Working Papers

FEDERAL RESERVE BANK OF BOSTON™

No. 18–10

The Dollar During the Global Recession: US Monetary Policy and the Exorbitant Duty

Vania Stavrakeva and Jenny Tang

Abstract:

We document that during the Global Recession, US monetary policy easings triggered the "exorbitant duty" of the United States, the issuer of the world's dominant currency, by causing a dollar appreciation and a transfer of wealth from the United States to the rest of the world. This dollar appreciation runs counter to the predictions of standard macroeconomic models and works through two channels: (i) a flight-to-safety effect which lowered the expected excess returns of holding safe US government debt relative to foreign debt and (ii) lowered expected future inflation in the United States relative to other countries. We show that the signaling channel of monetary policy, whereby US policy easings are perceived to signal weaker future growth, can reconcile the novel empirical findings that we document.

Keywords: exchange rates, currency risk, risk premia, monetary policy, forward guidance, Federal Reserve information, interest rates, Global Financial Crisis

JEL Classifications: E52, F31, G01

Vania Stavrakeva is an assistant professor of economics at the London Business School. Her email address is vstavrakeva@london.edu. Jenny Tang is an economist in the research department at the Federal Reserve Bank of Boston. Her email address is jenny.tang@bos.frb.org.

Nikhil Rao provided invaluable research assistance on this project. We thank the participants and discussants at numerous conferences and seminars. We are also grateful to Emmanuel Farhi, Domenico Giannone, Gita Gopinath, Stephen Morris, Ali Ozdagli, Paolo Pesenti, Ricardo Reis, Hélène Rey, Kenneth Rogoff, Eric Swanson, and Michael Weber for useful comments.

This version: October, 2018

1 Introduction

An influential literature has documented that serving as the dominant currency not only confers an "exorbitant privilege" but also exacts an "exorbitant duty" (see Gourinchas and Rey (2007a), Gourinchas and Rey (2007b), Gourinchas, Rey, and Govillot (2018)).[1] The United States provides insurance to the rest of the world in exchange for having the privilege of paying low interest rates on its safe dollar-denominated assets. The usual narrative is that, during global crises, the dollar appreciates due to a flight to safety which increases the demand for safe dollar-denominated assets such as US government bonds. The dollar appreciation raises the net asset valuations of foreign countries, which tend to invest in safe dollar-denominated assets and borrow in risky non-dollar currencies. This valuation channel represents a sizable wealth transfer from the United States to the rest of the world. Gourinchas, Rey, and Govillot (2018) estimate that the net asset valuation loss that the United States suffered during the global financial crisis and the Eurozone debt crisis was 14 percent and 16 percent, respectively, of US GDP.[2]

This literature leaves unanswered an important set of questions: what types of exogenous shocks trigger this exorbitant duty appreciation of the dollar and through what channels do these shocks operate. Answering these questions will improve our understanding of what determines the relative values of foreign currencies and is of first-order importance for policymakers in the United States and around the world, given the dollar's role as the dominant currency.[3] This paper makes two contributions that, taken together, offer answers to these questions.

The first contribution is evidence documenting that expansionary US monetary policy shocks during the Global Recession—dated from 2008:Q4 to 2012:Q2—triggered the exorbitant duty associated with being the hegemon currency. That is, US policy easings during the Global Recession caused the dollar to appreciate and led to a significant transfer of wealth from the United States to the rest of the world. We find important heterogeneity in the dollar's response to US monetary policy shocks. More precisely, a surprise US rate cut induced a larger appreciation against currencies for which the dollar acts as a better hedge; that is, currencies that tend to depreciate more strongly against the dollar when US real output growth falls. We use a novel decomposition of the exchange rate response to show

[1]See also Gourinchas, Rey, and Truempler (2012), Gourinchas and Rey (2014), and Gourinchas, Rey, and Sauzet (forthcoming).

[2]The valuation losses can be perceived as a transfer from the United States to the rest of the world and these losses capture both the effect due to the dollar's movement and the change in asset valuations in local currency.

