions of standard business cycle models. In Figure **??** in appendix A.2 we run a Monte Carlo simulation on the spectral density of output using a textbook Real Business Cycle model and the New-Keynesian model by Smets and Wouters (2007). We find that the spectral density of output from model simulated data is counterfactually increasing in the periodicity.

[4]Because filtering the series could induce a spurious hump in the spectral density, we check that results are robust to various detrending techniques and frequency bands.

[5]The choice of the data sample does not affect the results. We start from 1967 as it is consistent with the longest data sample available for the analyses carried in Section 1.4.

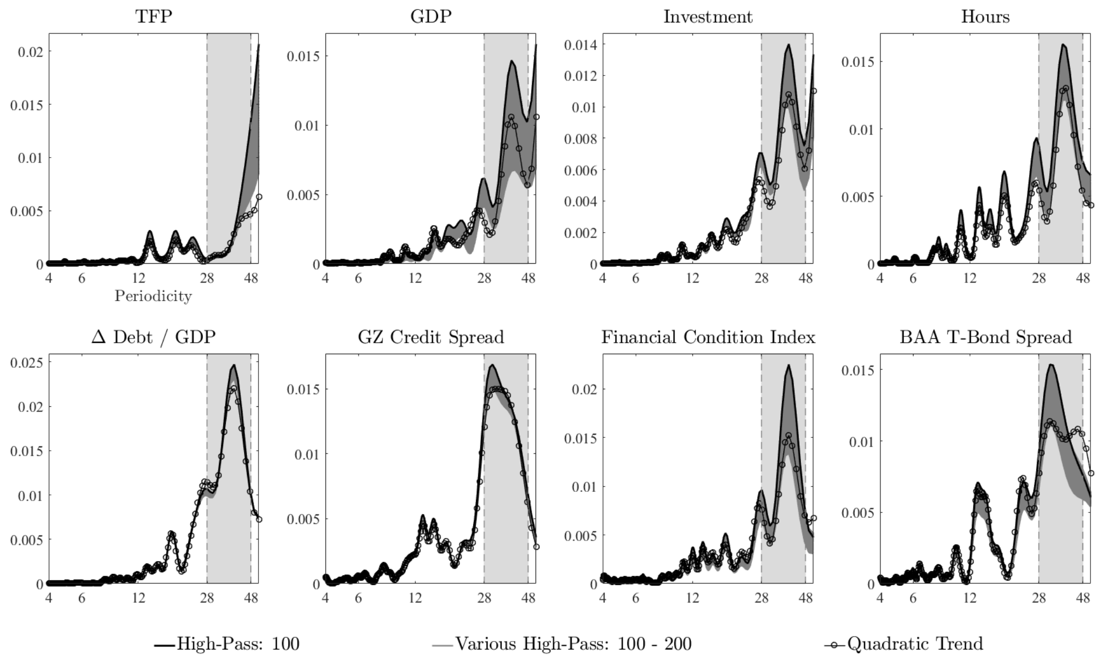


Figure 1.1: Unconditional spectral densities of quarterly U.S. signal systemic cyclicality

Note: Data from 1967:q1 to 2018:q4. TFP is utilization-adjusted total factor productivity. GDP is real gross domestic product. Investment is real consumption of durables plus real gross private domestic investment. Hours is hours of all persons in non-farm business sector. Change in debt is the flow of nonfinancial business debt securities and loans. GZ Credit Spread is the measure of credit spread described in Gilchrist and Zakrajšek (2012). Financial Conditions Index is provided by Chicago Fed. BAA T-Bill Spread is the difference between the yield of BAA corporate bonds and the treasury note at 10-year horizon. Series are detrended using a quadratic trend (circle-solid line), a filter that excludes fluctuations of period greater than 100 (black line), or from 101 to 200 (dark grey lines).

The presence of a systemic cyclical component in the data implies that the probability that a recession occurs should increase after an expansion. To verify whether this is true, we estimate a linear probability model and compute the probability that the economy enters in a recessions after k quarters since the previous expansion. We define expansions as periods in which real GDP growth is above the top quintile for at least two consecutive quarters. Likewise, we construct a recession indicator that takes value equal one if the real GDP growth falls into the bottom quintile for at least two consecutive quarters. Figure 1.2 plots the probability that the economy will be in a recession in a two-quarter window around time $t+k$ given an expansion at time t. Results confirm the evidence of cyclicality described above. The conditional probability of a recession increases after an expansion and peaks approximately after two years. The picture also shows the prediction from data simulated using standard business cycle models such as the one described in Smets and Wouters (2007), the textbook Real Business Cycle model. In addition, we run the same experiment using the incomplete information model of Blanchard et al. (2013). All models predict that recessions are effectively unforecastable, in that the probability of a recession quickly converges to its unconditional mean after an expansion. To see this, we plot the results from simulating a random walk process in levels and show that the results from all models considered are indistinguishable from the predictions obtained after simulating a random walk for real GDP.

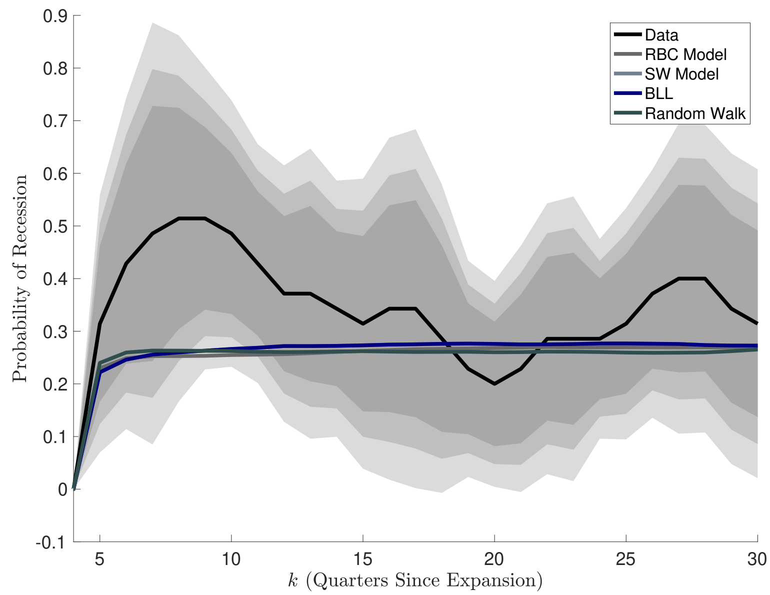


Figure 1.2: Probability of a recession peaks two years after an expansion

Note: Probability of recession in a two-quarter window after k quarters since expansion. Confidence intervals are 68%, 80%,and 90% (shaded areas) around the point estimate (solid black line).

1.2.2 Conditional rejection of cycles

Ultimately, we are interested in understanding the *sources* of the oscillatory behaviour documented above. To this end, we ask whether technology shocks account for these empirical regularities. We use quarterly utilization-adjusted TFP (Basu et al., 2006) and identify technology shocks as the innovation of detrended TFP after regressing it on its own lags, lags of the first principal component of a large dataset of aggregate economic variables and news shocks estimated following Barsky and Sims (2011a).[6] We estimate impulse responses

[6]Results are robust to different detrending techniques, additional controls, and different number of lags and principal components. See Appendix A.3 for results and additional details.

using the method of local projections proposed by Jordà (2005). Specifically, we estimate the h-th coefficient of the impulse response function by regressing each variable at time $t+h$ on the shock at time t.[7] We choose to implement the method of local projections because unlike vector autoregressions (VAR), it does not require to specify the lag structure of the data generating process.

[7] Details on local projections are in the Appendix A.5.

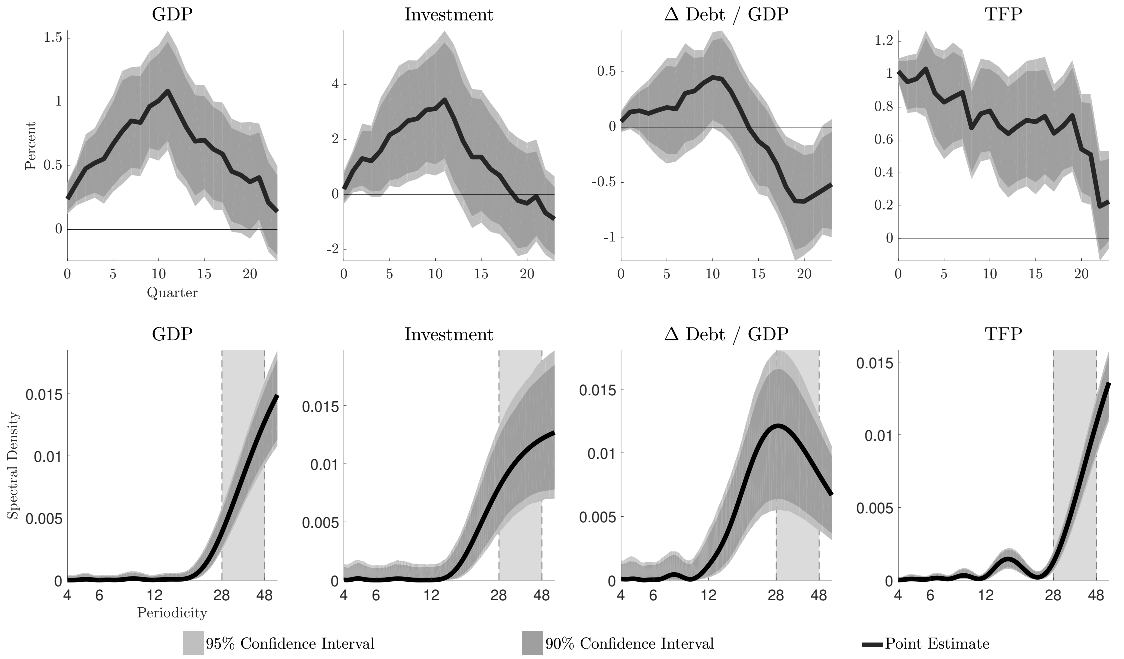


Figure 1.3: Impulse responses and spectral densities of a TFP shock.

Note: Technology shocks are the innovation of detrended TFP after regressing it on its own lags, lags of the first principal component of a large dataset of aggregate economic variables and news shocks estimated as in Barsky and Sims (2011a). Impulse responses (top panel) are estimated using local projections method. Confidence intervals are computed using the block-bootstrap method described in Kilian and Kim (2011). Conditional spectral densities (bottom panel) are computed from the Fourier transform of the estimated MA.

The top panel of Figure 1.3 shows the impulse responses of real GDP, investment and the change in nonfinancial corporate debt as a fraction of GDP, to a positive transitory technology shock. An unanticipated improvement of TFP leads to a hump-shaped response of real GDP and investment, aggregate debt rises during the initial build-up and decreases while the economy returns to its long run trend. To verify whether these impulse responses can account for the spectral properties of the data, we compute the spectral densities implied by the estimated coefficients of the moving averages. The bottom panel of Figure 1.3 shows that the spectral densities of real GDP and investment conditional to a TFP shock are monotonically increasing over business cycle periodicities. This poses a challenge to TFP-based explanations of boom-bust cycles.

Conditional test for the presence of a local peak The lack of a local peak in the spectral density of output, investment, and TFP observed in Figure 1.3 suggests that technology shocks cannot account for spectral properties of the data shown in Figure 1.1. To make the point, we construct a test for the presence of a significant local peak in the spectral density conditional to a structural shock. The test procedure echoes Canova (1996) and Reiter and Woitek (1999) who design a test for the presence of a peak for the unconditional spectral density. Details of our procedure are presented in the Appendix A.7. The idea is to test if the shape of the conditional spectral density around a particular frequency range is not statistically different from the spectral density implied by an autoregressive process of order one. More specifically, define D_1 the average estimated spectral density over a range around 34 quarters, and D_2 the average estimated spectral density over a range around 45 quarters. The test statistic is the ratio $D \equiv D_1/D_2$. A value of D bigger than one indicates the spectral density is decreasing in the range 34 to 45 quarters. The spectral density associated to an AR(1) process, in contrast, is monotonically increasing in the periodicity. Therefore we test the null hypothesis $H_0 : D = D^*$ where D^* is the value implied by an AR(1) with persistent parameter estimated from the data, against the alternative $H_1 : D > D^*$. Results for the technology-implied spectral density are reported in Table A.1. We fail to reject the null hypothesis of absence of a local peak for GDP, investment, and TFP.

Taken together our reduced form and conditional evidence points at the presence of oscillatory properties of the data that do not appear to be captured by movements in TFP. In the next section we build a model that helps us rationalizing the findings and propose "pure" sentiment shock - defined as shifts in expectations unrelated to fundamental - as a natural candidate to explain the spectral properties of the data. In section 1.4 we construct novel empirical evidence in favor of this hypothesis and show that the model can reproduce the responses to sentiment and technology shocks together with the unconditional spectral densities of the data.

1.3 A model of conditional cycles

In this section we show that a standard Real Business Cycle model augmented with financial frictions can rationalize the cyclicality conundrum. Azariadis et al. (2015) document that unsecured firm credit is procyclical whereas collateralized debt is acyclical. Building on their findings, we assume a type of solvency constraint that allows firms to borrow up to a fraction of their market value. Furthermore, we introduce short and long term debt as in Jermann and Quadrini (2012). This form of financial friction combined with procyclical fluctuations of long-term debt generate strong internal amplification and cyclical dynamics in response to serially uncorrelated shifts in expectations. For plausible parametrizations of the financial constraint, we find that the model displays dynamic multiplicity of equilibria due to self-fulfilling changes in expectations (a.k.a sunspots). In this environment, waves of optimism unrelated to present and future fundamentals, generate temporary expansions followed by recessions.

Importantly, our model stands in stark contrast to the class of models of self-fulfilling business cycle due to aggregate increasing returns to scale as described in Benhabib and Farmer (1994).[8] Amplification in the form of increasing returns would strongly influence the transmission of technology shocks, thus, while these models can generate endogenous oscillatory dynamics, they cannot *simultaneously* account for the empirical evidence on technology shocks.

For expositional reasons, we present first a benchmark model featuring intertemporal debt as the only state variable. In the next section we identify sentiment shocks in the data and augment the model with capital and external consumption habit to match empirical responses. We further validate model's performance by showing that it does a good job in matching the spectral properties of the data.

[8] Examples in this class are Farmer and Guo (1994), Wen (1998), and Liu and Wang (2014).

1.3.1 Firm sector

There is a continuum $i \in [0,1]$ of firms with a gross revenue function $F(z_t, k_t, n_t) = z_t k_t^{\theta} n_t^{1-\theta}$. The variable z_t is the stochastic level of productivity common to all firms, n_t is the labor input, k_t is the capital input which we assume to be constant and equal to one for now. Firms issue noncontingent bonds b_{t+1} at a price b_{t+1}/R_t. We assume that firms receive a tax advantage such that given the interest rate r_t, the effective gross interest rate for the firm is $R_t = 1 + r_t(1-\tau)$ where τ is the tax benefit. Thus, firms are effectively more impatient than households so that if financial markets are not too tight the equilibrium stock of debt will be positive. In addition to the intertemporal debt, firms raise funds with an intraperiod loan, ℓ_t, to finance working capital. Because revenues are realized at the end of the period, working capital is required to cover the intraperiod cash flow mismatch. The loan ℓ_t is paid at the end of the period with no interest.[9]

The timing of the events is the same as in Jermann and Quadrini (2012). Shocks realize at the beginning of the period. Firms enter the period with outstanding debt equal to b_t and choose labor n_t, the new intertemporal debt b_{t+1} and distribute dividends d_t. Since payments are made before producing, the intraperiod loan is

$$\ell_t = w_t n_t + \phi(d_t) + b_t - b_{t+1}/R_t,$$

where $\phi(d_t) = d_t + \kappa(d_t - \bar{d})^2$ includes a convex distribution cost of dividends which captures

[9] The assumption of two types of debt is made for analytical convenience. In particular the intratemporal debt can be replaced with cash that firms carry from the previous period. Cash would then be used to finance working capital and pay part of dividends.

documented evidence of preferences for dividend smoothing (Lintner, 1956). The end of period firm's budget constraint is

$$b_{t+1}/R_t + F(z_t, n_t) = w_t n_t + \phi(d_t) + b_t. \tag{1.1}$$

It follows that firm's revenues are equal to the intraperiod loan, that is $\ell_t = F(z_t, n_t)$.

Incentive constraint. When production is complete, firms decide whether or not repay the intraperiod loan they owe to the household. Consistent with recent evidence on the procyclicality of unsecured debt (see Azariadis et al., 2015), we assume that contract enforcement is imperfect so that firms have incentives to default. If a firm defaults it can divert its end of period revenues $y_t \equiv F(z_t, n_t)$. However, a defaulting firm can be caught with probability γ, in which case its assets will be liquidated and the firms will cease to operate. If a firm is not caught, it continues to retain access to credit in future periods.[10]

Formally, a firm defaults if

$$y_t + (1-\gamma)E_t m_{t,t+1} V_{t+1} > E_t m_{t,t+1} V_{t+1},$$

where $m_{t,t+1}$ is the households' stochastic discount factor, and V_{t+1} is the firm's future value defined as the net present value of future dividends.

Because shocks realize at the beginning of period, there is no intraperiod uncertainty.

[10] Assuming that in the case of being caught a firm would also loose its revenues does not quantitatively alter our results.

Thus we can write the following incentive constraint that deters default in equilibrium,

$$\gamma E_t m_{t,t+1} V_{t+1} \geq y_t. \tag{1.2}$$

The left hand side of the constraint is equal to γ times firms' market value and decreases with the amount of intertemporal debt $b_{t,t+1}$. Whereas the right hand side is equal to the end-of-period revenues y_t which are equal to firms' intra-period loan. Hence, the incentive constraint in eq. (1.2) is effectively limiting both types of firms' debt. Importantly, in deciding between short and long-term debt, firms understand that an increase in b_{t+1} tightens their borrowing constraint as it limits their future ability to distribute dividends, but they do not internalize the effects that a change in production have on their market value through movements in the discount factor $m_{t,t+1}$. This type of externality will turn out to be crucial to generate both amplification and boom-bust phenomena.

The problem of the individual firm can be written recursively as

$$V_t = \max_{d_t, n_t, b_{t+1}} \left\{ d_t + E_t \Big[m_{t,t+1} V_{t+1} \Big] \right\} \tag{1.3}$$

subject to (1.1) and (1.2).

Firm's first order conditions are

$$(1 + \mu_t \gamma) R_t E_t \left[m_{t,t+1} \frac{\phi'(d_t)}{\phi'(d_{t+1})} \right] = 1 \tag{1.4}$$

$$\frac{w_t}{1 - \mu_t \phi'(d_t)} = (1 - \theta) \frac{y_t}{n_t} \tag{1.5}$$

where μ_t is the Lagrange multiplier associated to the incentive constraint. Equation (1.4) is the first order condition of new intertemporal debt b_{t+1}, and captures the fact that the marginal cost of debt increases with $\mu_t$and with the effective firm's discount factor defined as the household's discount factor, $m_{t,t+1}$ times the expected decrease in the cost of dividends. The first order condition of labor input (1.5) shows that financial frictions introduce a time varying labor wedge. When debt limits are looser the labor wedge declines, that is μ_t decreases, so that firms borrow more intra-period and the labor demand increases.

Furthermore, looser credit constraints also increase the intertemporal loan. To see this, combine the budget constraint of the firms with the optimality condition for labor:

$$\frac{b_{t+1}/R_t - b_t}{y_t} = \frac{\phi(d_t)}{y_t} - (1-\theta)\mu_t \phi'(d_t) - \theta.$$

As credit market relaxes, that is μ_t decreases, for a given dividend to output ratio, the intertemporal debt rises.

1.3.2 Households sector and general equilibrium

There is a continuum of homogeneous utility-maximizer households. Households are the owners of firms. They hold equity shares and noncontingent bonds issued by firms. Households' instantaneous utility function is

$$U(c_t, n_t) = \frac{c_t^{1-\omega} - 1}{1-\omega} + \alpha \log(1 - n_t).$$

The household's budget constraint is

$$c_t + s_{t+1}p_t + \frac{b_{t+1}}{1+r_t} = w_t n_t + b_t + s_t(d_t + p_t) - T_t \tag{1.6}$$

where s_t is the equity shares and p_t is the market price of shares. The government finances the tax benefits to firms through lump-sum taxes equal to $T_t = B_{t+1}/[1 + r_t(1 - \tau)] - B_{t+1}/(1 + r_t)$. The first order conditions with respect to n_t,b_{t+1}, and s_t are

$$w_t = -\frac{U_n(c_t, n_t)}{U_c(c_t, n_t)} \tag{1.7}$$

$$U_c(c_t, n_t) = \beta(1 + r_t)E_t U_c(c_{t+1}, n_{t+1}) \tag{1.8}$$

$$p_t = \beta E_t \left\{ \frac{U_c(c_{t+1}, n_{t+1})}{U_c(c_t, n_t)}(d_{t+1} + p_{t+1}) \right\} \tag{1.9}$$

Given the aggregate states $\mathbf{s}$, that are productivity z and aggregate bonds B we can define the general equilibrium as follows:

Definition: *A recursive competitive equilibrium is defined as a set of functions for (i) households' policies $c^h(\boldsymbol{s}, b)$, $n^h(\boldsymbol{s}, b)$ and $b^h(\boldsymbol{s}, b)$; (ii) firms' policies $d(\boldsymbol{s}, b)$, $n(\boldsymbol{s}, b)$, and $b(\boldsymbol{s}, b)$; (iii) firms' value $V(\boldsymbol{s}, b)$; (iv) aggregate prices $w(\boldsymbol{s})$, $r(\boldsymbol{s})$, and $m(\boldsymbol{s}', \boldsymbol{s})$; (v) law of motion for the aggregate states $\boldsymbol{s}' = \psi(\boldsymbol{s})$. Such that: (i) household's policies satisfy conditions (1.7) and (1.8); (ii) firm's policies are optimal and $V(\boldsymbol{s}, b)$ satisfies the Bellman's equation (1.3); (iii) the wage and the interest rate clear the labor and bond markets; (iv) the law of motion $\psi(\boldsymbol{s})$ is consistent with individual decisions and stochastic processes for productivity.*

1.3.3 Inspecting the mechanism

The key externality in the model is that firms do not fully internalize the effects of their production decisions on their market value. In particular, while they understand that a higher level of debt reduces their market value because it limits the ability to distribute future dividends, they do not internalize the feedback loop between output and their market value. Absent of adjustment cost of dividends, *i.e.* $\kappa = 0$, credit market amplification depends upon the elasticity of firms' production to the households' stochastic discount factor. This elasticity is equal to

$$\frac{\partial log(y_t)}{\partial log(m_{t,t+1})} = \frac{\beta\tau}{\gamma(1-\mu)(1-\tau+\tau\beta)^2}\left[\frac{(1-n)(1-\theta)}{(\omega-1)(1-n)(1-\theta)+1}\right] \equiv \xi,$$

where $\mu = \tau(1-\beta)/\gamma(1-\tau+\tau\beta)$.

If credit market frictions are severe, that is the probability of being excluded from financial market γ is low or the tax advantage on debt τ is high, firms are more responsive to changes in their continuation value reflected by changes in the stochastic discount factor. Sufficiently high values of ξ give rise to self-fulfilling equilibria. Suppose lenders and borrowers are optimistic regarding firms' market value, this relaxes the financial constraint and implies an increase in the credit supply. As a consequence, production and households' labor income increase which raise firms' market value through an increase in the stochastic discount factor $m_{t,t+1}$ validating the initial shift in expectations.

Formally, take a first order approximation around the steady state, aggregate output can be expressed as

$$\hat{y}_t = \frac{\omega\xi}{\omega\xi - 1} E_t \hat{y}_{t+1} - \frac{1}{\zeta(\omega\xi - 1)} \hat{z}_t \tag{1.10}$$

where $\zeta \equiv (\omega - 1)(1 - n)(1 - \theta) + 1$.

When $\omega\xi > 1/2$, current aggregate output is a convex function of future output which is sufficient to generate indeterminacy.

Note that the impact of technology shocks on aggregate output is ambiguous. By increasing end of period revenues, a positive technology shock raises firm's incentives to divert funds thereby increasing the right-end-side of the incentive constraint in eq. (1.2). Whether firm's market value increases more than firm's revenue depends upon firm's willingness to distribute dividends. We find that for plausible parametrizations, the Lagrange multiplier μ_t *increases* in response to a positive technology shock.

Thus financial constraints amplify shifts in expectations while they dampen the response to technology shocks. Yet, why do boom-bust episodes occur? Theorem 1 below lists the necessary conditions under which boom-bust fluctuations may obtain in response to perturbations from the economy's steady state.