[3]See Goldberg and Tille (2008), Shin (2012), Ivashina, Scharfstein, and Stein (2015), Casas et al. (2017), Gopinath (2016), and Gopinath and Stein (2018) for evidence on the dollar's use as the dominant currency in trade and financial transactions.

that US monetary policy caused the dollar to appreciate through two main channels.[4]

The first channel is a flight-to-safety effect resulting from US monetary policy easings during the Global Recession. Surprise US interest rate cuts lowered the expected future excess returns, or the currency risk premium, investors required to hold US government bonds and to be short the bonds of other countries. We present additional evidence consistent with this flight-to-safety effect by showing that US policy easings led to higher risk aversion during the Global Recession.[5] Moreover, we show that the cross-currency heterogeneity with respect to the dollar's response to US monetary policy surprises can be entirely explained by the heterogeneous response of expected future currency risk premia to US monetary policy.

This flight-to-safety effect is consistent with the narrative in the "exorbitant duty literature. However, besides this channel, we document evidence of a second important channel that led the dollar to appreciate against other currencies. Expansionary US monetary policy, by lowering the expected future path of US inflation relative to other countries, also caused the dollar's value to increase during the Global Recession.

This paper's second contribution is a theoretical model that shows how the signaling channel of monetary policy can reconcile all of the empirical facts that we document for exchange rates and risk aversion. We present a partial equilibrium model in which the central bank is perceived to have better information regarding the future economic environment. Therefore, forward guidance that is intended to be accommodative also can signal expected adverse future shocks to US GDP growth. The theory is similar in spirit to Tang (2015), Melosi (2017), and Andrade et al. (2018), but our model also introduces currency risk premia using a consumption habits framework akin to Campbell and Cochrane (1999).

Our theory shows that if the direct expansionary effect of promising lower future policy rates—forward guidance that will lower future real interest rates—is overshadowed by the signaling effect of this forward guidance, then real GDP growth expectations will decline. This fall in expected growth heightens expected future risk aversion, thus leading to a flight-to-safety-driven appreciation of the dollar against currencies for which the dollar is a hedge. At the same time, lower US growth expectations also cause the dollar to rise in value by lowering expected future US inflation relative to foreign inflation, provided that agents primarily interpret the economic signal from forward guidance to be about demand shocks. Therefore, the model features both mechanisms that are empirically shown to be the main drivers of

[4]More specifically, based on an accounting identity, we decompose the dollar's response to US monetary policy surprises into the responses of changes in expectations over future excess returns (which includes currency risk premia), relative policy rate paths, and relative inflation rate paths. We do so using estimates of these exchange rate components from Stavrakeva and Tang (2018b), which are based on a VAR approach that disciplines estimates of agents' expectations using survey forecasts.

[5]In the data we proxy for risk aversion by using either the Chicago Board of Trade's Volatility Index (VIX) or the measure constructed by Bekaert, Engstrom, and Xu (2017).

the dollar's response to US monetary policy shocks during the Global Recession. Moreover, the theory also predicts the same cross-currency heterogeneity in the dollar's response and in the strength of the flight-to-safety effect that we empirically document occurred over the Global Recession—namely, that the more that the dollar tends to hold its value against a particular currency in bad times, the more strongly it appreciates and the more pronounced is the flight-to-safety effect against that other currency in response to US monetary policy easings.

For this theory to be consistent with our results regarding how US monetary policy easings during the Global Recession caused the dollar to appreciate in a way that is consistent with fulfilling its exorbitant duty, the signaling effect had to dominate the direct expansionary effect of promising lower rates during this period. Indeed, we confirm empirically that over this crisis period, accommodative US monetary policy led to significantly negative downward revisions in future survey-based US growth expectations. To understand why the signaling effect overshadowed the direct effect of monetary policy during the Global Recession, we examine two sufficient conditions for this dominance that are implied by the model.