Theorem 1 *Boom-bust phenomena obtain only if*

i. The equilibrium is indeterminate.

ii. Adjustment costs are non zero, that is $\kappa > 0$.

Proof is relegated in Appendix A.8.

Condition i. states that if the credit market amplification channel is strong enough, so that indeterminacy obtains, then the economy can also be subject to oscillatory dynamics.[11] The intuition is that after an initial expansion, firms have accumulated large amount of debt which limits their ability to borrow and produce. As firms decrease production they do not internalize the adverse effects on their market value. The stronger are the effects of this externality the larger is the drop in current production. The reason why adjustment cost of dividends is necessary to obtain cycles is more subtle. Besides the static amplification mechanism described above, the model displays dynamic substitutability between current and future production generated by movements in firms' net worth. An increase in new debt brings about higher current production but it decreases future firms' net worth which negatively affects the subsequent level of production. Absent dividend adjustment costs, firms with a high level of outstanding debt would finance production by decreasing the amount of distributed dividends, therefore limiting the impact that changes of net worth on their production decisions, thus preventing the large accumulation of debt after the expansion to generate a recession.

1.3.4 Parametrization and theoretical impulse responses

The sunspot shock is defined as an i.i.d. expectation error of firm's value that is not correlated with fundamentals

[11]This property is not specific to the environment described here. Gu et al. (2013) discuss the link between indeterminacy and cycles in the context of financial frictions of different forms.

$$\widehat{V}_t - E_{t-1}\widehat{V}_t = u_t$$

where $u_t = \epsilon_{s,t} + \psi_z \epsilon_{z,t}$.

The terms $\varepsilon_{s,t}$ and $\epsilon_{z,t}$ are respectively the sunspot shock and the technology shock.[12] The natural logarithm of technology is assumed to follow an AR(1) process as

$$\widehat{z}_t = \rho_{z,t}\widehat{z}_{t-1} + \epsilon_{z,t}.$$

We calibrate the model to a quarterly frequency consistent with the frequency of the data. We set β in order to match a 3% annual interest yield on bonds. Following Jermann and Quadrini (2012) tax shield τ and capital's share of income θ are set equal to 0.35 and 0.36, respectively. With the aim of emphasizing the difference between the two shocks, we set the inverse of households' intertemporal elasticity of substitution ω to 1.2, the probability of being caught in case of default γ to 0.1 and the degree of adjustment cost to dividends κ to 2.3. The parameter ρ_z governs the persistence of the technology process and is set equal to 0.93 consistent with the law of motion of detrended TFP estimated in the data. We assume the expectation error u_t and the technology shock to be uncorrelated, so that ψ_z is equal to zero.[13]

[12] Note that inserting the sunspot on output would not alter our results. It is easy to show that

$$\widehat{V}_t - E_{t-1}\widehat{V}_t = \omega(\widehat{y}_t - E_{t-1}\widehat{y}_t).$$

[13] Note that ψ_z equal zero implies a zero-impact response of output and firm's value after a technology shock. While this is an implausible restriction that will be relaxed in the quantitative exercise, it allows to generate a starker difference between the dynamics induced by the two shocks.

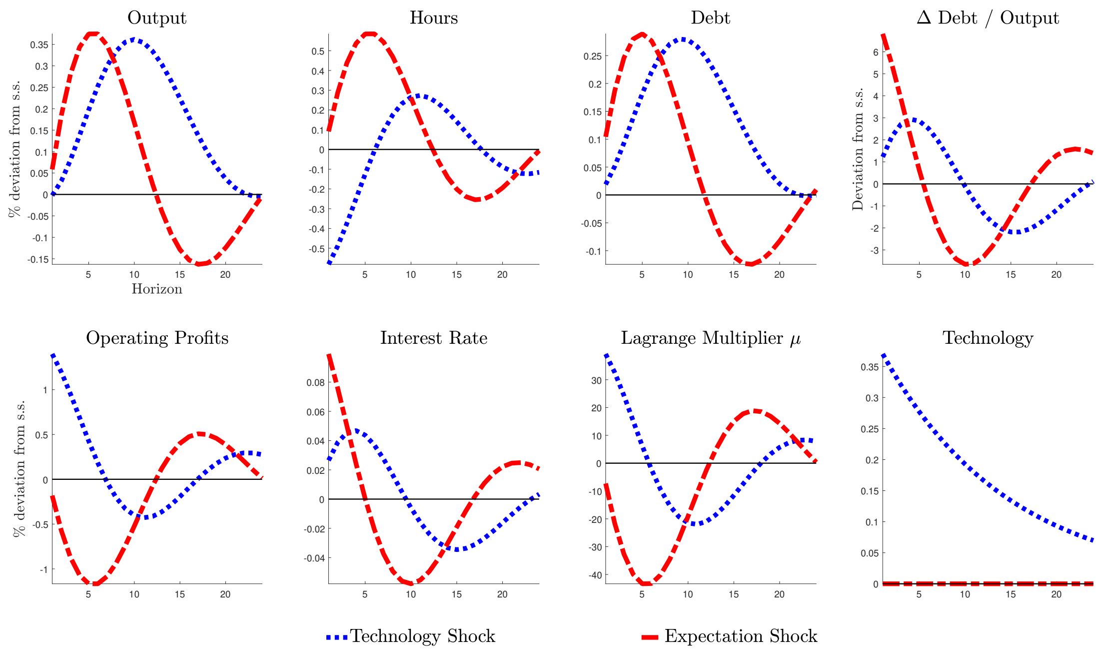


Figure 1.4: Model impulse responses to a technology shock and to sunspot shock

Figure 1.4 shows the theoretical impulse responses of the model to a sunspot shock and to a technology shock. In response to the sunspot shock the economy experiences an initial boom characterized by an increase output, consumption and hours. The associated increase in debt has two effects. On the one hand, it reflects an increase in households' savings which increases the supply of credit generating a decrease in the real rate and an increase in firms' market value. On the other hand, larger outstanding debt hinders firms' ability to pay current and future dividends which deteriorates their market value. Which of these two forces prevails depends upon the level of firms' profitability. As production increases firms' profitability falls so that firms' market value decreases, the financial constraint tightens and output starts declining. During the contraction phase, households are less willing to lend which results in an increase in the real rate, a decrease in firm's value and a further tightening of the financial market. This negative vicious circle reinforces as households' savings decline, ultimately bringing about a recession. Importantly, even though agents know about the incoming recession their actions magnifies the decline in output.

A positive technology shock generates hump-shaped dynamics in all the main macroeconomic variables. By increasing incentives to divert funds, a positive technology shock tightens the financial constraint which dampens the impact response of output. Importantly, the response of debt and output is comparable to the ones after a sunspot shock, suggesting that looking at measures of firms' indebtedness such as the debt to GDP ratio may not be the best predictor of a crisis.

Importantly, expectation-driven fluctuations arise also in an economy where fundamentals, that is technology, preferences, or government policies, do not change and this is common knowledge. This distinguishes them from noise shocks arising from *ex post* erroneous beliefs on future changes of technology. Bearing this distinction in mind, in the next section, we estimate expectation shocks unrelated to fundamentals and to rational expectations of fundamentals. We find that these shocks generate boom-bust dynamics consistent with the quantitative prediction of an extended version of the model.

1.4 Identifying sunspot shocks using survey data

In this section we estimate the sunspot shock as a "pure" sentiment shock, that is a shock that reflects a change in expectations disconnected from changes in expectations on future

TFP and realizations of TFP. To this end, we use quarterly one-year-ahead expectations on a number of key macroeconomic variables formed by both professional forecasters and households. We proceed in three steps.

The Survey of Professional Forecasters and the Survey of Consumer Expectations include expectation data on a number of variables, such as future real GDP growth, investment, and consumption. Our theory does not point at a particular variable, rather expectation shocks should be reflected into a change of expectations common across all variables in the surveys that capture information upon expected future business conditions. Therefore, as a first step, we construct an expectation indicator $\hat{S}_t$ from the first principal component of all the relevant available expectation data. The sample includes seven quarterly variables from 1982:Q2 to 2018:Q4.

Second, we regress the indicator $\hat{S}_t$ on a battery of controls in order to capture variations in expectations that are "extrinsic", that is, exogenous to fundamentals and to changes in expectations on future fundamentals. Formally, let the process of detrendend TFP be represented by the following news representation

$$\log(TFP)_t = A(L)\log(TFP)_{t-1} + \varepsilon_t^z + \sum_{k=1}^{\infty} \varepsilon_{t-k}^k$$

where ε_{t-k}^k is a news shock on TFP k-period ahead which is part of time t agents' information set, and ε_t^z is the surprise shock of technology. Let S_t^K be the indicator that summarizes revision of agents expectations on the economic activity K-period ahead. We assume that these revisions depend upon current technology shocks, expectations on future technology, and expectation shocks. Specifically,

$$S_t^K = \lambda_0 \log TFP_t + \sum_{k=1}^{K} \alpha_k \varepsilon_t^k + \varepsilon_t^s$$

where expectations on future technology are a linear combination of news upon technology up to K horizons. Hence, in order to identify *extrinsic* expectation shocks one needs to cleanse changes in expectations, proxied by $\hat{S}_t$, from the realized level of TFP and expectations about future TFP up to the horizon K. In other words, we want the estimated expectation shock to satisfy two conditions: (i) the estimated shock must be uncorrelated with future TFP realizations; (ii) the shock has to be uncorrelated with noise shocks, defined as ex-post wrong beliefs on future TFP. [14]

We proxy expectations on future TFP with TFP news shocks identified as in Barsky and Sims (2011a). However, this controlling set may no be large enough to satisfy the two conditions above. To overcome this issue we add two additional set of controls. First, we control for future realizations of TFP so as to guarantee that the estimated shock has no impact on future TFP. Second, as shown by Chahrour and Jurado (2018), one can recover noise shocks by adding future news and realizations of TFP to the econometrician's information set. Thus, we further control for future realizations of the identified news shock.

[14] As shown by Beaudry and Portier (2004) noise shocks in the form of ex-post wrong beliefs on future TFP can give rise to Pigouvian cycles and therefore are a competing candidate to the explanation of the reduced form evidence presented in Section 1.1. However, we find that controlling for this particular type of beliefs has small quantitative changes on the variance explained by the expectation shock, suggesting that noise shocks play only a minor role in shaping expectations.

Specifically, expectation shocks are estimated from the following equation:

$$\hat{\varepsilon}_t^s = \hat{S}_t - \sum_{k=0}^{\bar{k}} \widehat{\lambda}_k TFP_{t+k} - \sum_{k=0}^{\bar{k}} \widehat{\alpha}_k \varepsilon_t^{BS} - \mathbf{X}_t \widehat{\beta}$$

where ε_t^{BS} is the news shock estimated using the procedure in Barsky and Sims (2011a), and $\mathbf{X}_t$ is a vector of additional control variables, including past realizations of TFP and news, other shocks to fundamentals such as monetary policy and fiscal shocks, and past values of the first two principal components from a large data set of U.S. aggregate variables. Interestingly, even after controlling for virtually all available sources of fundamental fluctuations, estimated expectation shocks explain approximately half of the changes in the expectation indicator $\hat{S}_t$.

In the last step, we estimate the impulse response to an expectation shock using Local Projections as in Jordà (2005). Specifically, for each variable of interest Y, we run the following series of regressions

$$Y_{t+h} = \theta^h \hat{\varepsilon}_t^s + \sum_{j=1}^{J} \left[\delta_j \hat{\varepsilon}_{t-j}^s + \lambda_j Y_{t-j} + \mathbf{PC}_{t-j} \Gamma_j \right] + \nu_{t+h} \quad \text{for } h = 0, 1, \ldots, H \tag{1.11}$$

where θ^h is the response of Y to an expectation shock after h periods, and PC is a vector including the first two principal component from a set of U.S. aggregate variables. We use four lags, that is $J = 4$, in the baseline specification.

Figure 1.5 shows the responses of real GDP, real investment, and the change of non-financial corporate debt divided by real GDP to a one standard deviation expectation shock. Real GDP, investment and debt flow exhibit significant oscillatory dynamics. In particular, after a positive expectation shock, the economy enters an expansion followed by a recession after about two years. Importantly, the conditional spectral densities exhibit a peak associated to periodicities of 8 to 10 years, in line with the reduced form evidence presented earlier. Table A.1 in Appendix A.7 reports the p-values for the test of a local peak in the spectral density implied by expectation shocks. The null hypothesis of absence of a local peak is rejected for all variables, with the exception of TFP.

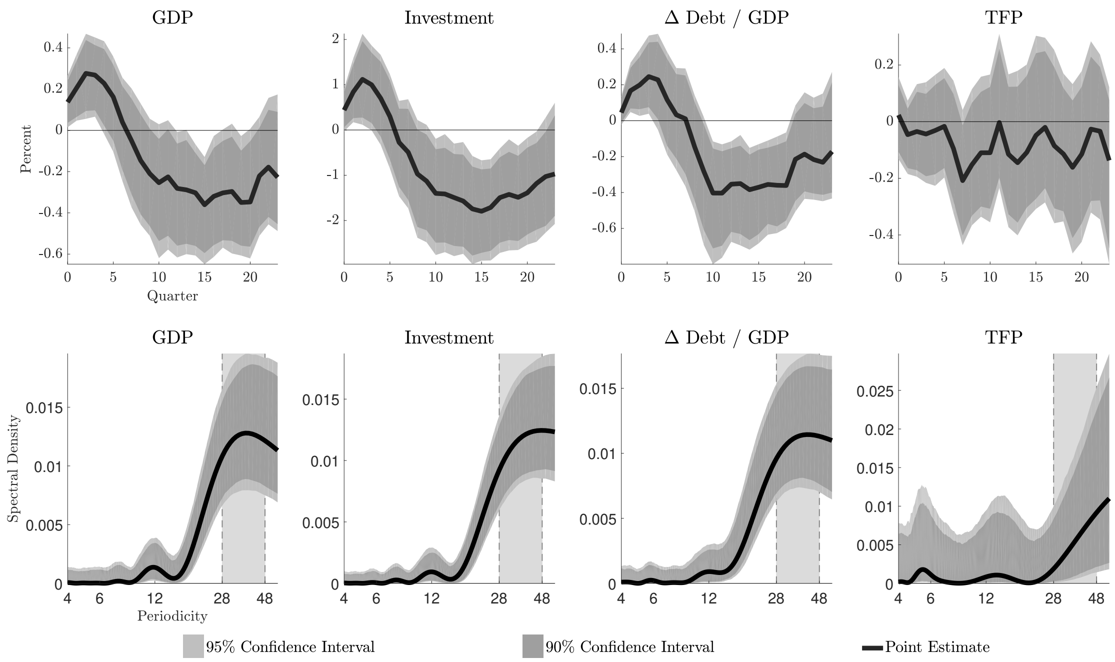


Figure 1.5: Impulse responses and conditional spectral densities to an expectation shock

Note: Expectation shocks are estimated as the innovations in S_t orthogonal to present, past, and future realization of TFP and expectations on TFP. Impulse responses (top panel) are estimated using local projections method. Confidence intervals are computed using the block-bootstrap method described in Kilian and Kim (2011). Conditional spectral densities (bottom panel) are computed from the Fourier transform of the estimated MA.

1.4.1 Robustness checks

In this section we show that the results in Figure 1.5 are robust to different detrending techniques, additional controls, and the expectation variables used to construct the indicator S_t. Given that our endogenous variables are non-stationary, in the baseline specification we detrend the variables using a Band-Pass filter which excludes periodicities above 100 quarters. In order to argue that the oscillatory dynamics implied by an expectation shock is not specific to the detrending technique, in Figure 1.6 we show robustness checks where endogenous variables are detrended using (*i*) first differences (and the cumulated), (*ii*) linear time trend, (*iii*) quadratic time trend, and (*iv*) Hodrick-Prescott filter. Results are in line with the baseline specification and most of the estimates lie between the confidence intervals of the main specification.

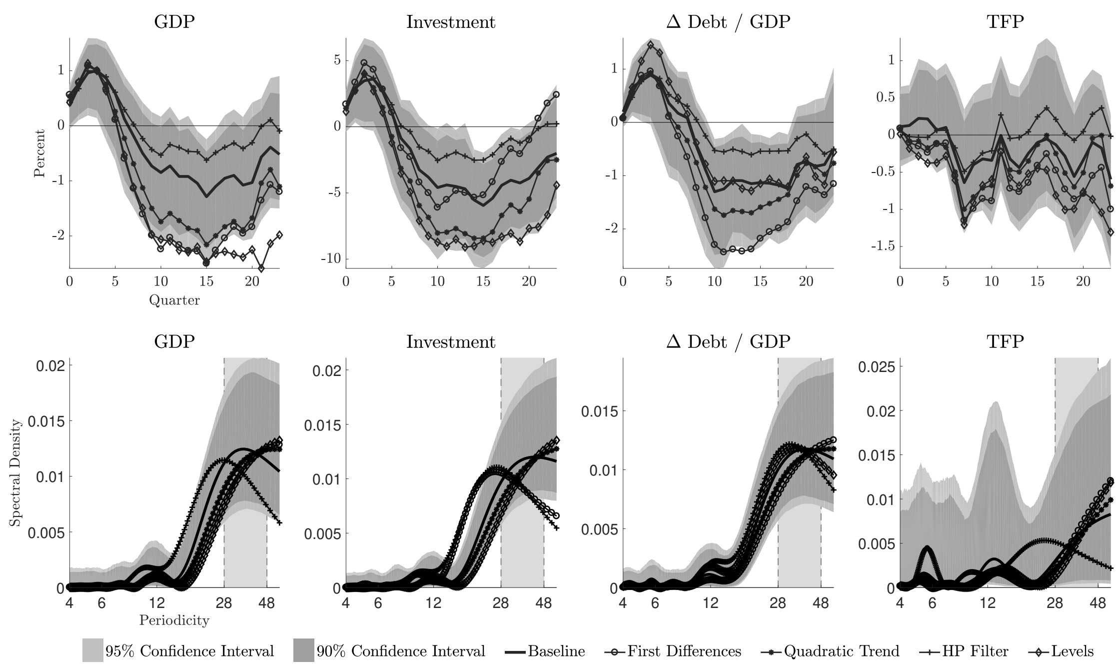


Figure 1.6: Impulse responses and conditional spectral densities to an expectation shock

Note: Point estimates (continuous line) are from the baseline specification presented in Figure 1.5. The figure shows the robustness of the point estimate to various detrending techniques.

Figure 1.7 reports results for four additional variations of the baseline specification. First, we increase the number of lags and the number of principal components in the regression equation of the expectation shock. Second, we control for the present and the past of other shocks to fundamentals such as oil shocks, fiscal shocks, military spending news shocks and monetary policy shocks. Third, we check whether results are sensitive to the choice of the indicator for the revisions of expectations. Specifically, we use only revisions on one-year-ahead output growth from the SPF and find results that are not significantly different from the baseline. Finally, we check that results are robust to the number of lags and principal components used in the LP.

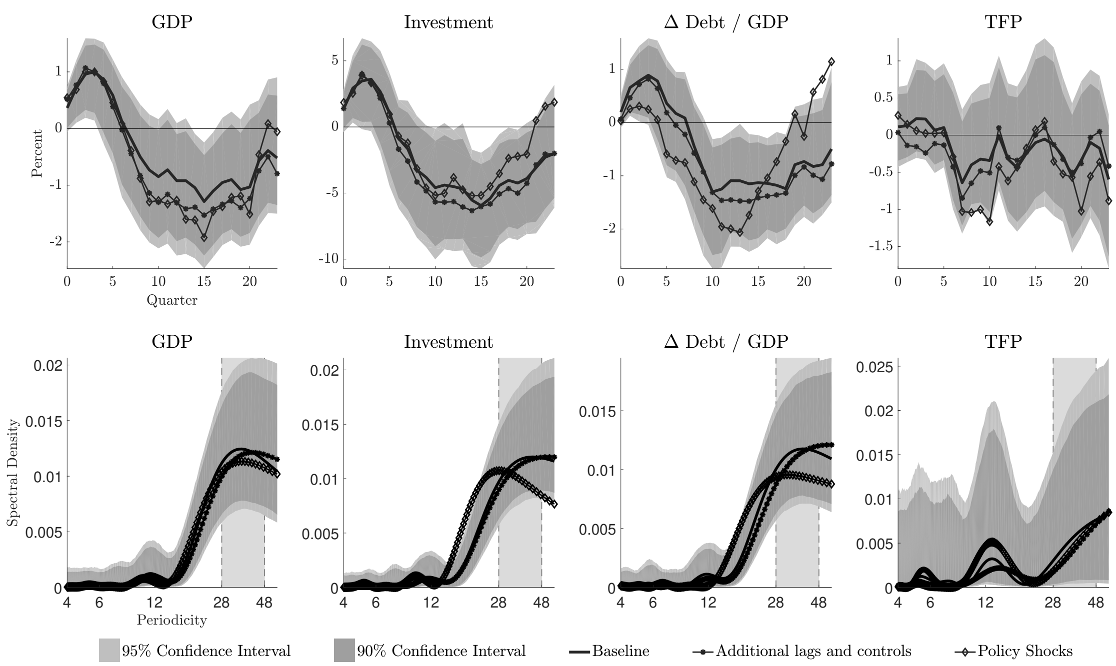


Figure 1.7: Impulse responses and conditional spectral densities to an expectation shock

Note: Point estimates (continuous line) are from the baseline specification presented in Figure 1.5. The figure shows the robustness of the point estimate to various controls (see text).

1.5 Model with capital and external consumption habit

In this section we augment the model with variable capital, investment-adjustment costs and external consumption habit. The equilibrium equations of the extended model are:

$$w_t U_c(c_t, c_{t-1}, n_t) = -U_n(c_t, c_{t-1}, n_t) \tag{1.12}$$

$$E_t[m_{t,t+1}(R_t - \tau)] = 1 - \tau \tag{1.13}$$

$$w_t n_t + b_t - \frac{b_{t+1}}{R_t} + d_t = c_t \tag{1.14}$$

$$\big[1 - \mu_t \phi'(d_t)\big] F_n(z_t, k_t, n_t) = w_t \tag{1.15}$$

$$k_{t+1} = (1-\delta)k_t + \left[\frac{\varsigma_1}{1-\nu}\left(\frac{i_t}{k_t}\right)^{1-\nu} + \varsigma_2\right] k_t \tag{1.16}$$

$$\begin{aligned} E_t\Bigg\{ m_{t,t+1}\frac{\phi'(d_t)}{\phi'(d_{t+1})}(1+\mu_t\gamma)\Bigg\{ \big(1-\phi'(d_{t+1})\mu_{t+1}\big)F_k(z_{t+1}, k_{t+1}, n_{t+1}) + \\ + \frac{1}{\varsigma_1}\left(\frac{i_{t+1}}{k_{t+1}}\right)^{\nu}\left[1-\delta+\frac{\varsigma_1\nu}{1-\nu}\left(\frac{i_{t+1}}{k_{t+1}}\right)^{1-\nu} + \varsigma_2\right]\Bigg\}\Bigg\} = \frac{1}{\varsigma_1}\left(\frac{i_t}{k_{t-1}}\right)^{\nu} + E_t\big[m_{t,t+1}\phi'(d_t)\mu_t\gamma\big] \end{aligned} \tag{1.17}$$

$$(1+\mu_t\gamma)E_t\left[m_{t,t+1}\frac{\phi'(d_t)}{\phi'(d_{t+1})}R_t\right] = 1 \tag{1.18}$$

$$y_t - w_t n_t - b_t + \frac{b_{t+1}}{R_t} - i_t = \phi_t(d_t) \tag{1.19}$$

$$\gamma E_t\big[m_{t,t+1}V_{t+1}\big] = y_t \tag{1.20}$$

where $y_t = F(z_t, k_t, n_t) = z_t k_t^{\theta} n_t^{1-\theta}$ and $\phi(d_t) = d_t + \kappa(d_t - \bar{d})^2$. Moreover, the stochastic discount factor is $m_{t,t+1} \equiv \beta(U_{c,t+1}/U_{c,t})$ and value of the firm is defined as $V_t = d_t + E_t\big[m_{t,t+1}V_{t+1}\big]$. Finally, $U_c(c_t, c_{t-1}, n_t) = (c_t - \iota c_{t-1})^{-\omega}$ and $U_n(c_t, c_{t-1}, n_t) = -\alpha(1 - n_t)^{-\omega_2}$.

1.5.1 Calibration and impulse response matching

Following Christiano et al. (2005) we divide the model parameters in two different groups. The first group is calibrated using unconditional moments or results from previous studies while the remaining parameters are estimated via impulse response matching. In both cases, the model is calibrate at a quarterly frequency. In the first group, the discount factor β, the capital share of income θ, and tax shield τ have the same values presented in section 1.3. The multiplicative parameter which governs the utility of leisure α is chosen such that the steady state value of n is equal to 0.3. Parameters ς_1 and ς_2 (capital-adjustment costs) are set such that in the steady state the depreciation rate is equal to $\delta = 0.025$ and the steady state Tobin's q is equal to one. Parameter ψ_z, which captures the correlation between technology shocks and the forecast error on firms' value, is set in order to match the impact of a 1% technology shock on real GDP. Moreover, the parameter γ, which governs the tightness of the incentive constraint, is set in order to match an empirical average debt-to-output ratio of 3.36. Finally, κ is calibrated in order to have a model standard deviation of equity payout over output equal to the empirical standard deviation.

The second group includes the vector of parameters $\Sigma = (\rho_z, \omega, \iota, \nu)$: the persistence of technology process, ρ_z; the inverse of households' intertemporal elasticity of substitution, ω; the external consumption habit parameter, ι; the degree of capital adjustment cost, ν. We choose Σ in order to minimizes the following object

$$J = \min_{\Sigma}[\hat{\Psi} - \Psi(\Sigma)]' V^{-1} [\hat{\Psi} - \Psi(\Sigma)]$$

where $\hat{\Psi}$ denotes the empirical impulse responses of GDP, Consumption, hours worked and TFP to both technology and expectation shocks, and $\Psi(\Sigma)$ is the model-implied counterpart of $\hat{\Psi}$. Finally, V is a diagonal matrix which gives different weights to the target estimates. Table 1.1 reports the parameter values of the model.

Parameter	**Interpretation**	**Value**	**Target**
α	Disutility of labor	8.785	Hours in steady state = 0.3
β	Discount factor	0.99	Annual bond yield = 3%
τ	Tax shield	0.35	Jermann and Quadrini (2012)
θ	Capital share	0.36	Standard
δ	Capital depreciation	0.025	Standard
ς_1	Capital adj. cost (1)	δ^{ν}	Depreciation rate = δ
ς_2	Capital adj. cost (2)	$\delta - \delta/(1-\nu)$	Tobin's $q = 1$
ψ_z	Corr tech and exp error	0.24	Impact of tech. shock on GDP
γ	IC parameter	0.12	$b/Y = 3.36$
κ	Dividend cost	3.01	$std(d/Y) = 0.024$
ρ_z	Technology persistence	0.93	IRF matching estimation
ω	CRRA consumption	1.25	
ι	Consumption habit	0.45	
ν	Capital adj. cost	0.55	

Table 1.1: Model's parameter values.

1.5.2 Model performance

Figures 1.8 and 1.9 plot the theoretical impulse response of the model against their empirical counterparts. The model does a good job in reproducing the empirical impulses to both shocks. In particular, we estimate the model consistent measure of labor wedge and find that the responses are in line with the predictions of the model.

Figure 1.10 shows the empirical conditional spectral densities against their model counterpart. The theoretical spectral densities implied by the model are within the range of the confidence bands of the empirical ones.

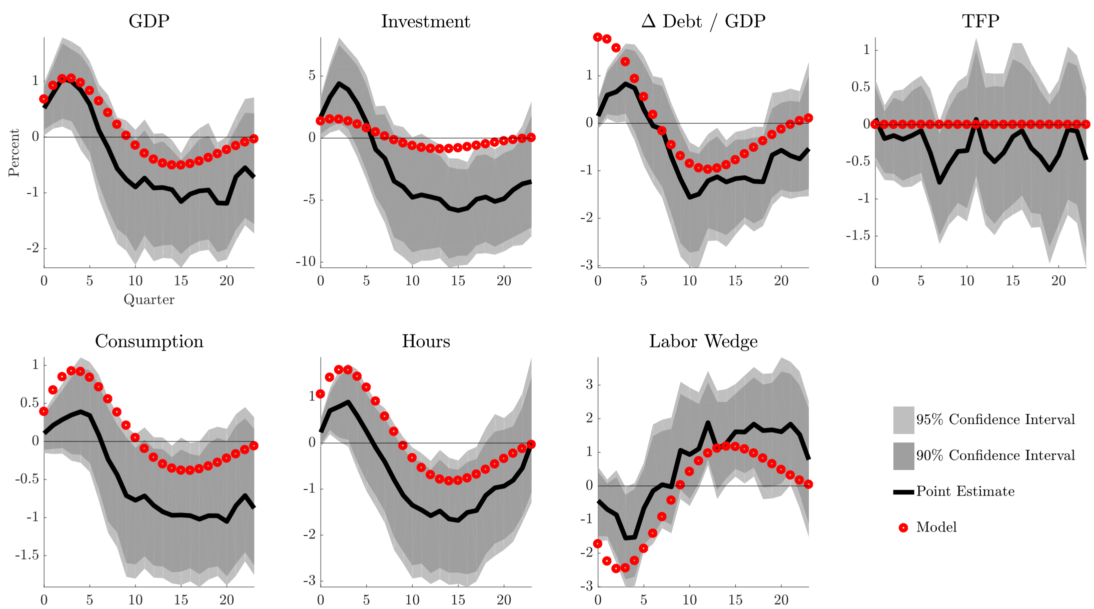


Figure 1.8: Model vs empirical IRFs to an expectation shock

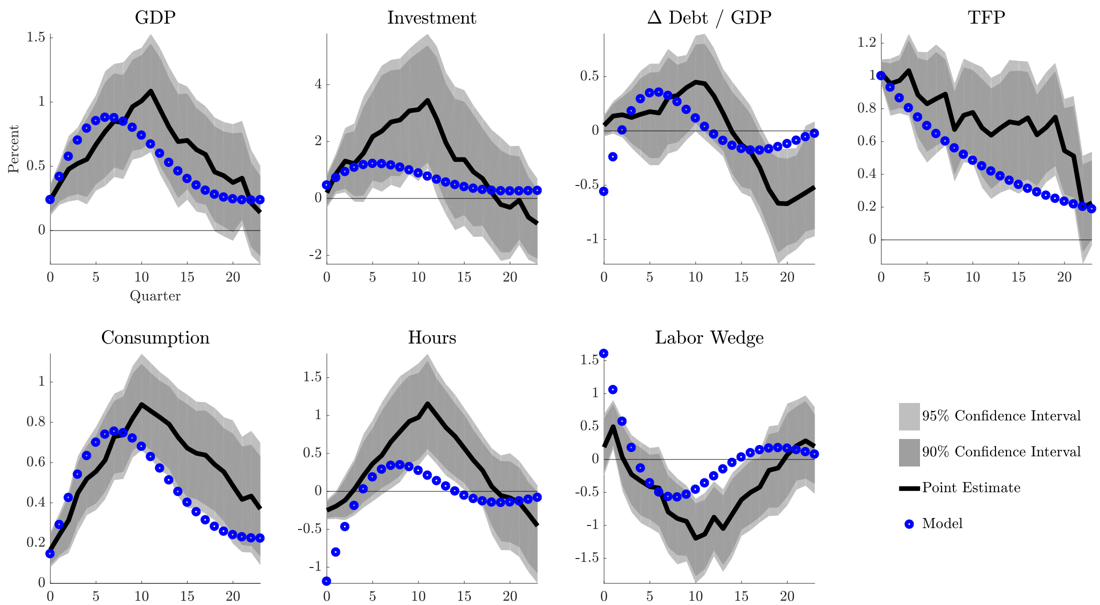


Figure 1.9: Model vs empirical IRFs to a technology shock

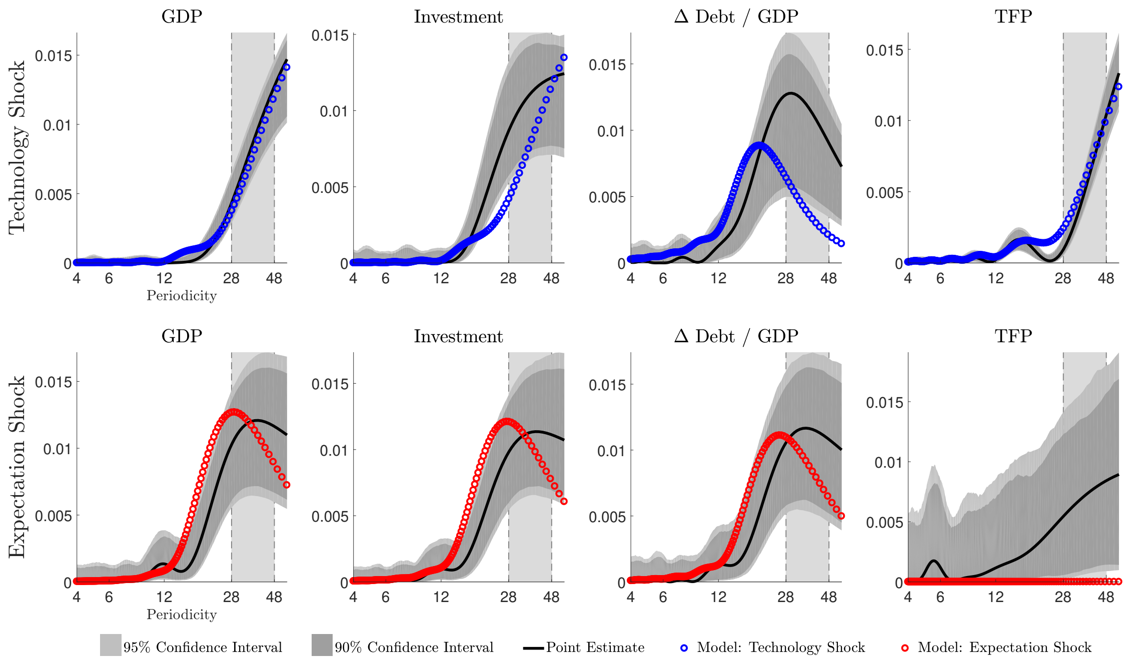


Figure 1.10: Model vs empirical spectral densities conditional on shocks

As a last validation exercise of the model, we simulate data and reproduce the results on the probability of recession presented in Figure 1.2. Figure 1.11 shows that the model can replicate the empirical probability of recession conditional on a previous expansion.

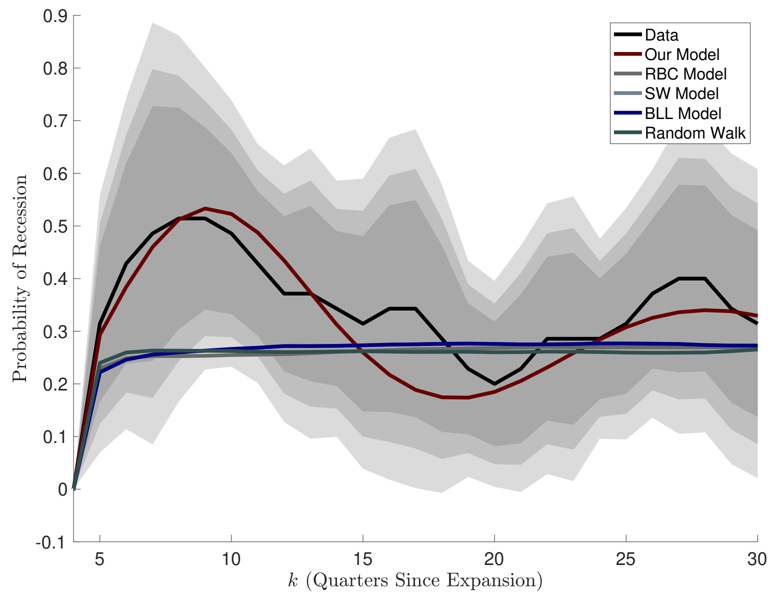


Figure 1.11: The model explains the dynamics of the recession probability

Note: Probability of recession in a two-quarter window after k quarters since expansion. Confidence intervals are 68%, 80%,and 90% (shaded areas) around the point estimate (solid black line).

1.6 Conclusion

We provide a simple synthesis of two major approaches to modeling business cycles. Under the first approach business cycles are driven by exogenous shocks that push the economy temporarily away from the long-run steady-state or balanced growth path. The second approach proposes models in which the economy experiences endogenous fluctuations even in the absence of fundamental shocks. However, both types of models fail to provide a unified explanation of the unconditional and conditional moments of the data. In the data, shocks to economic fundamentals induce dynamics that are consistent with the first view.

But unconditional moments and results from expectation shocks, suggest to write models consistent with the inherent instability class. Taken together, our findings speak in favor of a theory in which both views coexist. Thus, we provide a model that embeds a strong financial amplification channel which generates boom-bust dynamics in response to i.i.d. expectation shocks. Consistent with the data, the financial amplification channel barely contributes to the propagation of technology shocks which exhibit no systematic relation between expansions and recessions. In sum, a sizeable part of economic recessions is due to preceding expansions. More importantly, those expansions that are not generated by a change in fundamentals are more likely to end in recessions. As a consequence, policy makers should intervene more decisively during expectation-driven expansions than during fundamental-driven expansions. Characterizing the optimal policy in light of our findings is part of our future endeavors.

Chapter 2

Shocks and Exchange Rates in Small Open Economies

2.1 Introduction

Exchange rates are arguably the most important price for small open economies (SOEs), but the sources of their fluctuations are still far from understood. Partly because of scant disciplining evidence, the drivers of exchange rate dynamics largely differ across classes of open economy models. This paper studies the properties of domestic and external shocks on SOEs exchange rates, presents a new set of exchange rate facts, and explores their implications for open economy models.

We begin by showing that it is possible to separately identify domestic and external shocks in SOEs using minimal assumptions that hold in any class of SOE models. We observe that shocks originating from within a *small* economy should not influence world variables at any horizon, while external (or global) shocks should affect world variables at least at some horizon. In the context of vector autoregressions (VARs), we thus identify external shocks as those that explain all of contemporaneous and expected future movements in external variables. To do so, we apply a methodology developed by Uhlig (2003) to extract the exogenous shocks that explain as much as possible of the forecast error variance of an external variable in a VAR. Our approach then requires domestic shocks to be orthogonal to all external disturbances. We implement this methodology on monthly data for a large number of SOEs, study the properties of these shocks, and interpret them by analyzing the dynamic comovement that they imply.

Our first empirical finding is that external and domestic shocks display different patterns of deviations from uncovered interest parity (UIP). External shocks are associated with large and significant UIP deviations, and account for about 80% of all fluctuations in expected currency excess returns. To the contrary, domestic shocks generate exchange rate dynamics that are largely consistent with UIP, and do not substantially contribute to the variation in expected currency excess returns. The differences in conditional UIP deviations are such that domestic and external shocks display opposite comovement patterns between interest rate differentials and exchange rates. These facts indicate that country-specific UIP shocks are not a satisfactory representation of the data, and understanding UIP deviations requires inspecting the transmission channels of external disturbances.

Our second empirical finding is that one external shock drives a large fraction of fluctuations in exchange rates and expected currency excess returns, and it strongly comoves with global risk aversion and U.S. macroeconomic variables. Our approach does not require that a single shock accounts for a large fraction of external variation or that any shocks have an appealing interpretation. Yet, when applying our decomposition, we find that one single external shock can account for at least 2/3 of the external variation in exchange rates and currency excess returns. Moreover, we find that this shock is strongly correlated with innovations in the VIX – a common proxy of global risk aversion – and is associated with significant U.S. macroeconomic fluctuations. This external shock is characterized by the following comovement. When global risk aversion is low, U.S. output, inflation, and the Federal Funds rate are all significantly above their steady state. Across SOEs, interest rate declines in the short run, and their currencies appreciate relative to the U.S. dollar (with the exchange rate response being primarily shaped by the dynamic pattern of expected excess returns).[1] While closely related to the evidence of the "global financial cycle" (cf. Rey, 2013), the positive comovement among U.S. output, inflation and interest rates reveals that the bulk of external variation in exchange rates is not driven by U.S. monetary policy shocks.[2]

We show that an open economy model with segmented asset markets and global risk aversion shocks is consistent with the above findings. Building on a standard two-country

[1]This comovement implies that the external variation in exchange rates is not disconnected from U.S. macroeconomic dynamics.

[2]Relatedly, SOE models with only exogenous shocks to the external interest rate do not appear to be an adequate characterization of the data.

SOE framework with nominal rigidities (cf. Galí and Monacelli, 2005, and De Paoli, 2009), we assume that international financial markets are segmented, and financial traders, a subset of U.S. households, are averse to holding currency risk.[3] Besides a standard set of structural shocks, we introduce a "global risk aversion shock," modeled as an exogenous change in the level of risk aversion of U.S. households and financial traders, in the spirit of Gabaix and Maggiori (2015). In addition to the path of interest rate differentials, equilibrium exchange rates are determined by the level of traders' risk aversion and the external imbalance of the SOE. In our model, the net foreign asset (NFA) position to GDP of the SOE is the relevant measure of its external imbalance, and determines the amount of currency risk held by international traders. Importantly, the steady-state level of NFA/GDP of the SOE primarily governs its sensitivity to changes in global risk aversion. We solve our model around a non-zero steady-state NFA/GDP position, calibrated to -15% to reflect the median value in our empirical sample of SOEs. A negative steady-state NFA/GDP implies that U.S. financial traders are long in the SOE currency. Therefore, they demand a positive currency premium to hold the SOE currency risk, and this premium is increasing in the level of risk aversion.

Global risk aversion shocks lead to macroeconomic dynamics that are in line with our identified external variation, as well as the conditional patterns of expected excess returns.

[3] In this framework, economic developments in the large economy (the U.S.) affect the small economy, but not *vice versa*. This two-country SOE environment is thus consistent with our key empirical identification restrictions.

When global risk aversion declines, higher U.S. households' demand leads to an increase in U.S. output and inflation, as well as rising Federal Funds rate. Lower risk aversion induces traders to require lower excess returns on the SOE currency, bringing about a large currency appreciation in the SOE (despite an equilibrium decline in the interest rate differential).[4] Unlike global risk aversion shocks, other kinds of shocks have only mild effects on expected currency excess returns. In fact, any domestic shock that raises domestic interest rates – such as a domestic monetary policy contraction – leads to an impact appreciation of the domestic currency, with negligible deviations from UIP. Our parsimonious framework is therefore able to reproduce the conditional properties of UIP deviations that we documented empirically.

Our proposed model implies that the average NFA/GDP position of a country determines its exchange rate exposure to external shocks, and, as a result, its exchange rate properties. In our linearized model, the direct effect of a global risk aversion shock on the exchange rate solely depends upon a country's steady-state net foreign asset position. As a result, according to our model, countries with large net foreign debt should be more exposed to exogenous changes in global risk aversion relative to countries with a negligible net foreign position. Moreover, countries with large net foreign debt should experience larger exchange rate volatility and a positive comovement between the interest rate differential and their exchange rate.

[4] In response to this shock, domestic central banks cut their policy rate in the short run to avoid excessive fluctuations in consumer price inflation, in line with our empirical evidence.

We verify the key predictions of our model by resorting to the cross-country dimension of our data. Unlike countries with a small NFA/GDP position, we find that SOEs with large average net foreign debt (i) experience large exchange rate appreciations following expansionary external shocks, (ii) feature considerable exchange rate volatility, (iii) display a positive comovement between the interest rate differential and their exchange rate, and (iv) feature an exchange rate that is predominantly driven by external disturbances. In fact, the degree of external exposure in our data ranges from around 80% to around 20%, and is significantly correlated with a country's net foreign position. The untargeted moments in (i)-(iv) obtain naturally in our model when global risk aversion shocks are the predominant driver of external variations in exchange rates.[5]

Last, our analysis brings us to revisit recent puzzling evidence on the exchange rate response to domestic monetary policy shocks in SOEs. Our finding that the domestic variation in exchange rates is largely in line with UIP contrasts with recent evidence on the exchange rate effects of domestic monetary policy shocks. In particular, Hnatkovska et al. (2016) find that the domestic currency tends to depreciate after domestic monetary tightening in several SOEs, implying a significant UIP deviations from domestic shocks. This evidence, labeled "the exchange rate response puzzle," is primarily based on recursive identification schemes within the framework of VARs. The recursive identification strategy obtains as a special case of our proposed identification scheme, which allows us to understand

[5]Note that facts (i)-(iv) do not necessarily follow from Gourinchas and Rey's (2007) evidence that the level of external imbalances predict future exchange rates.

the nature of the differences in results. We show that VAR identification approaches based on recursive ordering are bound to commingle domestic and external shocks. In particular, we document that the structural shocks identified through recursive ordering and typically interpreted as "domestic monetary policy shocks" predict significant future movement in external variables. This feature emerges because domestic interest rates and exchange rates display strong "anticipated effects" – that is, they predict future external variables – and identification schemes based on contemporaneous restrictions do not account for all contemporaneous and expected variation of the external variables included in the VAR.[6,7] By identifying external shocks as those that explain movements in external variables *at any horizon*, our identification approach does not conflate shocks with different sources. We show that this misspecification problem is the source of "the exchange rate response puzzle," which disappears after controlling for the whole set of external disturbances.[8]

Furthermore, in a Monte Carlo estimation exercise we find that our identification strategy succeeds in recovering the effects of both external and domestic shocks: domestic monetary policy shocks are correctly identified, while the identified external shock maps into the innovation to global risk aversion – the main external driver of exchange rate fluctuations in the model. To the contrary, a recursive VAR analysis on model generated data reproduces

[6] The presence of anticipated effects is related to the Engel and West's (2005) observation that exchange rates tend to predict macro variables.

[7] The presence of anticipated effects invalidates the standard assumption of block exogeneity.

[8] The puzzle arose primarily in developing and emerging economies. In these countries, external shocks have a larger quantitative importance on exchange rates, since they feature large net foreign debt in the analyzed sample period.

the exchange rate response puzzle, exactly because it conflates domestic and external shocks.

Related literature. This paper builds on several strands of the literature concerned with understanding open economy fluctuations.

First, we contribute to the literature on the so-called UIP puzzle (see, e.g., Engel, 2014) by documenting a new conditional property of UIP deviations: they are large and persistent after external shocks, and small and insignificant after domestic shocks. In turn, we find that the bulk of external variation in exchange rates and currency excess returns is related to changes in the risk appetite of global investors. These findings confirm and extend recent evidence on the patterns of UIP deviations. Using firm-level data from Turkey, di Giovanni et al. (2017) document the presence of significant UIP deviations at both firm and country level, and show that these are strongly correlated with movements in the VIX. In the literature on carry-trade strategies, Lustig et al. (2011) identify a slope factor in exchange rate changes that is closely related to changes in volatility of equity markets around the world, while Della Corte et al. (2016) show that investors' exposure to countries' external imbalances explains the cross-sectional variation in currency excess returns.[9]

Second, by characterizing the nature of the main external driver of SOE exchange rates, we contribute to the literature on the empirical importance of global shocks, recently exemplified by Bruno and Shin (2015), Rey (2013) and Miranda-Agrippino and Rey (2015). These authors document large financial spillovers to global asset prices associated with vari-

[9] See also Lustig and Verdelhan (2019).

ations in global risk aversion, typically proxied by the VIX.[10] We similarly find that the main external driver of SOEs' exchange rate is indeed associated with variations in global risk aversion. In addition, we show that this shock leads to demand-like comovement among U.S. output, inflation, and interest rates, and a country's net foreign asset position explains the strength of their spillover effects. We provide a dynamic general equilibrium model that explains both the comovement and the cross-sectional exposure to shocks to global risk aversion.

Third, our empirical findings inform the literature on open economy models. In particular, our empirical analysis points to the presence of one external shock that generates large UIP deviations as well as U.S. demand-driven economic fluctuations. Exogenous global risk aversion shocks, in a model with a non-zero steady-state net foreign asset position, satisfies these properties. The workhorse New-Keynesian models in the literature á la Galí and Monacelli (2005) instead assume UIP and abstract from characterizing the sources of external variation. Recently, Itskhoki and Mukhin (2017, 2019) developed an open-economy model with UIP deviations generated by noise trader shocks in segmented international asset markets. Our proposed global risk aversion shock differs from noise-trader shocks in two key dimensions. First, in our framework global risk aversion shocks affect both U.S. financial traders and U.S. households, thereby generating the global comovement pattern

[10] Other papers that study the effect of specific U.S. or global shocks on SOEs include Canova (2005), Uribe and Yue (2006), Mackowiak (2007), Akinci (2013), Levchenko and Pandalai-Nayar (2015), Ben Zeev et al. (2017), Vicondoa (2019), Scott Davis and Zlate (2019), Iacoviello and Navarro (2018), Cesa-Bianchi et al. (2018), Bhattarai et al. (2017), and Fernández et al. (2016)

that we document empirically. Second, global risk aversion shocks affect SOEs differently depending on their average net foreign asset position, in line with our cross-country evidence on exchange rate dynamics.[11] To the contrary, noise trader shocks have generally negligible effects on U.S. macroeconomic aggregates, and cannot explain the documented cross-country differences in exposure to external disturbances.[12]

Last, our analysis highlights some challenges faced by the VAR literature in identifying shocks in SOEs. In this context, we revisit some empirical evidence on the exchange rate response to domestic monetary policy (Hnatkovska et al., 2016). We show that recent puzzling estimates of the exchange rate effects to monetary policy shocks arise because recursive identification approaches commingle domestic and external shocks, which feature opposite comovement patterns between interest rate differentials and exchange rates.[13]

[11]Nearly every open economy model assumes a zero steady-state NFA position. Some exceptions are Benigno (2009), Cavallo and Ghironi (2002), Ghironi (2008) and Ghironi et al. (2008) who focus on different issues relative to this paper.