First, the existence of the signaling channel requires conducting a type of monetary policy that has the potential to convey information regarding the central bank's forecast for future economic conditions. From 2008:Q4 to 2012:Q2, a period which coincides with our Global Recession sample, the Federal Reserve used "calendar-based" forward guidance. This type of forward guidance promised low rates until some later date, but the announcements were generally accompanied with an explanation that the FOMC expected that weak economic conditions for the foreseeable future would warrant such an extended period of low rates. This type of forward guidance is ripe for interpretation as a sign of a deteriorating economy, thereby weakening expectations for growth.[6,7]

Second, the model shows that the signaling effect dominates when uncertainty about economic fundamentals is sufficiently high relative to monetary policy uncertainty. As shown in our model and in Tang (2015) and Melosi (2017), among others, the more uncertain economic agents are about these fundamentals relative to their uncertainty about exogenous monetary policy shocks, the more weight agents will place on central bank policy announcements as indicators of economic fundamentals. Macroeconomic uncertainty—measured both by the estimates of Jurado, Ludvigson, and Ng (2015) and the dispersion of real GDP growth forecasts from the *Blue Chip Economic Indicators* survey—was particularly high during the

[6]In Stavrakeva and Tang (2018a), we attempt to disentangle the effect of forward guidance versus quantitative easing (QE) on the exchange rate and find that the dollar appreciation in response to the Federal Reserve cutting interest rates over the Global Recession can be attributed to calendar-based forward guidance and not to QE.

[7]This phase of forward guidance was replaced in 2012:Q4 by "threshold-based" forward guidance that gave information mainly about the Fed's policy reaction function.

Global Recession. In contrast, monetary policy uncertainty—measured as the monetary policy subcomponent of the Baker, Bloom, and Davis (2016) policy uncertainty index—was lower during this time relative to other periods.

This is the first paper, to our knowledge, to document that monetary policy easings can trigger the US dollar's exorbitant duty behavior. This evidence has important implications regarding which types of models should be used to inform our understanding of how US monetary policy affects the global macroeconomy. The vast majority of open economy models that study monetary policy and exchange rates ignore the signaling channel of monetary policy and, as a result, imply that the dollar will depreciate, not appreciate, when the Federal Reserve lowers rates during a crisis. Ingrained in the policy debate is the perception that cutting rates should depreciate the currency. During the crisis, the unprecedented expansionary monetary policy undertaken by the Federal Reserve, which many other countries believed would depreciate the dollar, prompted complaints that the United States was engaging in "currency wars" and "competitive devaluation" (see Bernanke (2017)). As we will show, this conclusion is not supported in the data. During the Global Recession, rather than depreciate, the dollar appreciated in a statistically significant way in response to the Federal Reserve's policy easings, which triggered substantial wealth transfers from the United States to the rest of the world.

The outline of the paper is the following. Section 2 reviews the related literature. Section 3 details our empirical strategy. Section 4 presents our results regarding the effect of US monetary policy during the Global Recession on the value of the dollar and its components, on risk aversion, and on net wealth transfers from the United States to the rest of the world. Section 5 presents the model describing the signaling channel of monetary policy and shows that this signaling channel was indeed strong enough during the Global Recession to reconcile the empirical results documented in the preceding sections. Section 6 concludes.

2 Related Literature

First and foremost, this paper is related to the literature on the exorbitant duty pioneered by Gourinchas, Rey, and Truempler (2012) and Gourinchas, Rey, and Govillot (2018). This series of empirical papers documents the fact that the United States, which benefits from the exorbitant privilege of borrowing at low rates, also has the exorbitant duty of providing insurance to the rest of the world. The fact that the dollar is a good hedge currency (i.e., it appreciates during bad times) also implies that, in times of distress, risk-averse investors flock into safe dollar-denominated assets, behavior which appreciates the dollar and triggers a wealth transfer from the United States to the rest of the world.[8] While the narrative is compelling, previous work has not empirically identified an exogenous shock that triggers the exorbitant duty behavior of the dollar.

Using high-frequency identification of monetary policy shocks, this paper documents one such shock—a monetary policy easing shock with a strong signaling component regarding the future of the US economy. In response to such a shock, the dollar appreciates and there are significant net wealth transfers from the United States to the rest of the world. Moreover, we add to the narrative by documenting that the dollar appreciates not only because of a flight-to-safety effect triggered by higher risk aversion, but also due to inflation expectations being relatively lower in the United States as a result of the lower growth expectations. We argue that the inflation channel of the dollar's appreciation will be present only if the shock originates in the United States. Another novel contribution to the literature is the evidence on cross-currency heterogeneity in the exorbitant duty properties of the dollar, which can be explained by the dollar's heterogeneous hedging properties against other currencies.