[12]Devereux and Engel (2002), Eichenbaum et al. (2017), Cavallino (Forthcoming), and Fanelli and Straub (2018) also present models with shocks to the UIP condition. These shocks share the same properties of noise-trader shocks. Akinci and Queralto (2018) propose a New Keynesian model in which endogenous UIP deviations arise from limits to arbitrage in private intermediation. In Akinci and Queralto's (2018) model, domestic and external shocks lead to UIP deviations of similar size.

[13]Jääskelä and Jennings (2011) and Carrillo and Elizondo (2015) use data simulated from specific models to examine the performance of different VAR schemes in recovering the effects of monetary policy in SOEs. A related literature is concerned with the ability of structural DSGE models to account for the substantial influence of external disturbances. See, for example, Justiniano and Preston (2010), Guerron-Quintana (2013), Alpanda and Aysun (2014), and Georgiadis and Jancoková (2017).

2.2 Decomposing exchange rate variation in SOEs

We are interested in decomposing the exchange rate variation of SOEs according to its sources. In this section, we briefly describe our dataset, outline our identifying assumptions, and explain how to implement our proposed approach in a VAR framework.

Data. We focus on a group of advanced and emerging SOEs: Australia, Austria, Belgium, Brazil, Canada, France, Germany, Indonesia, Italy, Japan, Mexico, New Zealand, Norway, Philippines, South Africa, South Korea, Sweden, Switzerland, Thailand, and United Kingdom. We analyze time periods that are characterized by a flexible exchange rate regime, following Ilzetzki et al.'s (2017) classification.[14] Further details on data sources and selection criteria are reported in Appendix B.1.

Identifying assumptions. At this stage, our objective is to decompose the sources of exchange rate variation in SOEs, while being agnostic about their structural interpretation. To do so, we impose a set of identifying restriction that is consistent with any class of SOE models – in fact with the very definition of a SOE – regardless of the underlying set of structural disturbances or transmission mechanisms. In an *open* economy, domestic variables respond to external shocks. In a *small* economy, domestic (i.e. idiosyncratic) shocks do not affect external variables. Thus, our identifying assumptions hold that any domestic

[14]The longest sample period covers 1974:1-2010:12. For Eurozone countries, we used their national exchange rates before the introduction of the Euro as separate episodes.

shock of the SOE does not affect external variables *at any horizon*, while external shocks affect external variables at least at some horizon.

Baseline SOE VAR. Throughout the paper, we present a number of VARs that feature domestic (SOE) variables and external (U.S.) variables. Our baseline is a three-variable VAR that features U.S. interest rates, domestic interest rates, and the exchange rate. A three-variable VAR allows us to compare our results to those obtained in standard UIP regressions (Section 2.3), and transparently compare the implications of different identification strategies (Section 2.7). In Section 2.4, we extend our VARs to feature additional macroeconomic and financial variables in order to trace out the effects of identified shocks on other macroeconomic variables.[15]

VAR implementation. Consider a three-variable VAR with the Federal Funds rate ($r^{\star}$), the policy-controlled interest rate of SOE k (r_k), and the logarithm of the bilateral nominal exchange rate between country k's currency and the U.S. dollar (s). Exchange rates are in domestic currency units per US dollar, so that an increase is a depreciation of local currency relative to the US dollar. The model is specified in levels and the number of lags is chosen according to the Akaike information criterion. Unlike the case of a vector error correction model, the estimators of the impulse responses of a VAR in levels are consistent

[15] We verify that our VAR is informationally sufficient, by applying the test proposed by Forni and Gambetti (2014).

in the presence of nonstationary but cointegrated variables where the form of cointegration is unknown. Furthermore, estimators are consistent even in the absence of a cointegrating relations among the variables, provided that enough lags are included in the VAR (see Hamilton, 1994).

Thus, let $y_t \equiv [r_t^\star \;\; r_{k,t} \;\; s_t]'$ be the 3×1 vector of observable variables that have length T, including the Federal Funds rate, the policy-controlled interest rate of country k, and the log of the nominal exchange rate, respectively. Denote by $y_t = B(L)u_t$ the reduced-form moving average representation in the levels of the observable variables, formed by estimating an unrestricted VAR in levels. The relationship between reduced-form innovations and structural shocks is given by:

$$u_t = A_0 \varepsilon_t \tag{2.1}$$

which implies the following structural moving average representation:

$$y_t = B(L) A_0 \varepsilon_t. \tag{2.2}$$

We assume that the structural shocks are orthogonal with unitary variance, so that the impact matrix A_0 satisfies $A_0 A_0' = \Sigma$, where Σ is the variance-covariance matrix of innovations. In order to identify A_0, one needs to impose $n(n-1)/2$ additional restrictions, where n is the number of variables included in the VAR.

Within the above three-variable VAR, we propose an identification strategy designed to separately identify the effects of external shocks from those of idiosyncratic shocks stemming from country k. Specifically, we assume that the external variable in the VAR, the Federal

Funds rate, is properly characterized as following a stochastic process driven by unanticipated and anticipated shocks (their respective statistical properties are described below). The domestic source of variation of the SOE is then identified as the linear combination of the VAR innovations that is orthogonal to (unanticipated and anticipated) external shocks.

To implement our identification scheme in the three-variable VAR presented above, we note that the impact matrix A_0, defined in Eq. (2.1), is unique up to any rotation D of the structural shocks. Specifically, for any 3×3 orthonormal matrix D, the entire space of permissible impact matrices can be written as $\widetilde{A}_0 D$, where $\widetilde{A}_0$ is an arbitrary orthogonalization (e.g. the one implied by a recursive identification scheme).

Here, the h-step ahead forecast error is

$$y_{t+h} - E_{t-1} y_{t+h} = \sum_{\tau=0}^{h} B_\tau \widetilde{A_0} D \varepsilon_{t+h-\tau}$$

where B_τ is the matrix of moving average coefficients at horizon τ. The share of the forecast error variance of variable i attributable to the structural shock j at horizon h is then:

$$\Omega_{i,j}(h) = \frac{\sum_{\tau=0}^{h} B_{i,\tau} \widetilde{A}_0 \gamma \gamma' \widetilde{A}_0' B_{i,\tau}'}{\sum_{\tau=0}^{h} B_{i,\tau} \Sigma B_{i,\tau}'}$$

where γ is the j-th column of D, while $B_{i,\tau}$ corresponds to the i-th row of B_τ.

To separately identify domestic and external sources of SOE fluctuations, we adopt a procedure that extends the identification scheme proposed by Barsky and Sims (2011b).[16] This approach can be explained as composed of two steps. First, we recover the unanticipated and the anticipated movements in the Federal Funds rate. The former is identified

[16] In using a maximum forecast error variance approach, Barsky and Sims (2011b) build on earlier work by Faust (1998), Uhlig (2003). See also Francis et al. (2014).

as the orthogonal innovation in $r^\star$. The latter is identified as the shock that maximizes the contribution to the forecast error variance of the Federal Funds rate up to a truncation horizon H, subject to the restriction that this shock has no contemporaneous effect on the Federal Funds rate.[17] Formally, the identification of the anticipated external shock boils down to solving the following maximization problem:

$$\gamma * = \arg\max \sum_{h=0}^{H} \Omega_{1,2}(h) = \frac{\sum_{\tau=0}^{h} B_{i,\tau} \widetilde{A}_0 \gamma \gamma' \widetilde{A}_0' B_{i,\tau}'}{\sum_{\tau=0}^{h} B_{i,\tau} \Sigma B_{i,\tau}'}$$

s.t.

$$\widetilde{A}_0(1,j) = 0 \quad \forall j > 1$$

$$\gamma(1,1) = 0$$

$$\gamma' \gamma = 1$$

where the first two constraints ensure that the anticipated external shock has no contemporaneous effect on the Federal Funds rate, and the third restriction narrows the solution space to the one of possible orthogonalizations of the reduced form, by preserving the orthonormality of the rotation matrix D. By imposing that γ must be a unit vector, the second column γ of matrix D is identified. The second step consists in recovering the domestic shock of SOE k. This shock can be identified by making use of the condition that the matrix D must be orthonormal, i.e. $DD' = D'D = I$. More specifically, letting $\gamma * = [0 \ \gamma_1 \ \gamma_2]$

[17] Our empirical results are robust to relaxing this contemporaneous restriction.

where $\gamma_2 = -\sqrt{1-\gamma_1^2}$, then one can express D as:[18]

$$D = \begin{bmatrix} 1 & 0 & 0 \\ 0 & \gamma_1 & \gamma_2 \\ 0 & -\gamma_2 & \gamma_1 \end{bmatrix} \tag{2.3}$$

where the first column ensures that the unanticipated external shock ($\varepsilon_t^{\star}$) is the orthogonal innovation to the Federal Funds rate, the second column results from the maximization problem above and therefore captures the whole set of shocks that induce future movements in the Federal Funds rate ($\varepsilon_t^{\star\star}$), and the third column identifies the domestic shock of country k (ε_t^{SOE}) that may affect both the nominal exchange rate and the policy controlled interest rate, while it has no contemporaneous or future impact on the external variable ($r^{\star}$).[19] Last, *for any* orthogonalization $\widetilde{A}_0$ of residuals u_t which satisfies the first constraint of the above maximization problem, the structural shocks can be recovered from the relation

$$u_t = \widetilde{A}_0 D \varepsilon_t. \tag{2.4}$$

where D is the rotation matrix previously identified, and $\varepsilon_t \equiv [\varepsilon_t^{\star} \;\; \varepsilon_t^{\star\star} \;\; \varepsilon_t^{SOE}]'$.

[18]The negative sign in front of γ_2 is just a normalization. Specifically, to preserve the orthonormality of D, one needs the 2×2 lower right submatrix of D to have either opposite diagonal elements or opposite off-diagonal elements.

[19]By construction, this condition is subjected to the maximization above, therefore results can still deliver that a domestic shock has some, but likely insignificant, future effects on the Federal Funds rate.

2.3 Conditional properties of exchange rates in SOEs

How important are domestic and external sources of fluctuations in SOEs' exchange rates and currency excess returns? Do different shocks generate different dynamic patterns of currency excess returns? In this section, we illustrate the relative contribution of different shocks for our variables of interest, and discuss their properties. The empirical evidence reported below is the result of estimating a set of individual-country VARs using the approach described in Section 2.2. We frame our main results in the form of impulse response functions (IRFs). Bias-corrected bootstrapped 90% confidence intervals are based on 1000 replications (see Kilian, 1998).

Definition of currency excess returns. In line with the relevant literature, the ex ante excess return on the domestic bond held from period t to period $t+m$, inclusive of the expected currency return, is defined as:

$$\mathrm{E}_t\,\hat{x}_{t+m} \equiv \hat{r}_{t|m} - \hat{r}^{\star}_{t|m} - \mathrm{E}_t\,\Delta\hat{s}_{t+m} \tag{2.5}$$

where hatted variables denote series generated by our VAR, E_t is the expectation operator conditional on time-t information, and $\hat{r}_{t|m}$ ($\hat{r}^{\star}_{t|m}$) are m-month domestic (foreign) interest rates.[20] Non-zero ex ante excess returns point to violation of so-called UIP. In fact, under UIP the exchange rate is expected to depreciate at a rate that equals the interest rate

[20]Below, we report the returns from an investment of one year maturity on the domestic bond. That is, $m = 12$ months, which is the typical maturity of the domestic interest rates in our sample.

differential.

In addition, let us define the counterfactual response of the exchange rate that one would observe under UIP. Following Engel (2016), we iterate Eq. (2.5) forward and obtain a relationship between the the level of the exchange rate and the expected path of interest rate differentials and excess returns:[21]

$$\hat{s}_t = \hat{s}_t^{UIP} + \mathrm{E}_t \sum_{j=0}^{\infty} \hat{x}_{t+j+1} \tag{2.6}$$

where $\hat{s}_t^{UIP} \equiv -\,\mathrm{E}_t \sum_{j=0}^{\infty} \left(\hat{r}_{t+j} - \hat{r}^{\star}_{t+j}\right)$ is the exchange rate level consistent with UIP. The difference between $\hat{s}_t$ and $\hat{s}_t^{UIP}$ is accounted for by the infinite sum of ex ante excess returns. Below we report $\mathrm{E}_t\, \hat{x}_{t+m}$ and $\hat{s}_t^{UIP}$ conditional on domestic (ε^{SOE}) and external shocks ($\varepsilon^{\star}$ and $\varepsilon^{\star\star}$). These objects are constructed using the expectations implied by the VAR.

Relative importance of domestic and external shocks. Figure 2.1 reports the variance decomposition for our baseline variables, along with expected currency excess returns. The Federal Funds rate appears to be exclusively explained by external disturbances. This outcome indicates that our two external shocks capture all the unpredictable fluctuations in the Federal Funds rate. The domestic interest rate is also predominantly driven by external shocks, in line with the observation that SOE monetary policy is largely devoted to respond to external sources of fluctuations. In the typical SOE, the exchange rate is

[21] In deriving Eq. (2.6) we impose that $\lim_{j \to \infty} \hat{s}_{t+j} = 0$, consistent with the observation that our VAR generates stationary time series.

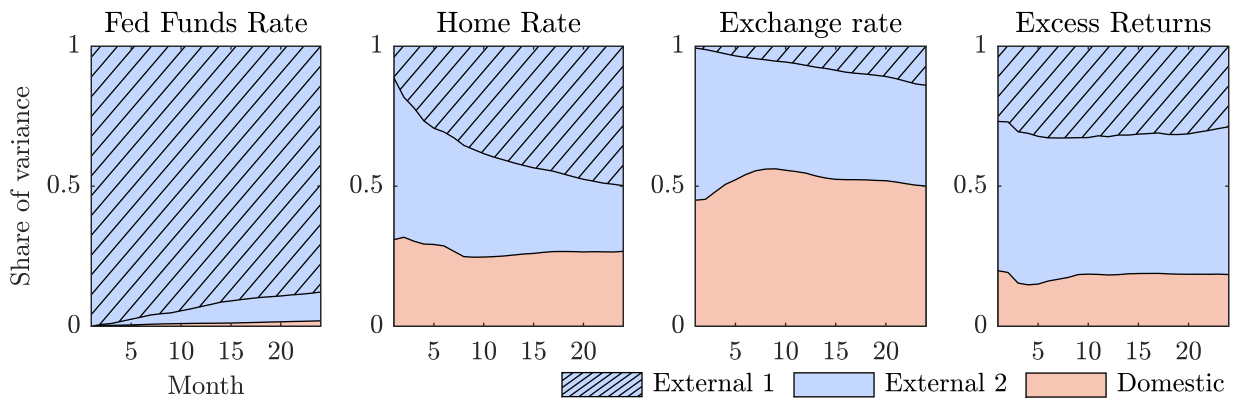


Figure 2.1: Relative contribution of domestic and external shocks

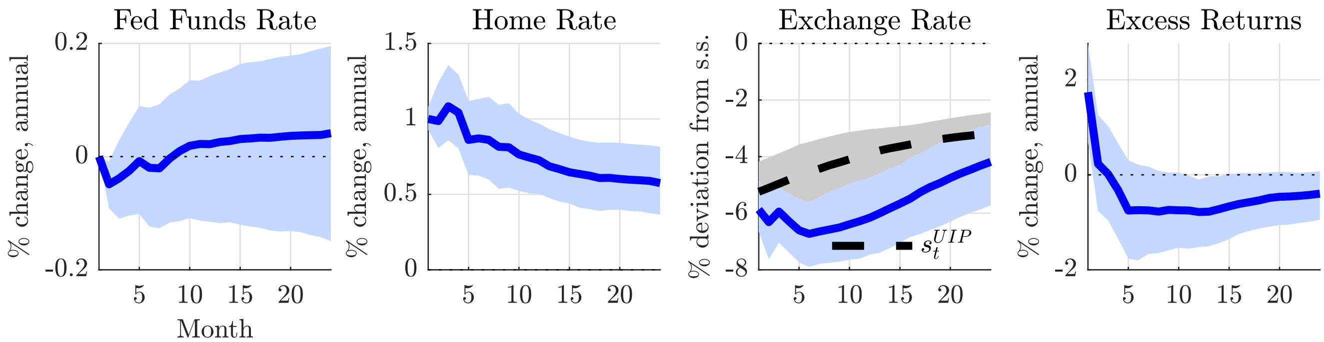
Note: The horizontal axes refer to forecast horizons, while the vertical axes denote the fraction of forecast error variance from each shock. External shocks consist of unanticipated (External 1) and anticipated (External 2) variation in the Federal Funds rate.

explained by domestic and external shocks in almost equal parts.[22] However, expected currency excess returns are predominantly part explained by external disturbances, suggesting that domestic and external shocks imply significantly different exchange rate dynamics.

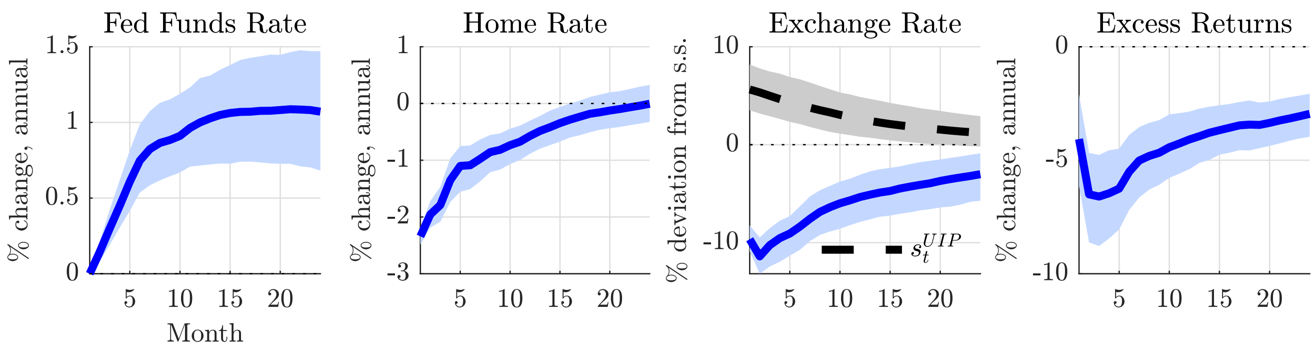
We note that between the two external shocks that we identify, the anticipated external shock ($\varepsilon^{\star\star}$) is by far the main external driver of exchange rates and excess returns. In fact, it explains more than 3/4 of the external variation in exchange rates, and more than 2/3 of the external variation in currency excess returns. For this reason, below we will solely focus on this source of external fluctuations, and we will refer to it as "the external shock."

Conditional interest rate and exchange rate dynamics. We are interested in under-

[22]Section 2.6 explores the cross-country differences in exchange rate exposure to external shocks.



(a) Empirical impulse responses to a domestic shock



(b) Empirical impulse responses to an external shock

Figure 2.2: Conditional properties of interest rates and exchange rates

Note: The lines denote median IRFs by countries with corresponding 90% confidence intervals from 1000 bias-corrected bootstrap replications of the reduced-form VAR. Domestic shocks are normalized to deliver a 1% impact increase in the home interest rate, while external shocks are normalized to deliver a 1% increase in the Fed Funds rate at one-year horizon. Excess returns are one-year ahead expected excess returns.

standing the comovement among interest rates, exchange rates, and one-year ahead ex ante excess returns implied by domestic and external shocks.

Figure 2.2 collects our findings. A domestic shock that leads to a 1% increase in the domestic interest rate is associated with an impact exchange rate appreciation (Figure 2.2a) and a largely insignificant response of currency excess returns. In fact, exchange rate dynamics under domestic shocks are both qualitatively and quantitatively in line with the UIP-consistent exchange rate response, $\hat{s}_t^{UIP}$.

After an external shock that leads to an increase in the foreign interest rate, the domestic interest rate declines significantly. Because the interest rate differential is persistently negative, UIP predicts a significant currency depreciation. However, the observed exchange rate response implies a significant currency appreciation, accounted for by large and persistent decline in excess returns required on the domestic bond.

Therefore, our evidence points to large and persistent UIP deviations due to external shocks, but not in response to domestic shocks. Importantly, the conditional differences in UIP deviations are so large that they generate an opposite comovement patterns in interest rate differentials and exchange rates across these two sources of variation.

Conditional UIP regression coefficients. Figure 2.3 documents that these conditional patterns also hold in country-specific VARs, with only few exceptions. The external variation in the exchange rates is associated with significant and predictable deviations from UIP. To the contrary, the Fama's (1984) coefficient computed under domestic shocks is largely insignificant.

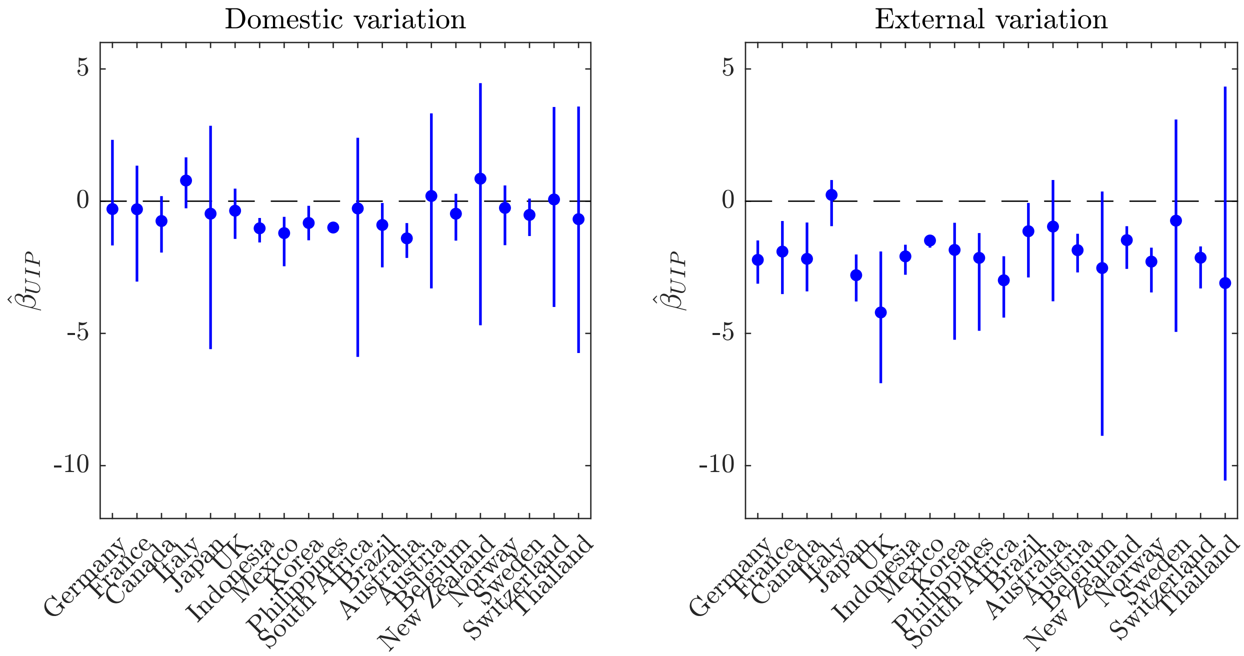


Figure 2.3: UIP violations

Note: The figure reports the UIP regression coefficients $\hat{\beta}_{UIP}$ in conditional versions of the Fama's (1984) regression: $-\operatorname{E}_t \hat{x}_{t+m} = \alpha + \beta_{UIP}(\hat{r}_{t|m} - \hat{r}^{\star}_{t|m}) + \varepsilon_t$. Excess returns on the domestic currency $\operatorname{E}_t \hat{x}_{t+m}$, defined in Equation (2.5), are constructed using the conditional expectations implied by the VAR. We set $m = 12$ months. For each country, we report the median value of the coefficient along with 90% confidence intervals from 1000 bias-corrected bootstrap replications of the reduced-form VAR.