This paper also presents a partial equilibrium model that can reconcile the empirical facts that we document. The approach taken in our model is related to that in Gourinchas, Rey, and Govillot (2018) in that both models emphasize that receiving a signal of lower expected future GDP growth can lead to higher risk aversion and, thus, trigger a global flight to safety.[9] Gourinchas, Rey, and Govillot (2018) develop a rich general equilibrium environment which matches a number of the facts in the literature related to both the exorbitant privilege and the exorbitant duty of being the hegemon currency.[10] In contrast, the purpose of our partial equilibrium model is to highlight the conditions under which monetary policy easing can trigger the exorbitant duty behavior of the dollar via the strong signaling effect of forward guidance, while also illustrating a microfoundation that matches the wide range of empirical facts that we document.

Maggiori (2017) also presents a model which provides a microfoundation for both the exorbitant duty and the exorbitant privilege of the United States, but the channel which leads to the dollar appreciation is somewhat different. In Maggiori (2017), the reason that

---

[8] See also Gourinchas and Rey (2007b), Gourinchas and Rey (2014), and Gourinchas, Rey, and Sauzet (forthcoming). For evidence on the hedging properties of the dollar, see Maggiori (2013) and the empirical evidence presented in this paper.

[9] Gourinchas, Rey, and Govillot (2018) model a shock that leads to a higher probability of a rare disaster occurring in the future, while we model a monetary policy shock which is interpreted as a strong signal regarding future economic growth.

[10] Gourinchas, Rey, and Govillot (2018) interpret periods when expected stock market volatility is high as those in which the dollar exhibits the exorbitant duty properties based on their finding that net foreign asset positions of the United States tend to fall when the Chicago Board Options Exchange Volatility Index (VIX) is high. In contrast, we estimate that exogenous US monetary policy easing shocks triggered an increase in risk aversion and a transfer of wealth from the United States to the rest of the world during the Global Recession.

the dollar appreciates in times of crisis is due to more resilient demand for US goods relative to the goods produced in the rest of the world, which arises from a shock to relative export costs. In contrast, in our paper, as in Gourinchas, Rey, and Govillot (2018), the mechanism that triggers the exorbitant duty behavior of the dollar operates through higher risk aversion which generates a flight to safety.[11]

There is also a growing theoretical literature that tries to elucidate the source of the exorbitant privilege of the dollar. This privilege has been attributed to the greater ability of the United States to create store-of-value assets relative to developing countries (Caballero, Farhi, and Gourinchas (2008)), to the larger size of the US economy (Hassan (2013), Hassan, Mertens, and Zhang (2016)), and to the lower risk aversion of US consumers relative to consumers in developing countries (Gourinchas, Rey, and Govillot (2018)), a preference that can be rationalized by considering that the US financial sector faces relatively more lax financial constraints (Maggiori (2017)).[12] Our paper is silent on the source of the exorbitant privilege of the dollar and takes as given the cross-country heterogeneity in the composition of net foreign asset portfolios, which contributes to the wealth transfers that we document.

Other related literature is work on asset pricing that studies the "flight-to-safety" or "flight-to-quality" mechanisms. The most relevant empirical paper to our study is Baele et al. (2018) who, unlike us, do not find an exogenous shock that triggers a flight to safety, but instead identify days on which asset prices behaved in ways consistent with a flight to safety. Baele et al. (2018) show that these days also coincide with an increase in risk premia and portfolio rebalancings in which savings flow out of equity funds and into money market funds or government bond funds.[13] In contrast, the shock that we identify also triggers net wealth transfers from the United States to the rest of the world, which is why we interpret the shock as triggering the dollar's exorbitant duty behavior more generally, not just triggering a flight-to-safety effect.[14]

Understanding the linkage between US monetary policy and the exorbitant duty proper-

[11] The models in Farhi and Maggiori (2018) and He, Krishnamurthy, and Milbradt (forthcoming) feature, among other results, multiple equilibria where self-fulfilling expectations can render the dollar a safe asset currency.

[12] More recent contributions in the literature include Chahrour and Valchev (2017) and Gopinath and Stein (2018).

[13] They also find that the Japanese yen, the Swiss franc and, to a lesser degree, the US dollar appreciate on the flight-to-safety dates. See also Beber, Brandt, and Kavajecz (2009) and Baele, Bekaert, and Inghelbrecht (2010), among others, for other empirical papers in this literature.