2.4 External shocks are global risk aversion shocks

Our evidence indicates that one external source of fluctuations is responsible for a large fraction of the observed variation in expected currency excess returns. A natural question is whether this external shock has an appealing interpretation. In this section, we trace out the effects of external shocks on key U.S. macroeconomic variables.

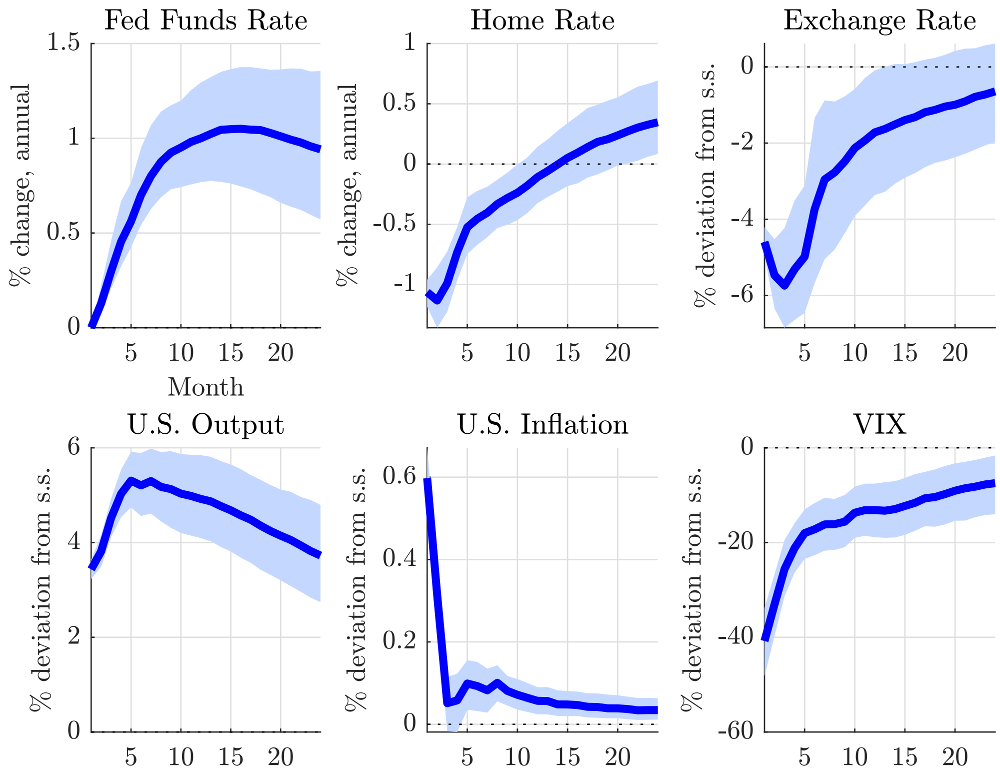


Figure 2.4: Empirical impulse responses to an external shock (Extended VAR)

Note: This figure features the estimated IRFs to an external shock in VARs that include a set of external variables. We run four-variable VARs that include the three baseline variables and either U.S. industrial production, U.S. CPI inflation, or the VIX ordered fourth. The lines denote median IRFs across countries. The shaded areas are the corresponding 90% confidence intervals from 1000 bias-corrected bootstrap replications of the reduced-form VAR.

In particular, we study the effects of the external shock in a set of extended VARs, that include U.S. industrial production, inflation in the U.S. Consumer Price Index (CPI), and the Chicago Board Options Exchange Volatility Index (VIX), a forward-looking measure of uncertainty and risk aversion. Figure 2.4 shows that the external shock leads to an increase

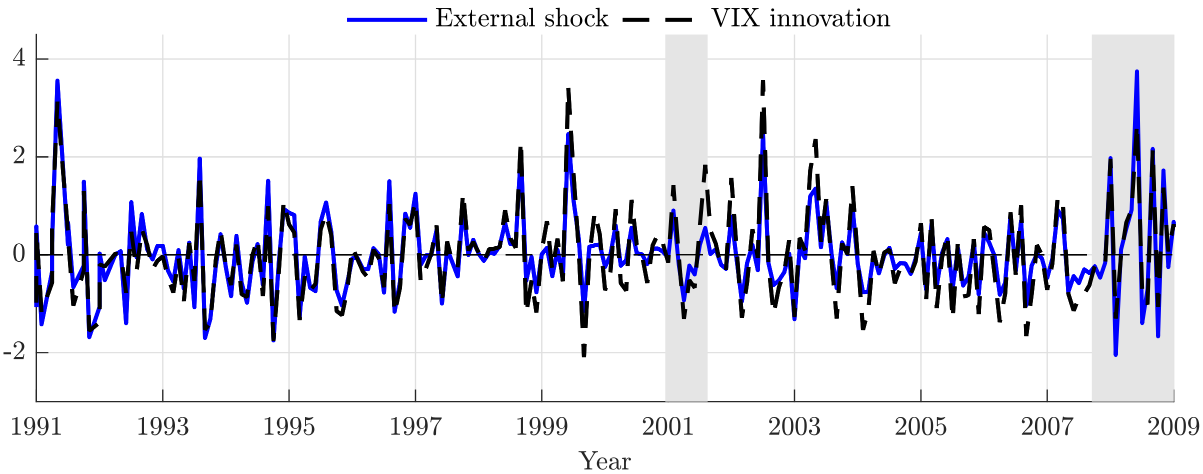


Figure 2.5: Identified external shocks and the VIX

Note: The figure plots the identified series of external shocks and the innovation in the VIX. The innovation in the VIX is computed as the residual of an AR(1) process. We report the IRFs for three calibrations of the steady-state value of the NFA/GDP position of the SOE ($\mathtt{b}$)

in U.S. output, U.S. inflation and the Federal Funds rate – a comovement that is typical of demand-driven expansions. In addition, these U.S. economic expansions are accompanied by a temporary decline in the VIX, and generate significant appreciations of SOEs exchange rates against the U.S. dollar.

The international finance literature has documented that global asset prices display significant comovement with the VIX, a common proxy of gloabl risk aversion (see, e.g., Bruno and Shin, 2015, Rey, 2013, and Miranda-Agrippino and Rey, 2015). In Figure 2.5 we report the historical series of our external shock, along with the innovation in the VIX, computed as the residual of an AR(1) process. We find that our estimated external shocks are intimately associated with movements in global risk aversion. In fact, the correlation between our identified series of external shocks and the innovation in the VIX is around 0.8. This evidence suggests that the core of the external variation in the exchange rates may be the result of fluctuations in risk appetite in international asset markets that also give rise to U.S. economic fluctuations. In the next sections, we formalize this interpretation in a dynamics two-country SOE model, and test its main predictions.

2.5 A SOE model with global risk aversion shocks

To rationalize our empirical findings, we build a two-country SOE dynamic general equilibrium model. After a brief introduction of the model environment, we present a summary of the equilibrium conditions and highlight the key economic mechanisms. Appendix B.3 contains the full derivation of the model.

2.5.1 Environment

Our model economy consists of two countries, the SOE and a large economy. To characterize the SOE, we follow De Paoli (2009) in taking the limit of the home economy size to zero.[23] The foreign (large) economy is then interpreted as the U.S.. The core of our model belongs to the international macroeconomic tradition initiated by Obstfeld and Rogoff (1995), in that it consists of a dynamic general equilibrium open-economy model with monopolistically competitive producers, sticky prices, and complete exchange rate pass-through.[24] Asset

[23]The limit is taken after having derived the equilibrium conditions for the two-country model.

[24]Complete exchange rate pass through obtains because prices are set in the producer's currency.

markets are both incomplete and segmented. The only assets available in the economy are two nominal riskless bonds denominated in home and foreign currency. We assume that households in each economy can only trade the bond of their respective country, and all international transactions are intermediated by a set of U.S. financial traders who are averse to taking risky positions (Jeanne and Rose, 2002, Gabaix and Maggiori, 2015, Itskhoki and Mukhin, 2017). In our model, financial traders are a subset of U.S. households, and we assume that their risk aversion is exogenous and time-varying.[25]

Households and the financial sector

The world economy is populated with a continuum of agents of unit mass, where the population in the segment $[0, n)$ belongs to the home (H) country and the population in the segment $(n, 1]$ belongs to the foreign (F) country.

Domestic economy. The domestic economy is populated by a representative household whose preferences are given by

$$\mathrm{E}_t \sum_{j=0}^{\infty} \beta^j \left[\frac{C_t^{1-\omega}}{1-\omega} - \frac{N_t^{1+\eta}}{1+\eta} \right] \tag{2.7}$$

where N_t denotes hours worked, and C_t is a composite consumption index defined by

$$C_t \equiv \left[(\nu)^{\frac{1}{\theta}} (C_{H,t})^{\frac{\theta-1}{\theta}} + (1-\nu)^{\frac{1}{\theta}} (C_{F,t})^{\frac{\theta-1}{\theta}} \right]^{\frac{\theta}{\theta-1}}$$

[25] Because financial traders are a subset of U.S. households, the U.S. is interpreted as the center of the international financial system.

where $C_{H,t}$ is an index of consumption of domestic goods given by the CES function

$$C_{H,t} \equiv \left[\left(\frac{1}{n} \right)^{\frac{1}{\iota}} \int_0^n C_{H,t}(i)^{\frac{\iota-1}{\iota}} \mathrm{d}i \right]^{\frac{\iota}{\iota-1}}$$

where $i \in [0,1]$ denotes the good variety. $C_{F,t}$ is an index of goods imported from the foreign country given by an analogous CES function:

$$C_{F,t} \equiv \left[\left(\frac{1}{1-n} \right)^{\frac{1}{\iota}} \int_n^1 C_{F,t}(i)^{\frac{\iota-1}{\iota}} \mathrm{d}i \right]^{\frac{\iota}{\iota-1}}$$

Parameter $\iota > 1$ denotes the elasticity of substitution between varieties (produced within any given country). Parameter $1 - \nu \in [0,1]$ governs the home consumers' preferences for foreign goods, and is a function of the relative size of the foreign economy, $1 - n$, and of the degree of openness, λ, namely $1 - \nu = (1 - n)\lambda$. Parameter $\theta > 0$ measures the substitutability between domestic and foreign goods, from the viewpoint of the domestic consumer.

Domestic households can trade only a one-period nominal bond, which is denominated in domestic currency. The domestic household's flow budget constraint is given by

$$\frac{B_{t+1}}{R_t} + P_t C_t = W_t N_t + B_t$$

where B_{t+1} denotes the nominal balance of home bonds, R_t is the nominal interest rate on the home bond, P_t is the price index of the composite consumption good, C_t, and W_t is the nominal wage rate. The problem of the domestic household consists in maximizing its utility (Eq. 2.7) subject to the budget constraint (Eq. 2.8). The first-order conditions of this problem are standard and therefore relegated to Appendix B.3.

Foreign economy. The foreign economy is populated by a continuum of households. At the beginning of each period, all members of a household are identical and share the household's assets. During the period, the members are separated from each other, and each member receives a shock that determines her role in the period. A member will be a trader with probability m_t, and a worker with probability $1 - m_t$. These shocks are i.i.d. among the members. We assume that the share of members that operate as traders in the international financial market is proportional to the output of the home economy (that is, $m_t = \mu n P^{\star}_{H,t} Y_t$). This assumption entails that traders devote a larger part of their balance sheets to bonds issued by larger economies. The members' preferences are aggregated and represented by the following utility function of the household:

$$\mathrm{E}_t \sum_{j=0}^{\infty} \beta^{\star j} \left[m_t \mathcal{U}(\widetilde{C}^{\star}_t) + (1 - m_t)\mathcal{U}(C^{\star}_t, N^{\star}_t) \right]$$

where

$$\mathcal{U}(\widetilde{C}^{\star}_t) \equiv \frac{\left(\widetilde{C}^{\star}_t\right)^{1-\omega^{\star}_t}}{1 - \omega^{\star}_t} \tag{2.8}$$

and

$$\mathcal{U}(C^{\star}_t, N^{\star}_t) \equiv \frac{(C^{\star}_t)^{1-\omega^{\star}_t}}{1 - \omega^{\star}_t} - \frac{(N^{\star}_t)^{1+\eta}}{1 + \eta}$$

Here, $\widetilde{C}^{\star}_t$ is the consumption of traders, $C^{\star}_t$ is the consumption of workers, and $\omega^{\star}_t$ governs the degree of (relative) risk aversion of both household's members. We assume that foreign households' risk aversion is time varying. In particular, $\omega^{\star}_t = \omega^{\star} \exp(\xi_t)$ and its time-varying component evolves according to the following autoregressive process:

$$\xi_t = \rho_\xi \xi_{t-1} + \varepsilon_{\xi,t} \tag{2.9}$$

where $\varepsilon_{\xi,t}$ are i.i.d. disturbances drawn from a Normal distribution with mean zero and standard deviation σ_ξ. The problem of the worker-members of the foreign household is standard, and analogous to the one of the domestic household. Her intertemporal budget constraint reads

$$\frac{B^\star_{t+1}}{R^\star_t} + P^\star_t C^\star_t = B^\star_t + W^\star_t N^\star_t - \frac{m_t}{1-m_t} T^\star$$

where the last term is an intrahousehold transfer that accrues to the trader-members of the households, and ensures that their consumption is always positive. The other foreign variables are interpreted analogously to their domestic counterparts. The first-order conditions of this problem are standard and therefore relegated to Appendix B.3.

Traders on the foreign exchange market. The trader-members of the foreign household are the only agents who can trade bonds internationally.[26] Traders collectively take a zero-capital position $\widetilde{D}_{t+1}$ in home-currency bonds and short $\widetilde{D}^\star_{t+1} = -\widetilde{D}_{t+1}/S_t$ foreign-currency bonds, or *vice versa*. Here, S_t is the nominal exchange rate, defined to be the price of the foreign currency unit, as in the empirical section. The exchange rate is relevant for the balance sheet of international traders because each economy offers a bond in its own currency. A one U.S.-dollar position generates a U.S.-dollar return of $\widetilde{R}_{t+1} = R^\star_t - R_t \frac{S_t}{S_{t+1}}$. The problem of each individual trader consists in choosing a position $d^\star_{t+1}$ to maximize (2.8) subject to the budget constraint $P^\star_t \widetilde{C}^\star_t = T^\star + \widetilde{R}_{t+1} d^\star_{t+1}$.[27] In Appendix B.3.1, we show that

[26] Since traders are part of the foreign household, the foreign economy is interpreted as the center of the international financial system.

[27] Again, $T^\star$ denotes a constant intra-household transfer that ensures that each trader's consumption is

the individual trader's problem is approximately equivalent to maximizing a mean-variance utility of returns. The resulting demand for home-currency bonds by the financial traders is then:

$$\widetilde{D}^{\star}_{t+1} = \frac{m_t}{\omega^{\star}_t}\frac{\mathrm{E}_t\,\widetilde{R}_{t+1}}{\mathrm{Var}_t(\widetilde{R}_{t+1})} \Rightarrow \frac{\widetilde{D}_{t+1}}{S_t} = -\frac{m_t}{\omega^{\star}_t}\frac{\mathrm{E}_t\,\widetilde{R}_{t+1}}{\mathrm{Var}_t(\widetilde{R}_{t+1})} \tag{2.10}$$

The financial market clears when the interest rates R_t and $R^{\star}_t$ are such that $B_{t+1}+D_{t+1}=0$ and $B^{\star}_{t+1}+D^{\star}_{t+1}=0$. This condition implies that in equilibrium the net foreign asset position of home equals net foreign liabilities of foreign, $nB_{t+1}=-(1-n)B^{\star}_{t+1}S_t$, in aggregate per-capita terms.[28] Thus, Eq. (2.10) becomes:

$$-\frac{B_{t+1}}{P_{H,t}Y_t} = \frac{\mu}{\omega^{\star}_t}\frac{\mathrm{E}_t\left(R_t\frac{S_t}{S_{t+1}}-R^{\star}_t\right)}{\mathrm{Var}_t(\widetilde{R}_{t+1})} \tag{2.11}$$

Finally, we follow De Paoli (2009) in taking the limit for $n\to 0$ to portray our SOE. This implies that economic developments in the large economy affect the SOE, but the reverse is not true. Under this assumption, the mass of household-traders $m_t\to 0$, $\forall t$. As a result, traders influence the model's behavioral equations only through their pricing of the exchange rate. The resulting profits from their trading activity are infinitesimally small from the standpoint of the foreign economy, and don't affect the household's budget constraint.

We solve the model by log-linearization around a steady state with a non-zero net foreign asset position, and use $\mathtt{b}\equiv {}^{B}/_{P_H Y}$ to denote the steady-state net foreign asset position relative to GDP of the home economy. Using the international bond market clearing condition,

always non-negative

[28]Here, $nD_t=\widetilde{D}_t$ and $(1-n)D^{\star}_t=\widetilde{D}^{\star}_t$.

the linearized version of the traders' bond demand (Eq. 2.11) reads:[29]

$$\chi\left(-b_{t+1}-\mathtt{b}\xi_t\right)\approx r_t-r_t^{\star}-\mathrm{E}_t\,\Delta s_{t+1} \tag{2.12}$$

where $\chi \equiv \frac{\sigma_s^2}{\mu/\omega^{\star}}$ governs traders' risk bearing capacity in steady state.

Before we close the model, we can outline the mechanism and a testable implication of our framework. Eq. (2.12) is the exchange rate determination equation of our model economy. As in Itskhoki and Mukhin (2017), the standard UIP condition obtains as a special case when the risk-bearing capacity of traders $\chi = 0$. In our model, $\chi = 0$ if traders are risk neutral ($\omega^{\star} = 0$), the size of the financial sector $\mu \to \infty$, or the exchange rate is non-stochastic ($\sigma_s^2 \equiv \mathrm{Var}_t(\Delta s_{t+1}) = 0$). The variance of the innovation to the nominal exchange rate, σ_s^2, is endogenously determined. If $\chi > 0$, the model economy features two sources of time-varying UIP deviations - endogenous movements in the net foreign asset position to GDP, b_{t+1} and exogenous changes in global risk aversion ξ_t.[30] First, as emphasized by Gabaix and Maggiori (2015), an equilibrium imbalance that requires traders to be long in a currency generates a positive expected excess return of this currency. In this model, a country's imbalance is directly related to its net foreign asset position to GDP. A negative net foreign asset position requires traders to be long in that country's currency and therefore requires a positive expected return on this currency. Second, for a given level of the net foreign position, changes in global risk aversion affect the degree of expected returns demanded

[29] For illustration purposes, Eq. (2.12) is an approximation in that it ignores the terms arising because of steady-state UIP deviations.

[30] Here, b_{t+1} denotes the equilibrium deviation of net foreign assets to GDP relative to its steady state value. That is $b_{t+1} \equiv {}^{B_{t+1}}/_{P_{H,t}Y_t} - {}^{B}/_{P_H Y}$

by traders in equilibrium. In our linearized model, changes in risk bearing capacity have a *direct* effect on exchange rate determination if a country's steady-state net foreign asset position is non-zero. If the steady-state net foreign asset position of a country is negative, higher global risk aversion requires higher expected returns on this currency to provide the incentive for risk-averse traders to keep absorbing the imbalance. The opposite reasoning holds for countries that are net creditors in steady state.

Firms

Each country features a continuum of firms that produce output under a constant-returns-to-scale production function. The economy-wide production functions are thus $Y_t = AN_t$ and $Y_t^\star = AN_t^\star$ for the domestic and foreign goods, respectively.

We assume that each producer sets its price in her own currency. In this case the law of one price holds. Under these conditions, $P_{H,t} = S_t P_{H,t}^\star$ and $P_{F,t} = S_t P_{F,t}^\star$ for each t. However, the home bias specification leads to deviations from purchasing power parity; that is, $P_t \neq S_t P_t^\star$. Prices follow a partial adjustment rule as in Calvo (1983). Producers of differentiated goods know the form of their individual demand functions, and maximize profits taking overall market prices as given. In each period a fraction, $\alpha \in [0, 1)$, of randomly chosen producers is not allowed to change the nominal price of the goods they produce. The remaining fraction of firms, given by $1 - \alpha$, chooses prices optimally by maximizing the expected discounted value of profits.

Monetary authorities

In each country, the monetary authority is assumed to follow a Taylor (1993)-type rule with interest-rate smoothing:

$$r_t^\star = \rho_r r_{t-1}^\star + (1-\rho_r)\phi\pi_t^\star + \varepsilon_{r^\star,t} \qquad r_t = \rho_r r_{t-1} + (1-\rho_r)\phi\pi_t + \varepsilon_{r,t}$$

where $\varepsilon_{r^\star,t}$ and $\varepsilon_{r,t}$ are i.i.d. disturbances drawn from a Normal distribution with mean zero and standard deviations $\sigma_{r^\star}$, and σ_r, respectively.[31] In line with central banks' practices, we assume that they target a measure of consumer price (CPI) inflation.

2.5.2 Calibration and equilibrium conditions

In our model, the size of traders' balance sheet depends on risk perceptions. To account for risk in the computation of the model, we follow Coeurdacier et al. (2011) in deriving the "risky" steady state – a steady state in which agents expect future risk and the realization of shocks is zero at the current date. The risky steady state differs from the deterministic steady state only by second order terms related to variances and covariances of the endogenous variables. These second moments pin down the size of traders' long-run balance sheet. To analyze model dynamics, we then look at a first order log-linear approximation around the risky steady state. Crucially, we allow the steady-state net foreign assets, $\mathtt{b}$, to be non-zero.[32]

[31] Monetary authorities are assumed to target a zero inflation steady state.

[32] In our model, we allow for different discount factors across countries, that is $\beta \neq \beta^\star$. This gives rise to different steady-state returns on the two countries' bonds, and a non-zero steady-state net foreign

Calibration. Our benchmark value for $\mathtt{b}$ is a net foreign asset position relative to (annual) GDP of around -15%, the median value in our sample of SOEs.[33] Our model is calibrated to a monthly frequency. We set $\beta^\star = 0.9967$ which implies a steady state annual interest rate of about 4%, and $\eta = 1$ which implies a unit Frisch elasticity. Our calibration of the Calvo parameter ($\alpha = 0.9167$) implies an average duration of price contracts of one year. We set the consumption share of imports $\lambda = 0.4$, and the trade elasticity $\theta = 1$. The Taylor-rule coefficient on consumer price inflation, ϕ, equals 1.5, while the parameter that governs the degree of interest rate smoothing, ρ_r, equals 0.947, in line with typically estimated values in the DSGE literature. We set $\rho_\xi = 0.90$.[34]

We choose the variances of the structural shocks so that the model reproduces three empirical moments: the unconditional standard deviation of nominal exchange rate changes ($\mathrm{Std}(\Delta s_t)$), the observed unconditional deviation from UIP (α_1 in $\Delta s_{t+1} = \alpha_0 + \alpha_1(r_t - r_t^\star)$) and the unconditional contemporaneous correlation between the exchange rate and the interest rate differential (β_1 in $\Delta s_t = \beta_0 + \beta_1 \Delta(r_t - r_t^\star)$).

Equilibrium conditions. We report below the model's log-linear equilibrium conditions, evaluated at the risky steady state.[35] The equilibrium conditions that govern economic

position.

[33]Data on annual net foreign asset position to GDP are from the updated and extended version of the dataset constructed by Lane and Milesi-Ferretti (2007).

[34]Without loss of generality we normalize the steady state so that $\ln(C^\star) = 1$.

[35]All variables are expressed as log deviations from their steady state, except for net foreign assets to GDP

dynamics in the large (foreign) economy read:

$$\omega^\star \operatorname{E}_t \Delta c^\star_{t+1} + \omega^\star \operatorname{E}_t \Delta\xi_{t+1} = r^\star_t - \operatorname{E}_t \pi^\star_{t+1} \tag{2.13a}$$

$$\pi^\star_t = \beta^\star \operatorname{E}_t \pi^\star_{t+1} + \kappa^\star((\eta + \omega^\star)c^\star_t + \omega^\star \xi_t) \tag{2.13b}$$

$$r^\star_t = \rho r^\star_{t-1} + (1-\rho)\phi\pi^\star_t + \varepsilon_{r^\star,t} \tag{2.13c}$$

Given the exogenous processes, the economic dynamics in the large economy are fully described by the consumption Euler equation (Eq. 2.13a), the New Keynesian Phillips curve (Eq. 2.13b), and the monetary policy rule (Eq. 2.13c).[36] Both Eqs. (2.13a) and (2.13b) are influenced by shocks to foreign households' risk aversion ("global risk aversion shocks"), which act as taste shocks (cf. Stockman and Tesar, 1995).