[14] There are also related theoretical asset pricing papers. Vayanos (2004), Caballero and Krishnamurthy (2008) and Brunnermeier and Pedersen (2009) microfound the flight from risky assets to safe bonds during times of distress. Our model shares with all of these models the feature that some initial shock leads to agents effectively becoming more risk averse, thus triggering the flight to safety or flight to quality. However, these models do not study monetary policy with a strong signaling effect as the initial shock which triggers the flight to safety, and do not model exchange rates and currency risk premia.

ties of the dollar also makes a contribution to the literature on the international spillover effects of monetary policy. More precisely, we uncover a surprising result that is novel to this literature—monetary policy easing may lead to dollar appreciation rather than depreciation when the signaling channel of monetary policy is strong [see Eichenbaum and Evans (1995) and the ensuing literature].[15] Methodologically, we identify monetary policy shocks using high-frequency changes in market-based interest rate expectations in the spirit of Kuttner (2001), Bernanke and Kuttner (2005), and Gürkaynak, Sack, and Swanson (2005).[16] Our model of the signaling effect of monetary policy captures elements explored in greater detail in Tang (2015), Melosi (2017), and Nakamura and Steinsson (2018).[17] Andrade et al. (2018) also presents a model where calendar-based forward guidance serves as a signal about economic fundamentals. Unlike our paper, none of these earlier papers models and studies the effect of monetary policy signaling on exchange rates and currency risk premia.[18]

3 Identifying Effects of Monetary Policy Shocks

To estimate the effect of US monetary policy, our empirical approach is as follows. We obtain monetary policy surprises using changes in interest rate expectations measured in tight time-windows around US monetary policy announcements. We then regress the outcome variables of interest on policy indicators instrumented using these surprise measures. These two-stage least squares (2SLS) estimates can be interpreted as a local projections instrumental variables estimate of the contemporaneous response of the outcome variable to US monetary policy shocks that raise the relevant policy indicator by one unit [see Section 2.3 of Stock and Watson (2018)].

We first describe the data that we use, including measures of US monetary policy surprises, and then provide more details on the empirical specifications. We also present a decompo-

[15] The only other paper that empirically studies the linkage between exchange rates and monetary policy in the recent period at a frequency lower than daily is Rogers, Scotti, and Wright (forthcoming). They find a result consistent with the conventional wisdom that cutting rates in the United States depreciates the dollar. However, they use an identification strategy that precludes a signaling effect of monetary policy.

[16] More recent papers using similar identification methods are Gertler and Karadi (2015), Gilchrist, Zakrajšek, and Yue (2016), Swanson (2017) and Nakamura and Steinsson (2018).

[17] The theoretical literature on the signaling channel of monetary policy goes back to earlier work such as Cukierman and Meltzer (1986), Ellingsen and Söderström (2001), and Berkelmans (2011). Our use of "signaling" refers to what Campbell et al. (2012) call "Delphic forward guidance" or what Nakamura and Steinsson (2018) call the "information effect." This terminology differs from how "signaling" is used in the QE literature to refer to QE actions conveying a commitment to maintaining low future policy rates.

[18] Our paper is also related to the empirical literature that studies the signaling effect of monetary policy during the Global Recession. Campbell et al. (2012), Tang (2015), and Nakamura and Steinsson (2018) find that accommodative monetary policy either lowers or fails to improve expectations of economic conditions over various sample periods that do not align precisely with the Global Recession.

sition of exchange rates changes that we will use to disentangle the channels through which US monetary policy surprises transmit to exchange rates.

3.1 Data Description

In order to identify monetary policy shocks, we rely on the methodology developed in the high frequency monetary policy identification literature—see Kuttner (2001), Bernanke and Kuttner (2005), Gürkaynak, Sack, and Swanson (2005) and the ensuing literature. More precisely, the monetary policy surprises that we use as regressors in the first stage of the 2SLS estimation are changes in interest rate futures prices over a one-hour window ranging from 15 minutes prior to 45 minutes after FOMC statements and quantitative easing (QE) announcements that were made outside of these regular statements.[19] Given that the Global Recession coincides with the zero lower bound (ZLB) period in the United States and with the use of unconventional monetary policy, we use futures prices that capture the whole spectrum of the yield curve in order to measure unconventional policies which were used during this period to impact yields at various maturities. In particular, we use federal funds rate futures expiring three months hence (FF4), eurodollar futures expiring three quarters hence (ED4), and two- and ten-year Treasury note futures expiring in the current quarter.[20] Since the payoffs of these futures contracts depend on the underlying short-term interest rates or Treasury yields that prevail upon settlement of the contracts, changes in these prices can be used to measure changes in interest rate expectations. Since these changes are measured over short time windows occurring around US monetary policy announcements, these changes in expectations only reflect information about current and future policy actions conveyed by these announcements.[21]