Domestic variables are determined according to the following system of log-linear equations:

$$\omega \operatorname{E}_t \Delta c_{t+1} = r_t - \operatorname{E}_t \pi_{t+1} \tag{2.14a}$$

$$\pi_{H,t} = \beta \operatorname{E}_t \pi_{H,t+1} + \kappa(\omega c_t + \eta y_t + \lambda(1-\lambda)^{-1} q_t) \tag{2.14b}$$

$$r_t = \rho r_{t-1} + (1-\rho)\phi\pi_t + \varepsilon_{r,t} \tag{2.14c}$$

$$\pi_t = (1-\lambda)\pi_{H,t} + \lambda(\Delta s_t + \pi^\star_t) \tag{2.14d}$$

$$y_t = \theta\lambda(1-\lambda)^{-1} q_t + (1-\lambda)(1 + \mathtt{b} - \widetilde{\beta}\mathtt{b})c_t + \left[1 - (1-\lambda)(1 + \mathtt{b} - \widetilde{\beta}\mathtt{b})\right](c^\star_t + \theta q_t) \tag{2.14e}$$

$$\widetilde{\beta}\,(b_{t+1} - \mathtt{b} r_t) - b_t + \mathtt{b}\,(\pi_{H,t} + \Delta y_t) = (1 + \mathtt{b} - \widetilde{\beta}\mathtt{b})\left(y_t - c_t - \lambda(1-\lambda)^{-1} q_t\right) \tag{2.14f}$$

(b_t), which is expressed as changes from its steady state. Also, $\widetilde{\beta} \equiv 1/\mathtt{R}$.

[36] The curvature parameter of the foreign economy's Phillips curve is given by $\kappa^\star \equiv \frac{(1-\beta^\star\alpha^\star)(1-\alpha^\star)}{\alpha^\star}$.

$$\Delta s_t = \Delta q_t - \pi_t^\star + \pi_t \tag{2.14g}$$

Since the SOE is effectively open to trade in goods and assets, it is affected by the dynamics of the exchange rate and foreign demand, as in the canonical model with complete exchange rate pass-through.[37] The key difference relative to the standard framework consists in the exchange rate determination, which is governed by Eq. (2.12), described above.

In this environment, there are three structural shocks: home and foreign monetary policy innovations ($\varepsilon_{r,t}$ and $\varepsilon_{r^\star,t}$), and shocks to global risk aversion ($\varepsilon_{\xi,t}$).

2.5.3 Equilibrium dynamics following a shock to global risk aversion

Figure 2.6 depicts the IRFs to a temporary reduction in global risk aversion. In the foreign economy, lower risk aversion induces households to increase current consumption, while firms' faced with higher demand raise their prices. The foreign central bank responds to the ensuing inflationary pressures by gradually raising the nominal interest rate, as per its desire for interest rate smoothing. In the foreign economy, a decline in global risk aversion is therefore associated with rising output, inflation, and nominal interest rate.

This shock affects the domestic economy through its effect on the exchange rate and foreign demand for home goods. *Ceteris paribus*, a decline in global risk aversion induces the financial sector to require lower excess returns on the domestic currency, thereby causing an

[37]Complete exchange rate pass-through implies that nominal exchange rate fluctuations directly translate into changes in home CPI (Eq. 2.14d), exactly because import prices are denominated in the (foreign) producer's currency, and these adjust sluggishly.

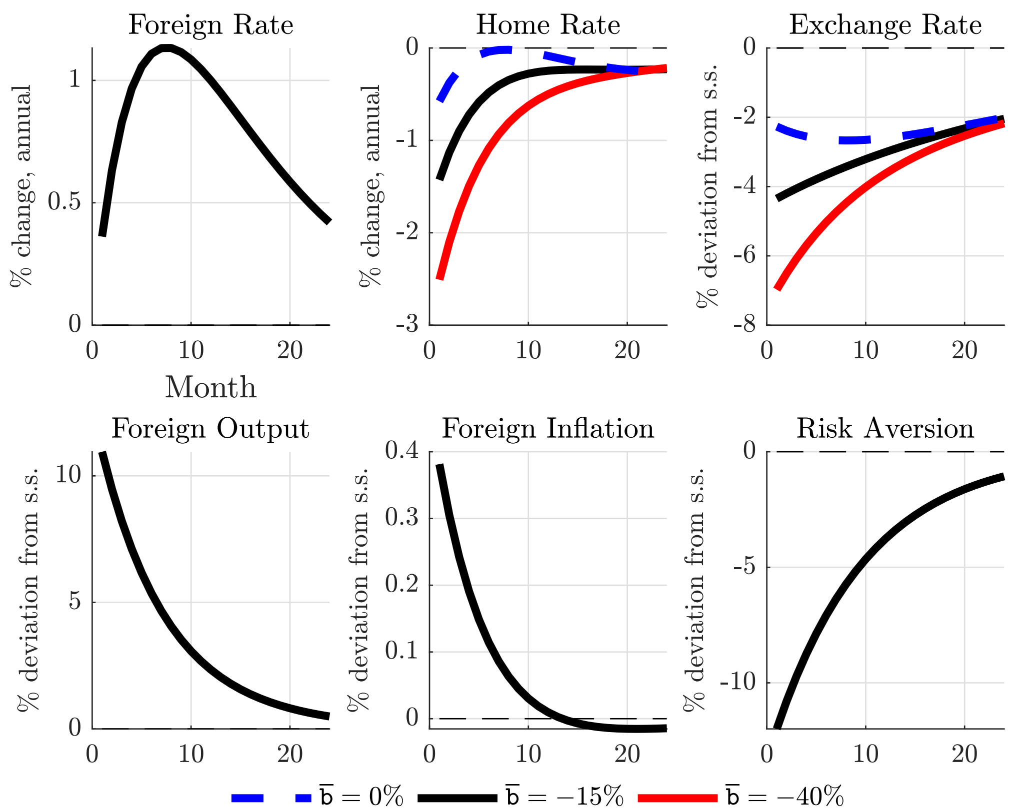


Figure 2.6: Theoretical IRFs to a temporary reduction in global risk aversion

Note: The impulse is an unanticipated 1% reduction in the foreign economy's degree of risk aversion.

instantaneous appreciation of the nominal exchange rate (Eq. 2.12). This effect is reinforced by higher external demand for domestic goods, which improves its net foreign asset to GDP position and reduces the degree to which international financial traders are exposed to home currency risk. These forces dominate over the nominal depreciation implied by the dynamics of the interest rate differential. In fact, the exchange rate response to this shock is largely shaped by the behavior of currency excess returns, as we will show below. In turn, the nominal appreciation of the small economy's exchange rate brings about a contemporaneous fall in import prices (in local currency) which puts downward pressure on domestic CPI inflation (see Eq. 2.14d). In our calibrated model, the deflationary forces implied by lower (domestic-currency) prices of imported goods govern the short-run dynamics of domestic CPI inflation.[38] As a result, the domestic central bank cuts the nominal interest rate. Thus, this shock acts as a favorable supply shock in the SOE, and leads to a procyclical response of the CPI-inflation-targeting monetary authority.

These impulse responses thus provide a natural interpretation of the comovement documented in Figure 2.4, as being driven by global risk aversion shocks.

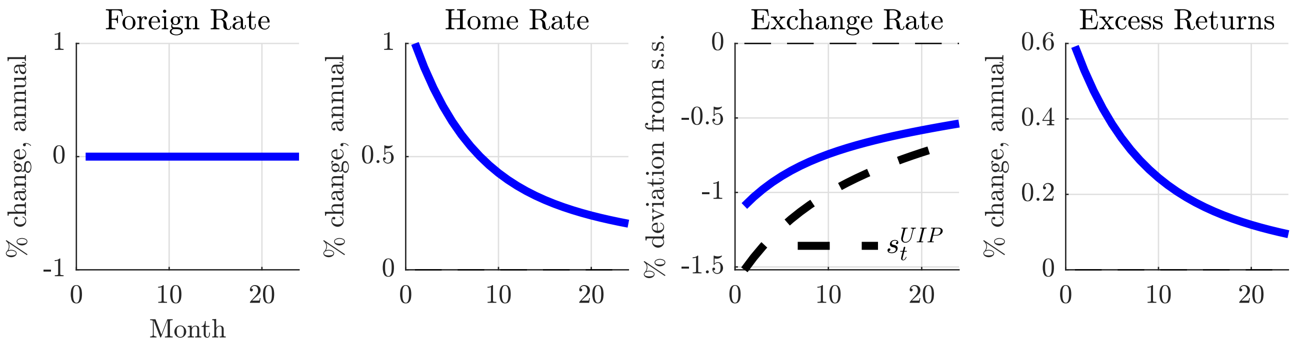
The role of net foreign assets. Figure 2.6 also reports the impulse responses across different levels of the SOE's net foreign asset position to GDP. The blue line reports the impulse responses for an economy with $\mathtt{b} = 0\%$, the highest quartile of our empirical sample, while the red line is for an economy with $\mathtt{b} = -40\%$, the bottom quartile of our empirical sample. In the economy with $\mathtt{b} = 0\%$, changes in global risk aversion only influence exchange rates via the general equilibrium responses of the net foreign asset position and the interest rate differential (see Eq. 2.12). In this economy, a reduction in global risk aversion brings about a lower degree of currency appreciation, relative to the benchmark economy. To the contrary, in the economy with high net foreign debt ($\mathtt{b} = -40\%$), changes in global risk aversion exert a magnified effect relative to the benchmark case (see Eq. 2.12). Their

[38] The domestic component of CPI inflation reflects two opposing forces: higher product demand and adverse expenditure-switching effect due to worsening of the terms of trade.

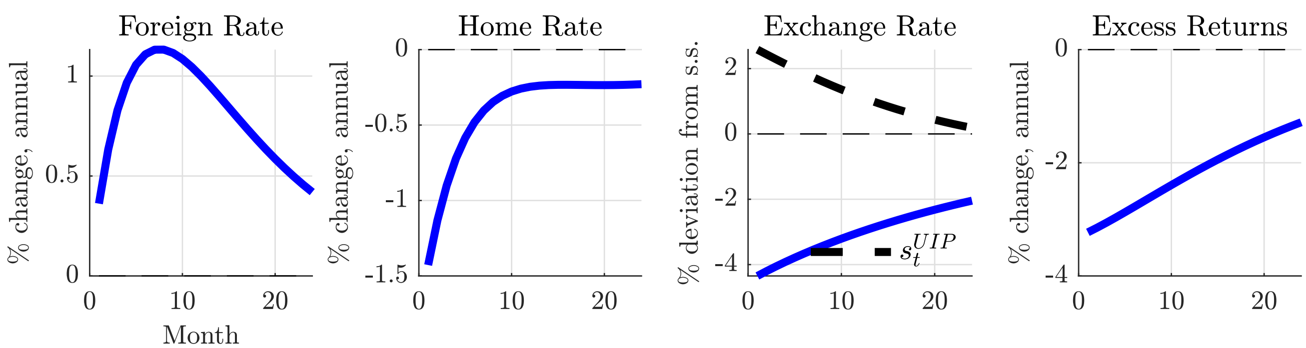
exchange rate appreciates considerably more relative to the benchmark economy following a reduction in global risk aversion. The ranking in domestic interest rate responses across net foreign asset positions mimics the exchange rate responses. In fact, home CPI inflation – the variable that home central banks target – is largely determined by imported inflation in the short run. In Section 2.6, we test these cross-country predictions of our model.

Conditional UIP deviations. Figure 2.7 depicts the theoretical IRFs of a country's exchange rate to transitory domestic and external shocks, which are taken to be the domestic monetary policy shock, and the global risk aversion shock, respectively.

In our model, an unexpected domestic interest rate increase leads to a domestic currency appreciation (2.7a), an exchange rate response that is largely in line with the its UIP-consistent counterpart. Domestic monetary policy shocks (and, in fact, any shocks other than ξ_t) do not affect the level of global risk aversion. The variation in excess returns are due to the equilibrium deterioration of the SOE net foreign asset position, which plays a minor quantitative role in determining exchange rate dynamics. To the contrary, the patterns of excess returns play a predominant role after shocks to global risk aversion (2.7b). The model is thus able to reproduce the patterns of conditional UIP deviations documented in Figure 2.2.



(a) Theoretical impulse responses to a domestic shock (Domestic monetary contraction)



(b) Theoretical impulse responses to an external shock (Reduction in global risk aversion)

Figure 2.7: Conditional properties of domestic and external shocks

Note: Domestic shocks are normalized to deliver a 1% impact increase in the home interest rate, while external shocks are normalized to deliver a 1% increase in the Fed Funds rate at one-year horizon. Excess returns are one-year ahead expected excess returns.

2.6 Net foreign assets and exchange rate dynamics

In the model presented in Section 2.5, the net foreign asset position of the SOE governs the transmission of global risk aversion shocks, and substantially shapes exchange rate properties. In this section, we verify some predictions of the model along untargeted moments of our sample of SOEs.

Exchange rate responses to external shocks, across NFA/GDP. In Figure 2.8, we report the empirical impulse responses to an external shock of the first ($\texttt{b} \leq 40\%$) and last ($\texttt{b} \geq 0\%$) NFA/GDP quartile. The ranking of responses of interest rates, exchange rates, and excess returns across NFA positions conform with the qualitative predictions of our model (see Figure 2.6). Countries with larger net foreign debt to GDP exhibit larger interest rate, exchange rate, and expected excess return responses relative to countries with a relatively balanced NFA/GDP position. This evidence favors the idea that external imbalances play an important role in the international transmission of external shocks. This is a natural feature of a model in which global risk aversion shocks are the key source of external variation in exchange rates.

Exchange rate properties, across NFA/GDP. Because the NFA/GDP position governs a SOE exposure to global risk aversion shocks, SOEs with different NFA/GDP position should feature different exchange rate properties. To begin with, a country's NFA/GDP position affects its degrees of exchange rate volatility. In Figure B.2a, we report the standard deviation of exchange rate changes across NFA/GDP quartiles, along with the values predicted by the model. In line with the model's prediction, exchange rate volatility is higher for countries with a relatively higher net foreign debt.

Because global risk aversion shocks impart a different comovement between interest rates

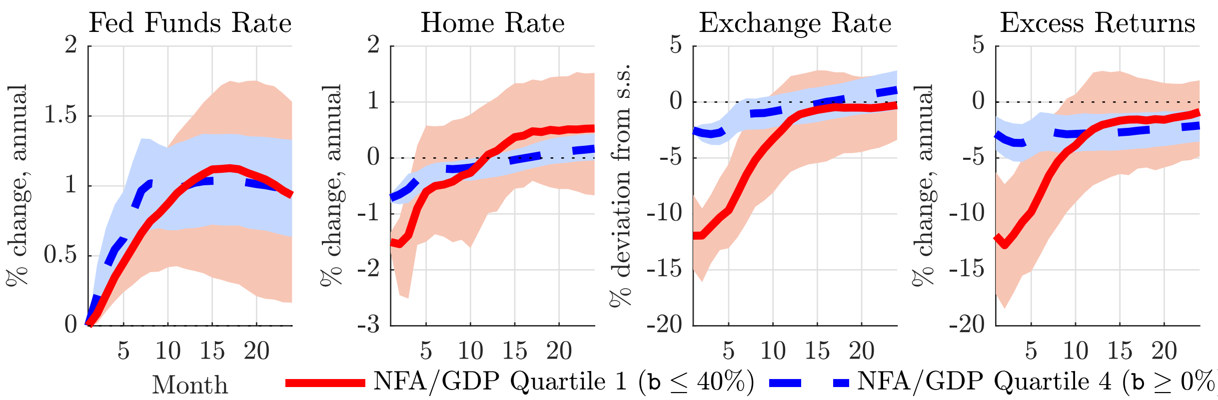
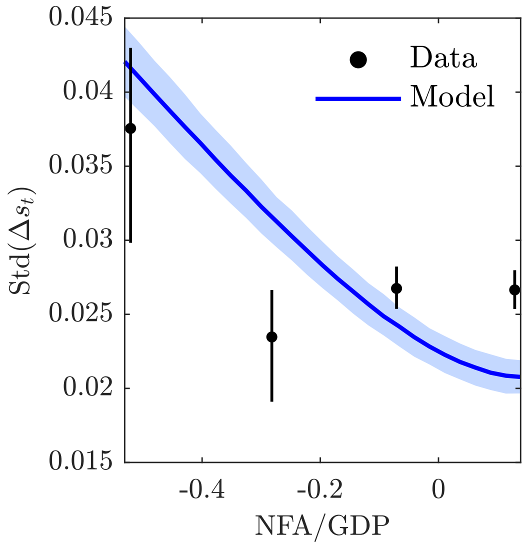


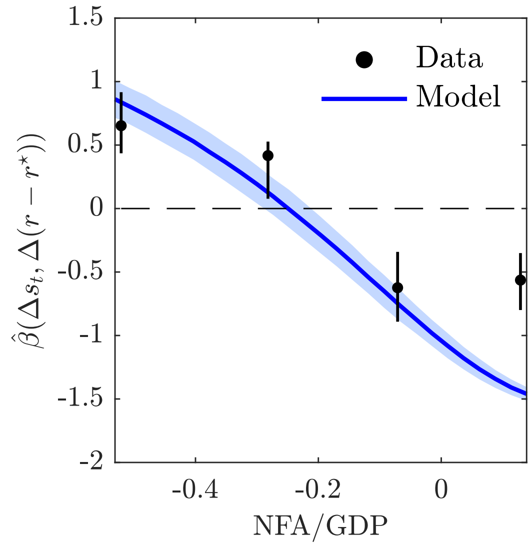
Figure 2.8: Empirical impulse responses to an external shock, by NFA position

Note: The lines denote median IRFs by group of countries with corresponding 90% confidence intervals from 1000 bias-corrected bootstrap replications of the reduced-form VAR. Countries are grouped depending on their average NFA/GDP position. The shock is normalized to deliver a 1% increase in the Fed Funds rate at one-year horizon.

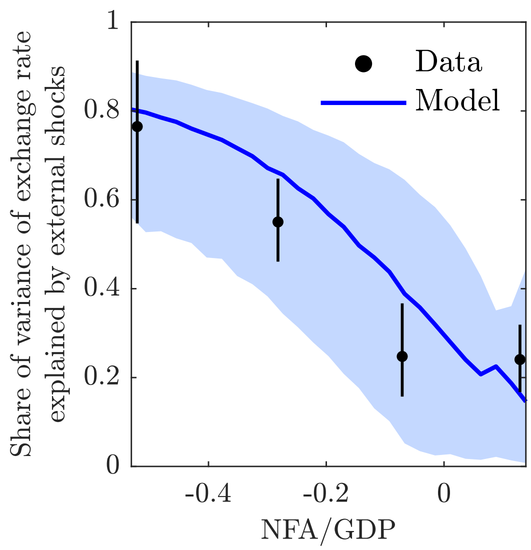
and exchange rates relative to other shocks, SOEs with different NFA/GDP position should feature a different comovement between exchange rate and interest rate differential. To test this prediction, we run $\Delta s_t = \alpha + \beta\Delta(r_t - r_t^\star) + u_t$ and report $\hat{\beta}$, our measure of comovement, both in the data and the model in Figure B.2b. The pattern of empirical comovement that we find in the data aligns well with the prediction of the model. For countries with large net foreign debt the regression coefficient is significantly positive, while it is significantly negative for countries with a relatively balanced NFA/GDP position.



(a) Standard deviation of exchange rate innovations



(b) Comovement between interest rate differentials and exchange rates



(c) Exchange rate exposure to external shocks

Figure 2.9: Exchange rate properties across NFA/GDP positions

Note: The figure reports moments of the data by NFA/GDP quartile, along with the respective values estimated on model-simulated data. For each NFA/GDP quartile, we report the median estimate along with 90% confidence intervals. The solid lines are the median estimates from a Monte Carlo simulation of the model across different values of steady-state NFA/GDP (`b`). In the Monte Carlo simulation, we perform 1000 repetitions with 400 observations per repetition. The shaded areas are the 90% confidence intervals.

Our model predicts that a country's net foreign asset position determines its exposure to external shocks. If global risk aversion shocks account for the bulk of external variation, a country with a small NFA position should be more insulated from external shock relative to a country with a large one. A natural measure of exposure to external shocks is the fraction of the forecast error variance of a country's exchange rate explained by external disturbances. In Figure 2.9c, we report a country's exposure to external shocks across NFA/GDP, along with the values predicted by our model. Countries with a higher ratio of net foreign liabilities to GDP indeed tend to be more exposed to external shocks, in a way that is quantitatively in line with our model's prediction. We take this as evidence that a country's external imbalance is a key determinant of exposure to external disturbances. This feature obtains naturally in a model in which exchange rates are predominantly driven by global risk aversion shocks.

2.7 Revisiting the "exchange rate response puzzle"

One of our main empirical findings is that the domestic variation in exchange rates is largely consistent with UIP (Figure 2.2a). This feature is also present in our model, where only global risk aversion shocks generate large UIP deviations. In our model economy, a domestic monetary contraction brings about an increase in the home rate and an exchange rate appreciation, in line with standard open economy models (Figure 2.7a). This conclusion contrasts with the evidence in Hnatkovska et al. (2016) that the domestic currency tends to depreciate in response to a monetary tightening, especially in developing and emerging economies. This evidence, labeled "the exchange rate response puzzle," is primarily based on recursive identification schemes within VARs. In this section, we show that recursive identification strategies are bound to confound the endogenous response of domestic variables to external shocks with the effect of domestic shocks. We then show that this misspecification problem is the source of the "exchange rate response puzzle."

A recursive identification scheme. The recursive identification scheme obtains as a special case of the identification approach we proposed in Section 2.2, when the D matrix is the identity matrix. In the context of a recursive identification, the exclusion restrictions consist in assuming that the impact matrix is lower triangular, that is

$$u_t = \begin{bmatrix} a_1 & 0 & 0 \\ a_2 & a_3 & 0 \\ a_4 & a_5 & a_6 \end{bmatrix} \widetilde{\varepsilon}_t \tag{2.15}$$

which is estimated with the Cholesky decomposition of Σ.

From Eq. (2.2), the restrictions on the impact matrix A_0 imply that the Federal Funds rate can respond contemporaneously only to its own innovations which are captured by the first element of the vector $\widetilde{\varepsilon}_t$. The policy controlled interest rate of the SOE is not allowed to react on impact to movements in the nominal exchange rate while it can respond to unanticipated movements in the Federal Funds rate. The second element element of the vector of structural shocks $\widetilde{\varepsilon}_t$ is thus typically interpreted as the monetary policy shock of the SOE. In this context, a domestic monetary policy shock influences the policy rate of the

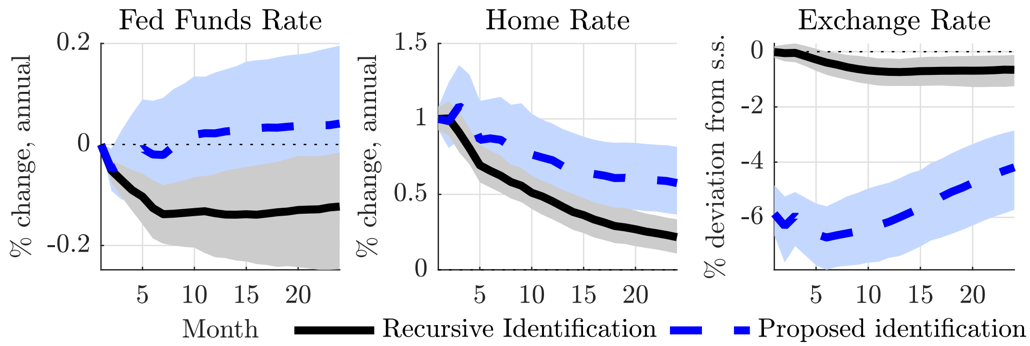


Figure 2.10: Empirical responses to a domestic shock across identification approaches

Note: The shaded areas are the 90% confidence intervals from 1000 bias-corrected bootstrap replications of the reduced-form VAR.

SOE (and possibly the exchange rate) contemporaneously, has no effect on the Federal Funds rate *contemporaneously*, and leaves the response of the Federal Funds rate unrestricted in the months following the shock.

Before discussing the estimated exchange rate response to monetary policy, we ask whether the identified monetary policy shocks are consistent with the assumptions of a SOE. To this end, Figure 2.10 depicts the impulse responses of the three variables in the baseline VAR to a domestic monetary policy shock obtained under recursive ordering, along with the impulse responses to a domestic shock obtained following our proposed identification approach. Under a recursive identification, a contractionary domestic monetary policy shock leads to a significant and persistent decline of the Federal Funds rate, the external variable of our VAR.

There are two possible interpretations of the results in Figure 2.10. First, the U.S.

economy, and, in turn, the Federal Reserve, may respond to disturbances that originate in SOEs, and in particular to their monetary policy innovations. Second, monetary policy in SOEs may respond to external shocks that affect the world interest rate with some delay. While both interpretations are valid in principle, we note that the first interpretation is both contrary to conventional wisdom and inconsistent with the very premise of a *small* open economy by which domestic shocks do not alter world interest rates and incomes.

We thus subscribe to the second interpretation, and argue that the domestic shocks identified through recursive schemes partly capture the endogenous response of domestic central banks to external shocks that influence the Federal Funds rate with some delay. These are the set of shocks that we identified as *anticipated* external shocks in Section 2.2. In addition, we note that these results question the applicability of the common block exogeneity restriction. In the context of the baseline VAR, block exogeneity is equivalent to setting the coefficients on domestic variables in the Federal Funds rate equation to zero. Under the null of no anticipated external shocks, these coefficients are in fact zero. However, if anticipated effects exist, as documented in Figure 2.10, these coefficients are not zero, and applying block exogeneity would be equivalent to imposing a counterfactual restriction.[39] While block exogeneity implies a restriction on the reduced-form parameters of the VAR, our proposed identification approach imposes a restriction on the propagation of shocks: domestic shocks have no effect on the world interest rate.

[39] It is important to stress that the above statements are conditional on the information set spanned by the variables included in the VAR.

Comparison between identification schemes. What is the relation between the shocks identified using a recursive identification and the ones identified with our proposed approach? By combining equations (2.1) and (2.4) one can show that

$$\widetilde{\varepsilon}_t^{mp} = \gamma_1 \varepsilon_t^{\star\star} + \gamma_2 \varepsilon_t^{SOE} \tag{2.16}$$

where $\widetilde{\varepsilon}_t^{mp}$ is the domestic monetary policy shock under a recursive identification, whereas $\varepsilon_t^{\star\star}$ and ε_t^{SOE} are the anticipated external shock and the domestic shock identified under the proposed alternative identification, respectively. Equation (2.16) implies the following. If the restrictions underlying a recursive identification were correct, both identification strategies would recover exactly the same set of shocks. In that case, the estimated value of γ_1 would be zero. However, if anticipated external shocks exist and spill over into the SOE (that is, if estimated $\gamma_1 \neq 0$), standard recursive identification schemes fail to correctly recover the true monetary policy shock. The empirical findings highlighted in Section 2.3 indicate that anticipated external shocks are important, and produce a comovement between domestic interest rates and exchange rates that is opposite from the one implied by domestic shocks. These observations imply that conflating domestic and external sources of exchange rate fluctuations can lead to incorrect inference about the effects of domestic shocks.

We argue that these observations point to the source of recent puzzling evidence on the exchange rate response to domestic monetary policy shocks. In particular, Hnatkovska et al. (2016) documented that the domestic currency tends to appreciate in advanced countries but depreciates in developing and emerging countries in response to a monetary tightening, evidence primarily based on recursive identification schemes within VARs. In Figure B.1, we show that the exchange rate response puzzle disappears after accounting for the effects of anticipated external shocks: in most countries, a monetary policy contraction is associated with a significant appreciation of the nominal exchange rate. Instead, a puzzle arises under a recursive identification scheme because it commingles domestic and external shocks, which give rise to opposite comovements between interest rates and exchange rates (see Figure 2.2).

Examining the performance of identification schemes. We examine the performance of our empirical approach, by estimating a three-variable system identical to the baseline empirical specification on model generated data. We show that our empirical approach correctly separates domestic and external sources of exchange rate variation, whereas a recursive VAR scheme reproduces the exchange rate puzzle.

Figure B.2a in Appendix B.2 indicates that the IRFs produced by our proposed identification approach correctly disentangle the different sources of variation. In fact, the identified domestic shock maps closely into the domestic monetary policy shock (the only domestic shock in our model), while the (anticipated) external shock maps into the global risk aversion shock. Figure B.2b in Appendix B.2 also presents the IRFs implied by a recursive identification presented above. The recursive VAR fails to correctly capture the exchange rate response to a domestic monetary policy innovation. In contrast to the theoretical response, the recursive VAR suggests that a policy-induced interest rate increase triggers a nominal depreciation. In addition, the monetary policy shock series identified under the recursive scheme predicts significant changes in the Federal Funds rate, as documented empirically in Figure 2.10. This happens exactly because the recursive scheme conflates the independent variation in the domestic interest rates and its endogenous response to changes in global risk aversion.

2.8 Conclusions

The exchange rate is at the core of the international transmission mechanism, and a large literature is concerned with understanding the nature of its fluctuations. In this paper, we investigated what role domestic and external shocks play in shaping exchange rate dynamics in SOEs. Using an agnostic decomposition approach, we find that one external shock drives a considerable part of the variation in exchange rate, and, especially, UIP deviations.[40] Moreover, this external shock is significantly correlated with movements in global risk aversion, and connected to U.S. economic fluctuations. We illustrated that these empirical comovements can be interpreted as the equilibrium of a two-country SOE model with international financial market imperfections. In our model, global risk aversion shocks are the main driver of exchange rates and UIP deviations, and a country's net foreign asset position governs their international transmission. Our evidence accords well with our model's predictions, suggesting that external imbalances are important in explaining exchange rate dynamics and their exposure to external shocks.

[40]To the contrary, UIP largely holds conditional on domestic shocks.

Appendix A

What are the Sources of Boom-Bust Cycles?

A.1 Unconditional Spectral Density

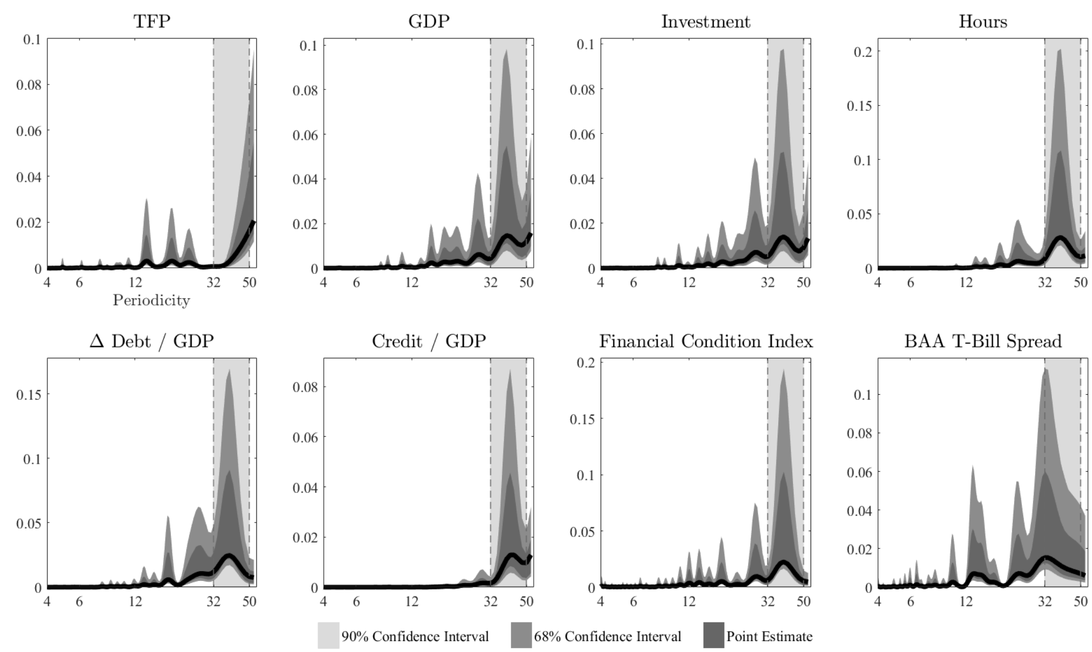


Figure A.1: Unconditional spectral density of quarterly and seasonally adjusted U.S. macroeconomic and financial variables from 1981 to 2018.

Note: All variables are stationarized using Band-Pass filter excluding periodicities above 100 quarters. Confidence intervals are computed following the procedure described in Beaudry et al. (2019).

A.2 Spectral density from model simulated data

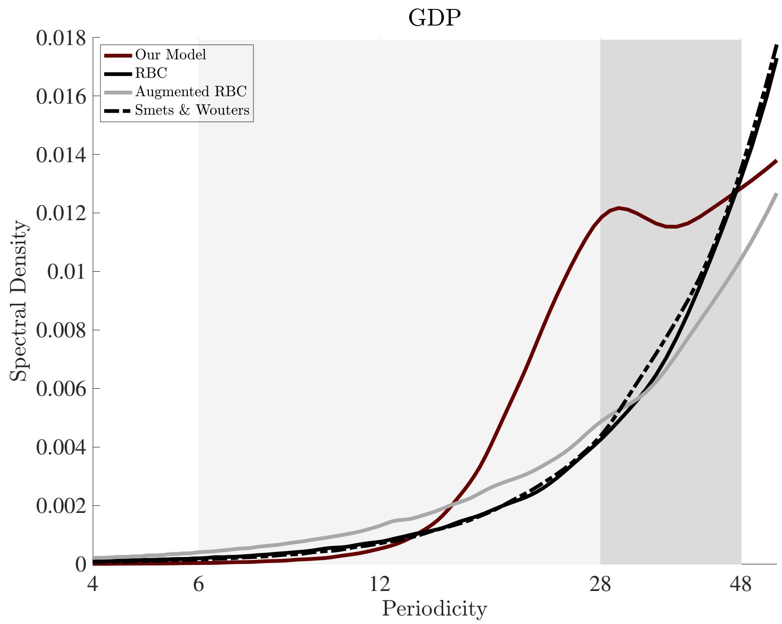


Figure A.2: Mean unconditional spectral density of GDP

Note: Monte Carlo simulation using various standard models and our model (red line). Simulated data are deterended using a band-pass filter that removes fluctuations at periodities greater than 100 quarters.

A.3 Robustness checks on technology Shocks

Figure A.5 reports impulse responses together with conditional spectral densities implied by a technology shock for the baseline specification presented in Figure 1.3 and a series of robustness checks. In particular, RC 1 and RC 2 are the first and the second robustness check where variables are linearly and quadratically detrended, respectively. RC 3 is the third robustness check where TFP is controlled using 8 lags of TFP, the first 2 principal components and news shocks. RC 4 is the last robustness check where we use different number of lags and principal component when we estimate LP impulse responses.

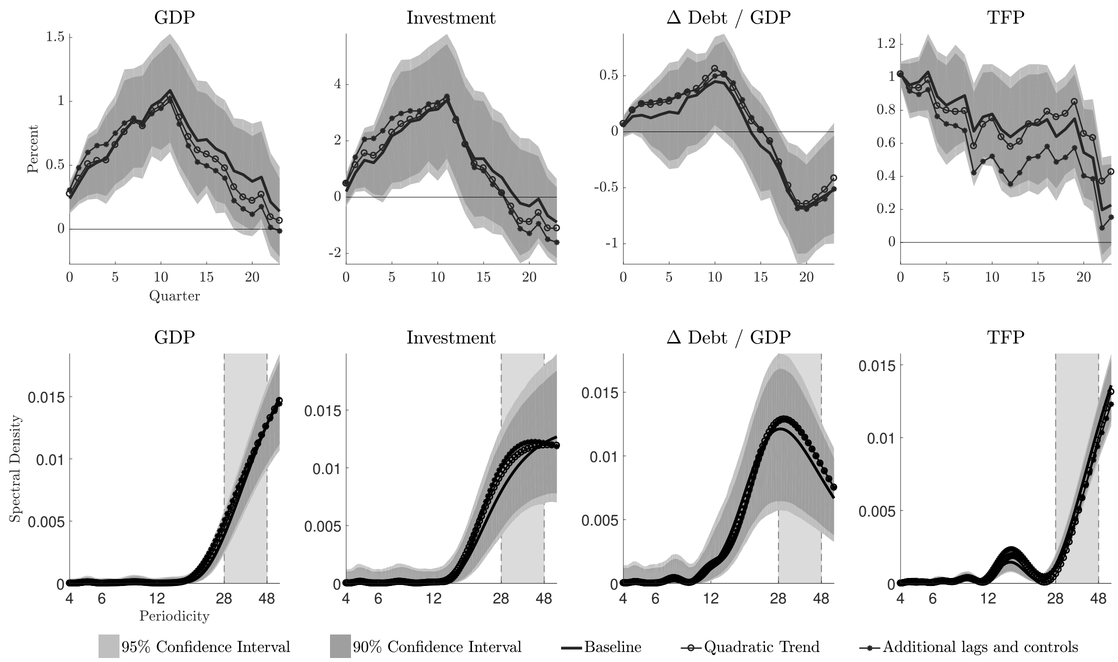


Figure A.3: Impulse responses and conditional spectral densities implied by a technology shock.

Note: Point Estimates is the baseline specification presented in Figure 1.3. RC 1 and RC 2 are the first and the second robustness check where variables are linearly and quadratically detrended, respectively. RC 3 is the third robustness check where we add more controls when we estimate a technology shock. RC 4 is the last robustness check where we use different number of lags and principal component when we estimate LP impulse responses.

A.4 Robustness checks on expectation shocks

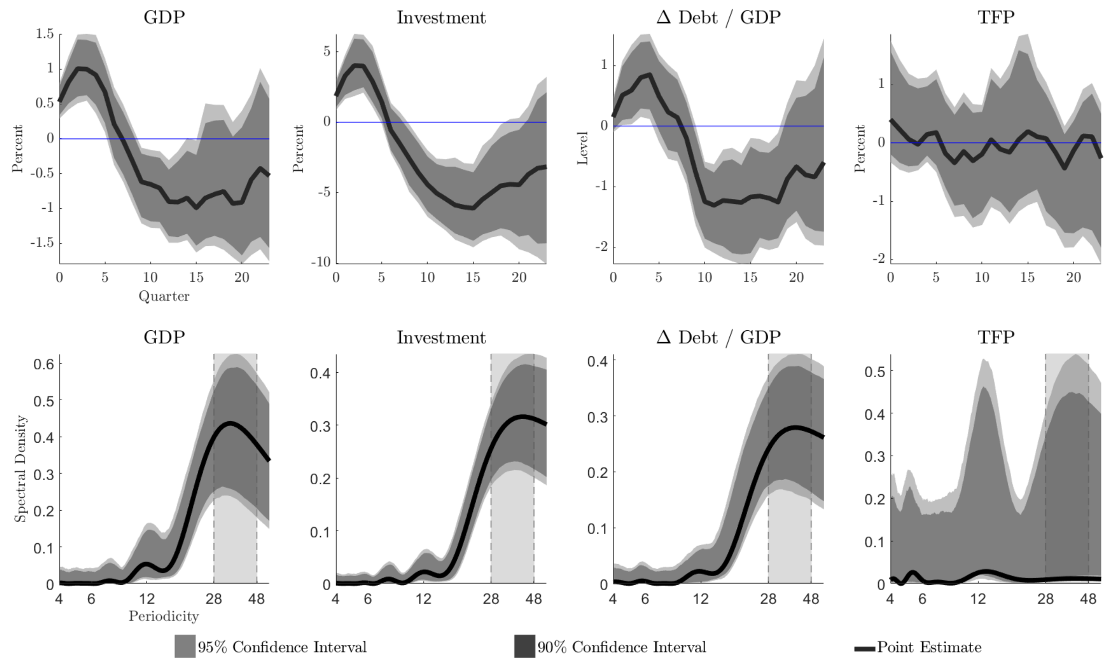


Figure A.4: Impulse responses to an expectations shock.

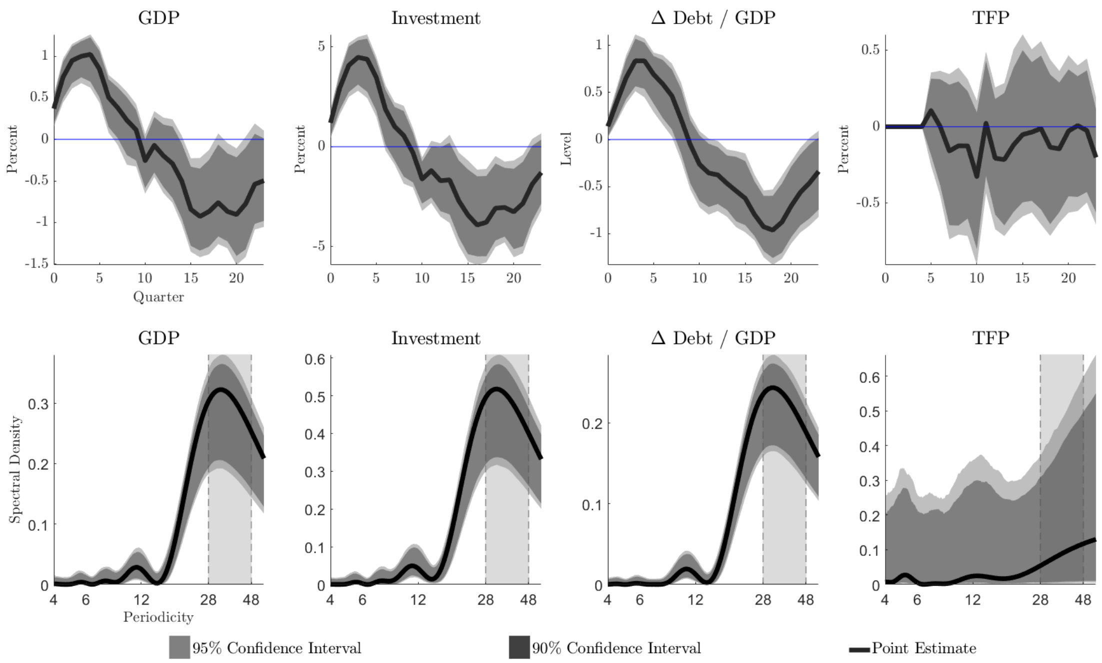


Figure A.5: Impulse responses to an expectations shock starting in 1967.

A.5 Local Projections

To estimate LP impulse responses we follow standard techniques as firstly introduced by Jordà (2005). Given the stationary series y_t and shock ε_t, impulse responses can be estimated as follows,

$$y_{t+h} = \theta_h \varepsilon_t + \sum_{j=1}^{J} \left[\delta_j \varepsilon_{t-j} + \lambda_j y_{t-j} + \gamma_j x_{t-j} \right] + \nu_{t+h} \quad \text{for } h = 0, 1, \ldots, H \tag{A.1}$$

where θ_h represents response of y_t to shock ε_t at horizon h and x_t are additional controls which in our estimation represent principal components from a large dataset of macroeconomic variables.

A.5.1 Inference

Following Kilian and Kim (2011) we estimate confidence interval using the block bootstrap procedure. As emphasized by Kilian and Kim (2011), we opt for this approach because the error term in the local projections regressions is most likely serially correlated. The LP impulse response estimator for horizon h depends on the tuple,

$$\mathcal{T}_h = [y_{t+h} \quad \varepsilon_t \quad \varepsilon_{t-1} \quad \dots \quad \varepsilon_{t-J} \quad y_{t-1} \quad \dots \quad y_{t-I}] \tag{A.2}$$

To preserve the correlation in the data, build the set of all $\mathcal{T}_h$ tuples for $h = 0, 1, \dots, H$. For each tuple $\mathcal{T}_h$, employ the following procedure:

1. Define $g = T - l + 1$ overlapping blocks of $\mathcal{T}_h$ of length l.[1]

2. Draw with replacement from the blocks to form a new tuple $\mathcal{T}_h^b$ of length T.

3. Estimate θ_h^b from $\mathcal{T}_h^b$ using LP estimator.

4. Repeat 1. to 3. B (≥ 2000) times and select confidence intervals.

A.6 Variance Decomposition

Variance decomposition is estimated following Gorodnichenko and Lee (2017). In particular, we define the population share of variance explained by the future innovations in ε_t to the

[1]Notice that $l = (T - I - J + 2)^{\frac{1}{3}}$ is defined following Berkowitz, Birgean and Kilian (1999). Results are not sensitive to alternative choices of l.

total variations in the unpredictable component of y_{t+h} as,

$$v_h = \frac{\sigma_\varepsilon^2 \sum_{i=0}^{h} \theta_i}{Var(f_{t+h|t-1})} \tag{A.3}$$

where $Var(\varepsilon_t) = \sigma_{\varepsilon_t}^2$ and θ_i are LP estimators. Moreover $f_{t+h|t-1}$ can be estimated from the following regression,

$$y_{t+h} = \sum_{j=1}^{J} \delta_j \varepsilon_{t-j} + \sum_{i=1}^{I} \lambda_i y_{t-i} + \sum_{q=1}^{Q} \gamma_q x_{t-q} + f_{t+h|t-1} \tag{A.4}$$

where x_{t-q} represents a vector of additional controls.

Since the estimator v_h does not guarantee estimates to be between 0 and 1, we use the following estimator,[2]

$$\tilde{v}_h = \frac{\sigma_\varepsilon^2 \sum_{i=0}^{h} \theta_i}{\sigma_\varepsilon^2 \sum_{i=0}^{h} \theta_i + Var(\nu_{t+h} - \sum_{i=0}^{h-1} \theta_i x_{t+h-i})} \tag{A.5}$$

where ν_{t+h} is coming from the LP regression,

$$y_{t+h} = \theta_h \varepsilon_t + \sum_{j=1}^{J} \delta_j \varepsilon_{t-j} + \sum_{i=1}^{I} \lambda_i y_{t-i} + \nu_{t+h}. \tag{A.6}$$

A.6.1 Inference

To estimate confidence intervals for $\tilde{v}_h$, we directly use the non-parametric confidence intervals estimated for θ_i. In particular, use simulated θ_i^b to estimate,

$$\tilde{v}_h^b = \frac{\sigma_\varepsilon^2 \sum_{i=0}^{h} \theta_i^b}{\sigma_\varepsilon^2 \sum_{i=0}^{h} \theta_i^b + Var(\nu_{t+h} - \sum_{i=0}^{h-1} \theta_i^b x_{t+h-i})} \tag{A.7}$$

and select confidence intervals.

[2] See Gorodnichenko and Lee (2017) for a detailed description.

A.7 Conditional Spectral Density and Cyclicality Test

Consider the case where stationary variable y_t is explained by two shocks: $\varepsilon_{1,t}$ and $\varepsilon_{2,t}$. In this case, y_t can be represented with the following infinite moving average,

$$y_t = \sum_{h=0}^{\infty} \theta_{1,h}\varepsilon_{1,t-h} + \sum_{h=0}^{\infty} \theta_{2,h}\varepsilon_{2,t-h} \tag{A.8}$$

Since the estimated impulse responses cannot cover an infinite number of lags consider the truncate moving average,

$$y_t \approx \sum_{h=0}^{H} \theta_{1,h}\varepsilon_{1,t-h} + \sum_{h=0}^{H} \theta_{2,h}\varepsilon_{2,t-h} \tag{A.9}$$

Since we are interested in the conditional cyclicality implied by the two shocks, we focus on the conditional moving average,

$$y_{k,t} \approx \sum_{h=0}^{H} \theta_{k,h}\varepsilon_{k,t-h} \quad \text{for} \quad k = 1, 2. \tag{A.10}$$

where $y_{k,t}$ represents the realized value of y_t only conditional on shock $\varepsilon_{k,t}$ for $k = 1, 2$.

Conditional spectral densities are parametrically estimated by taking the Fourier transform of the estimated truncated moving average. Estimators are,

$$s_k(\omega) \approx \left[\sum_{h=0}^{H} \theta_{k,h} e^{ih\omega}\right] \sigma_k^2 \left[\sum_{h=0}^{H} \theta_{k,h} e^{-ih\omega}\right] \quad \text{for} \quad k = 1, 2. \tag{A.11}$$

where $\omega \in (0\ \pi]$ represents frequencies, $i = \sqrt{-1}$, $\theta_{k,h}$ is the LP estimator, and σ_k^2 is a standard estimator for $Var(\varepsilon_{k,t})$.[3]

[3]Notice that for estimating $s_k(\omega)$ we need to build a grid for $\omega \in (0\ \pi]$. Although results are not sensitive to different grid size, in our main results grid is 0.001 in order to guarantee a precise estimate to ten-year frequencies.

A.7.1 Inference

Similarly to what we have done for the variance decomposition, to estimate confidence intervals for $s_k(\omega)$, we directly use the non-parametric confidence intervals estimated for θ_h. In particular, use simulated θ_h^b to estimate,

$$s_k^b(\omega) \approx \left[\sum_{h=0}^{H} \theta_{k,h}^b e^{ih\omega}\right] \sigma_k^2 \left[\sum_{h=0}^{H} \theta_{k,h}^b e^{-ih\omega}\right] \quad \text{for} \quad k = 1, 2. \tag{A.12}$$

and select confidence intervals.

A.7.2 Test

1. Filter each variable you want to test using a Band-Pass filter which excludes frequencies below 2 and above 100.

2. Estimate the autoregressive parameter ρ_y implied by this stationary variable using standard regression techniques.

3. Simulate - for each variable y - B (≥ 2000) $AR(1)$ processes with persistence parameter ρ_y fed with normally distributed random disturbances.[4]

4. For each simulated series estimate its disturbances, impulse response coefficients with LP estimator θ_h and conditional spectral density via $s_k(\omega)$ where k is the estimated innovation from each simulated $AR(1)$ process.

[4]This simulated series has the same length of the data used in the empirical section. Since our sample start slightly after 1980 then we have about 150 observations.

5. Following Canova (1998) and Beaudry et al. (2019) we test if the estimated conditional spectral densities for shocks ε_t ($\hat{s}_\varepsilon(\omega)$) are indistinguishable from the ones derived from the simulated $AR(1)$ process ($\hat{s}_a(\omega)$).
   - Notice that $H_0 : \hat{D}_\varepsilon = \hat{D}_a$ and $H_1 : \hat{D}_\varepsilon > \hat{D}_a$
   - $\hat{D}_k = \hat{s}_k(\omega_1)/\hat{s}_k(\omega_2)$
   - $\omega_1 \in (\pi/40, \pi/28)$ and $\omega_1 \in (\pi/72, \pi/48)$
6. Test statistic is estimated as follows
   - Define $\hat{D}_k^b = \hat{s}_k^b(\omega_1)/\hat{s}_k^b(\omega_2)$ as the simulation of $\hat{D}_k$ from $\hat{s}_k^b$.
   - Estimate, for each b, $\hat{\zeta}^b = \hat{D}_\varepsilon^b - \hat{D}_a^b$ as the difference between the simulation for $\hat{D}_\varepsilon^b$ and $\hat{D}_a^b$.
   - P-value is the number of $\hat{\zeta}^b > 0$ over the total number of simulations B.

	GDP	Investment	ΔDebt / GDP	TFP
Expectation Shock	3.64%	4.82%	2.24%	28.4%
Technology Shock	28.52%	5.54%	0.1%	89.84%

Table A.1: P-values for the test of a local peak in the spectral density implied by expectation shocks (first row) and technology shocks (second row).

A.8 Proof of Theorem 1

Cyclical dynamics obtain if at least two roots of the loglinearized deterministic version of the model are stable, complex and conjugate. Under equilibrium determinacy the model possess only one stable root, therefore the model does not generate cyclical dyanmics. Indeterminacy of equilibria is associated with at least an additional stable root, thus allows for the existence of complex dynamics. The loglinearized deterministic version of the model can be written as

$$\begin{pmatrix} 2\kappa d & \frac{\tau\beta\omega}{1-\tau+\tau\beta} \\ 1-\beta & \beta-\omega \end{pmatrix} \begin{pmatrix} \hat{d}_{t+1} \\ \hat{y}_{t+1} \end{pmatrix} = \begin{pmatrix} \frac{2\kappa d}{1+\mu\gamma} & M \\ 0 & 1-\omega \end{pmatrix} \begin{pmatrix} \hat{d}_t \\ \hat{y}_t \end{pmatrix} \tag{A.13}$$

where

$$M \equiv \frac{\tau\beta\omega}{1-\tau+\tau\beta} - \gamma\frac{1-\mu}{1+\gamma\mu}\left(\omega - 1 + \frac{1}{(1-\theta)(1-n)}\right) \tag{A.14}$$

With no adjustment cost of dividends, that is κ equal to zero, the dynamics of dividends is irrelevant for the evolution of y_t implying that the two eigenvalues of the system cannot be conjugate.

A.9 Data

Following We define the after-tax model-consistent labor wedge Λ as the log difference between the MRS and MPL:

$$\Lambda_t = \log(MPL_t) - \log(MRS_t)$$

where

$$MRS_t = \frac{u_3(c_t, c_{t-1}, 1-n_t)}{u_1(c_t, c_{t_1}, 1-n_t)} \frac{1+T_t^c}{1-T_t^n} = \alpha \frac{(c_t - \iota c_{t-1})^{\omega}}{(1-n_t)^{\omega_2}} \frac{1+T_t^c}{1-T_t^n}$$

and

$$MPL_t = (1-\theta)\frac{y_t}{n_t}.$$

In order to empirically construct the labor wedge we use the same data by Zhang (2018).

Table A.2: Details on aggregate US data

Variable	**Source and Construction**	**Transform**
TFP	Utilization-adjusted total factor productivity (dtfp_util) by San Francisco Fed	Cumulated
GDP	Real gross domestic product (GDPC1) by FRED	Logarithmic
Investment	Gross domestic investment (GDPIC1) plus consumption of durables (PCDGCC96) by FRED	Logarithmic
Δ Debt	Flow of debt securities and loans for the nonfinancial business sector (BOGZ1FA144104005Q) by FRED	Seasonally-adjusted level
Consumption	Consumption of non-durables (PCNGC96) plus consumption of services (PCESVC96) by FRED	Logarithmic
Hours	Hours of all persons for the nonfarm business sector (HOANBS) by FRED	Logarithmic
Credit	Total credit to private non-financial sector (QUSPAM770A) by FRED	Logarithmic
GZ Credit Spread	Measured of credit spread by Gilchrist and Zakrajšek (2012) available on Simon Gilchrist's website	Level
Financial Condition Index	Chicago Fed National Financial Condition Index (NFCI) by FRED	Level
BAA T-Bond Spread	Moody's seasoned Baa corporate bond yield relative to yield on 10-year treasury constant maturity (BAA10Y) by FRED	Level

Note: Seasonally-adjusted transformation is the 7-term Henderson filter.

Appendix B

Shocks and Exchange Rates in Small Open Economies

B.1 Dataset

- Nominal exchange rates (s_t, monthly): the preferred measure of exchange rates are official exchange rates. If these are not available, we use period average market rates, or period average principal exchange rates. The main data source is the International Financial Statistics (IFS) compiled by the International Monetary Fund (IMF).
- Policy-controlled interest rates (r_k, monthly): These rates are measured in the data as the period average T-bill rates, the closest to the overnight interbank lending rates. If these are not available, discount rates, or money market rates are used. The main data source is the International Financial Statistics (IFS) compiled by the

International Monetary Fund (IMF).

- U.S. policy-controlled interest rates ($r^{\star}$, monthly): This rate is measured by the Federal Funds rate.

- Exchange rate regimes: these are determined according to the historical exchange rate classification in Reinhart and Rogoff (2004), recently updated by Ilzetzki et al. (2017). A country is deemed to have a flexible exchange rate regime if, in a given year, its exchange rate was either (i) within a moving band that is narrower than or equal to +/2 percent; or (ii) was classified as managed floating; or (iii) was classified as freely floating; or (iv) was classified as freely falling in Reinhart and Rogoff (2004). We follow Hnatkovska et al. (2016) in including high-income OECD countries irrespective of their exchange rate classification.

- U.S. industrial production (monthly)

- U.S. CPI inflation (monthly)

- Chicago Board Options Exchange Volatility Index (VIX, monthly)

- Net foreign asset positions to GDP (annual): Updated and extended version of dataset constructed by Lane and Milesi-Ferretti (2007).

- Data used for information sufficiency test (monthly)

Country	Time period	Country	Time period
Australia	1974:1-2010:11	Austria	1974:1-1998:12
Belgium	1974:1-1998:12	Brazil	1999:2-2007:12
Canada	1974:1-2010:11	France	1974:1-1998:12
Germany	1975:7-1998:12	Indonesia	1997:8-2007:12
Italy	1977:3-1998:12	Japan	1974:1-2010:11
Korea, Rep. of	1997:12-2007:12	Mexico	1995:1-2007:12
New Zealand	1978:1-2010:11	Norway	1974:1-2009:5
Philippines	1997:7-1999:11	South Africa	1995:3-2007:12
Sweden	1974:1-2010:11	Switzerland	1980:1-2010:11
Thailand	2001:2-2007:12	United Kingdom	1974:1-2010:10

Table B.1: List of countries in the dataset

B.2 Additional tables and figures

	P-value of F-statistic						
	P.C. =1	*P.C. = 3*	*P.C. =5*		*P.C. =1*	*P.C. = 3*	*P.C. =5*
Germany	0.27	0.51	0.61	Indonesia	0.99	0.99	1.00
Canada	0.24	0.57	0.68	Brazil	0.72	0.97	0.87
Italy	0.21	0.17	0.33	South Africa	0.42	0.82	0.44
France	0.88	0.93	0.47	Korea	0.09	0.36	0.10
Japan	0.30	0.19	0.22	Mexico	0.74	0.52	0.55
United Kingdom	0.83	0.81	0.47	Philippines	0.80	0.97	0.82

Table B.2: Information sufficiency test (cf. Forni and Gambetti, 2014)

Notes: The table reports the p-values of the F-statistic of a regression of the identified anticipated external shock on up to 5 principal components (P.C.) of a large data set capturing all the relevant U.S. macroeconomic information, described in Appendix B.1.

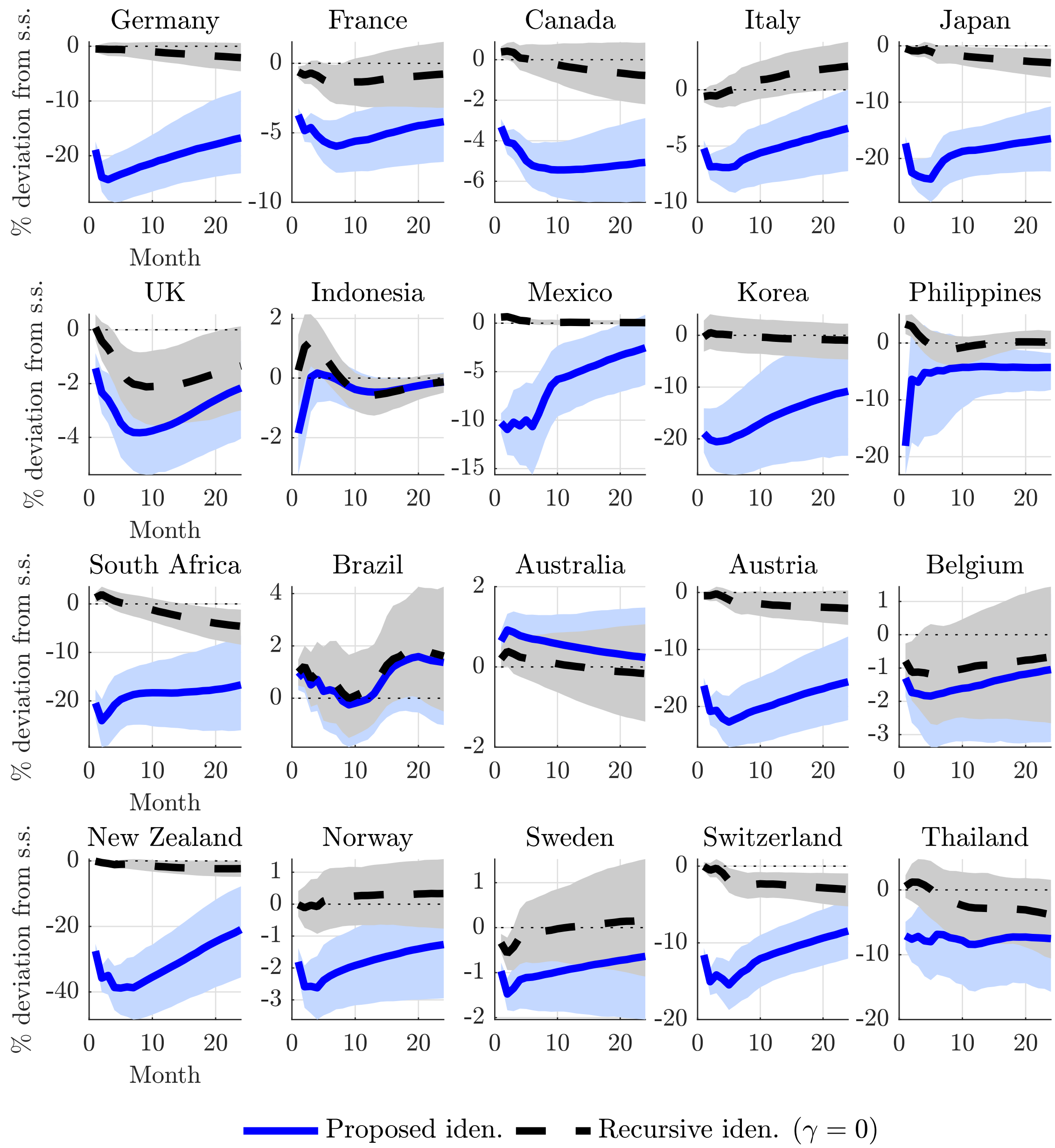
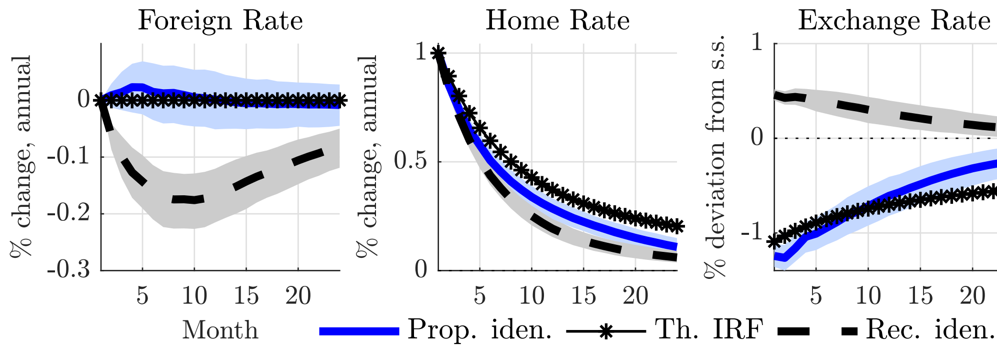
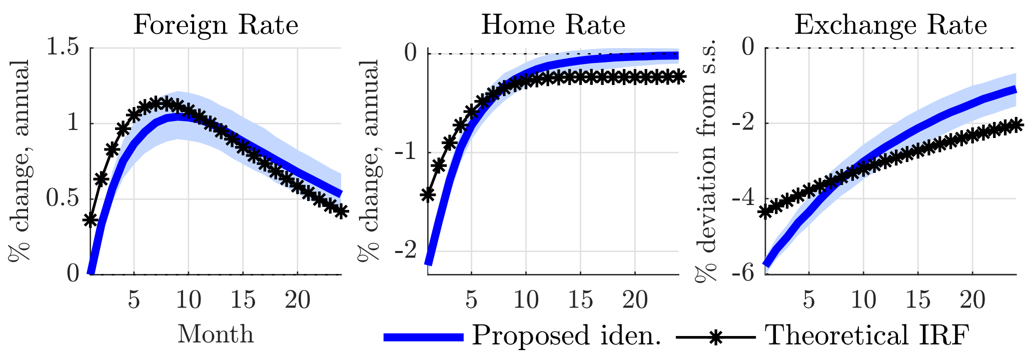


Figure B.1: Exchange rate response to a domestic shock across identification approaches

Note: The blue solid lines are the estimated exchange rate IRFs to domestic shock from the baseline three-variable VAR identified using our proposed identification. The black dashed lines are the estimated exchange rate IRFs to domestic shock from the baseline three-variable VAR identified using a recursive scheme. The shaded areas are the 90% confidence intervals from 1000 bias-corrected bootstrap replications of the reduced-form VAR. Impulse responses are normalized to deliver a 1% impact increase in the domestic interest rate.



(a) Domestic variation



(b) External variation

Figure B.2: Model and Monte Carlo estimated IRFs: three-variable VAR

Note: The black starred line shows the theoretical IRF from the model presented in Section 2.5. Panel B.2a reports the theroretical IRFs to a domestic monetary policy shock, while Panel B.2b reports the theretical IRFs to a global risk aversion shock. The solid lines are the average estimated IRF from a Monte Carlo simulation with 45 repetitions (countries) and 150 observations per repetition. The shaded areas are the 90% confidence intervals from 1000 bias-corrected bootstrap replications of the reduced-form VAR. In Panel B.2a both the recursive identification scheme ($\gamma_1 = 0$) and our proposed alternative are estimated on model-generated data.

B.3 Additional model details

B.3.1 Traders's decision problem

This section shows that a CRRA utility has a mean-variance representation. The problem of the international trader reads as follows:

$$\max_{d_{t+1}} \mathrm{E}_t \left[\frac{(T^\star + \widetilde{R}_{t+1} d_{t+1})^{1-\omega_t^\star}}{1-\omega_t^\star} \right] = \mathrm{E}_t \left[\frac{\exp\left\{ (1-\omega_t^\star) \log(T^\star + \widetilde{R}_{t+1} d_{t+1}) \right\}}{1-\omega_t^\star} \right] \tag{B.1}$$

where $T^\star$ is such that $(T^\star + \widetilde{R}_{t+1} d_{t+1}) > 0$.

Take second order Taylor expansion around $\widetilde{R} = 0$:

$$\begin{aligned} \log(T^\star + \widetilde{R}_{t+1} d_{t+1}) &\approx \log(T^\star) + \frac{d_{t+1}}{T^\star} \widetilde{R}_{t+1} - \frac{d_{t+1}^2}{2\,(T^\star)^2} \widetilde{R}_{t+1}^2 \\ &\approx \log(T^\star) + \frac{d_{t+1}}{T^\star} \widetilde{R}_{t+1} - \frac{d_{t+1}^2}{2\,(T^\star)^2} \operatorname{Var}_t(\widetilde{R}_{t+1}) \end{aligned}$$

where $\widetilde{R}_{t+1}^2$ is replaced by the conditional variance of $\widetilde{R}_{t+1}$.[1,2] Then Eq. (B.1) is approximated by:

$$\max_{d_{t+1}} \mathrm{E}_t \left[\frac{\exp\left\{ (1-\omega_t^\star) \left(\log(T^\star) + \frac{d_{t+1}}{T^\star} \widetilde{R}_{t+1} - \frac{d_{t+1}^2}{2(T^\star)^2} \operatorname{Var}_t(\widetilde{R}_{t+1}) \right) \right\}}{1-\omega_t^\star} \right]$$

[1]Note that $\mathrm{E}_t[\widetilde{R}_{t+1}]^2 \approx 0$.

[2]As the time interval shrinks, the higher order terms that are dropped from (B.1) become negligible relative to those that are included, and the deviation of $\widetilde{R}_{t+1}^2$ from $\operatorname{Var}_t(\widetilde{R}_{t+1})$ also become negligible. In particular in the limit of continuous time the approximation is exact and can be derived using Ito's Lemma.

$$\approx \max_{d_{t+1}} \exp\left\{(1-\omega_t^\star)\left(\log(T^\star) - \frac{d_{t+1}^2}{2\,(T^\star)^2}\operatorname{Var}_t(\widetilde{R}_{t+1})\right)\right\} \mathrm{E}_t\left[\exp\left\{(1-\omega_t^\star)\left(\frac{d_{t+1}}{T^\star}\widetilde{R}_{t+1}\right)\right\}\right].$$

Assume normal distribution of $\widetilde{R}_{t+1}$, then

$$\approx \max_{d_{t+1}} \log(T^\star) - \frac{d_{t+1}^2}{2\,(T^\star)^2}\operatorname{Var}_t(\widetilde{R}_{t+1}) + (1-\omega_t^\star)\frac{d_{t+1}^2}{2\,(T^\star)^2}\operatorname{Var}_t(\widetilde{R}_{t+1}) + \frac{d_{t+1}}{T^\star}\mathrm{E}[\widetilde{R}_{t+1}]$$

$$\approx \max_{d_{t+1}} \mathrm{E}_t[\widetilde{R}_{t+1}]d_{t+1} - \frac{\omega_t^\star}{2T^\star}\operatorname{Var}_t(\widetilde{R}_{t+1})d_{t+1}^2$$

In equilibrium, the individual trader's asset decisision reads

$$d_{t+1} = \frac{T^\star\,\mathrm{E}_t[\widetilde{R}_{t+1}]}{\omega_t^\star \operatorname{Var}_t(\widetilde{R}_{t+1})}$$

Without loss of generality, we set $T^\star = 1$. Then, aggregating over the m_t measure of traders, the overall demand for domestic bonds from traders is

$$\widetilde{D}_{t+1} = \frac{m_t}{\omega_t^\star}\frac{\mathrm{E}_t\,\widetilde{R}_{t+1}}{\operatorname{Var}_t(\widetilde{R}_{t+1})}$$

which is Eq. (2.10) in the text.

B.3.2 Model equilibrium equations

Besides each country's Phillips Curve, the model's equilibrium equations in levels are given by:

$$\beta^\star\,\mathrm{E}_t\left[\left(C_{t+1}^\star\right)^{-\omega^\star\exp(\omega_{t+1}^\star)}\frac{R_t^\star}{\Pi_{t+1}^\star}\right] = \left(C_t^\star\right)^{-\omega^\star\exp(\omega_t^\star)}$$

$$\frac{R_t^\star}{R^\star} = \left(\frac{R_{t-1}^\star}{R^\star}\right)^{\rho_R}\left(\frac{\Pi_t^\star}{\Pi^\star}\right)^{(1-\rho_R)\phi}\exp\left(\varepsilon_{r^\star,t}\right)$$

$$\beta\,\mathrm{E}_t\left[(C_{t+1})^{-\omega}\,\frac{R_t}{\Pi_{t+1}}\right] = (C_t)^{-\omega}$$

$$\frac{R_t}{R} = \left(\frac{R_{t-1}}{R}\right)^{\rho_R}\left(\frac{\Pi_t}{\Pi}\right)^{(1-\rho_R)\phi}\exp\left(\varepsilon_{r,t}\right)$$

$$\Pi_t = (\Pi_{H,t})^{1-\lambda}\left(\frac{S_t}{S_{t-1}}\Pi^\star_t\right)^{\lambda}$$

$$Y_t = Q_t^{\frac{\theta\lambda}{1-\lambda}}\left\{(1-\lambda)C_t + \lambda Q_t^{\theta}C^\star_t\right\}$$

$$\frac{{}^{B_{t+1}}\!/\!{}_{P_{H,t}Y_t}}{R_t} - {}^{B_t}\!/\!{}_{P_{H,t-1}Y_{t-1}}\frac{1}{\Pi_{H,t}{}^{Y_t}\!/\!{}_{Y_{t-1}}} = 1 - Q_t^{-\frac{\lambda}{1-\lambda}}\frac{C_t}{Y_t}$$

$$S_t = Q_t\frac{P_t}{P^\star_t}$$

$$-{}^{B_{t+1}}\!/\!{}_{P_{H,t}Y_t} = \frac{\mu\,\mathrm{E}_t\left(R_t - R^\star_t\frac{S_{t+1}}{S_t}\right)}{\mathrm{Var}_t\left(R_t - R^\star_t\frac{S_{t+1}}{S_t}\right)}$$

B.3.3 Model solution

We can represent the model outlined in Appendix B.3.2 as the following system of equations:

$$\mathrm{E}_t\left[f(X_{t+1})\right] = 0$$

where X_{t+1} contains all the variables in the model (including variables dated at time t and $t-1$) and f has as many rows as endogenous variables in the model. The risky steady state

(Coeurdacier et al., 2011) is obtained by taking a second-order approximation of f around $\mathrm{E}_t\, X_{t+1}$:

$$\Phi\left(\mathrm{E}_t\, X_{t+1}\right) = f\left(\mathrm{E}_t\, X_{t+1}\right) + \mathrm{E}_t\left[f''\left[X_{t+1} - \mathrm{E}_t\, X_{t+1}\right]^2\right]$$

where f'' is also evaluated at $\mathrm{E}_t\, X_{t+1}$. The risky steady state, $\mathtt{X}$, is then characterized by $\Phi\left(\mathtt{X}\right) = 0$, and the second moments $\mathrm{E}_t\left[f''\left[X_{t+1} - \mathrm{E}_t\, X_{t+1}\right]^2\right]$ are generated by the linear dynamics around $\mathtt{X}$.

The model's solution thus consists in a log-linear approximation around a risky steady state that is consistent with the second moments generated by the log-linear dynamics around it. This is achieved through an iterative algorithm, along the lines of Coeurdacier et al. (2011).

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