Our main outcome variables are the currency exchange rates of nine major countries against the United States: Australia, Canada, euro area, Japan, Norway, New Zealand, Sweden, Switzerland, and the United Kingdom. We also examine US net foreign asset positions measured by the US net international investment position or net valuation losses computed using net foreign asset positions and the current account balance. Lastly, we

[19] The list of QE announcements can be found in Appendix B.3 and was assembled from existing papers including Rogers, Scotti, and Wright (2014), Wu (2014), and Swanson (2017). All of the results are robust to excluding QE announcement dates. This is important to note as we will argue later that the effects we find can be attributed to "calendar-based" forward guidance.

[20] We thank Refet Gürkaynak for providing data on federal funds and eurodollar futures. The results are robust to using different sets of surprises, including ones that exclude measures based on near-term federal funds rate futures.

[21] These futures prices also contain risk premia, but Piazzesi and Swanson (2008) show that taking high-frequency differences in these prices effectively cleans out risk premia that vary at lower frequencies and that the remaining surprise measures are not contaminated by risk premia.

study the responses of US GDP growth forecasts from the *Blue Chip Economic Indicators* survey, the risk aversion estimates from Bekaert, Engstrom, and Xu (2017), and measures constructed using the VIX.

The policy indicators that we instrument are forward interest rates calculated using zero-coupon government bond yields. We consider forward rates at various horizons since unconventional US monetary policies have had different impacts on long- versus medium-term rates over the ZLB period. Policies such as forward guidance or QE have been found to have effects that peak in particular regions of the yield curve.[22] Therefore, rather than choosing interest rates at one particular maturity to act as a policy indicator, as is done in Gertler and Karadi (2015), Rogers, Scotti, and Wright (forthcoming), and other related papers, we examine forward rates at various horizons to more flexibly and agnostically capture the different dimensions of unconventional monetary policy. The exact empirical specifications are presented in the next sub-section.

3.2 Empirical Specification

We perform our analysis at a quarterly frequency and over a period which covers the Global Recession—2008:Q4–2012:Q2. Our sample begins with the contraction in real economic activity that began immediately after the Lehman Brothers collapse and also includes the European debt crisis. The end date coincides with Mario Draghi's mid-2012 speech stating that the European Central Bank would do "whatever it takes" to save the euro, which greatly calmed global financial markets.[23]

We focus on the crisis period for a number of reasons. First, to preview our argument presented in section 5, where we introduce our model, one of the conditions for US monetary policy easings to be interpreted as a signal that triggers the exorbitant duty behavior of the dollar requires that a relatively high level of uncertainty exists about economic fundamentals. This condition was surely met during the Global Recession as it was the worst economic crisis since the Great Depression.[24] Second, in Section C of the Appendix, we present formal structural break tests to show that the relationship between exchange rate changes and

[22]Swanson (2017) finds that the effect of forward guidance typically peaks for yields with maturities of between one and five years, while QE has its greatest impact on longer maturities, particularly 10 years. Greenwood, Hanson, and Vayanos (2016) find that the effects of QE announcements are the largest for one-year forward rates for five and seven years ahead.

[23]Gourinchas, Rey, and Govillot (2018) date the US financial crisis as 2007:Q4–2009:Q1 and the European sovereign debt crisis as 2010:Q4–2012:Q2. We chose to use 2008:Q4 as the beginning of the Global Recession rather than 2007:Q4 for methodological reasons, as this is an important structural break for monetary policy given that the United States entered the zero lower bound in 2008:Q4. Finally, considering the US financial crisis and the European sovereign debt crisis separately is not feasible given the sample size.

[24]We provide evidence on uncertainty in section 5.

changes in the monetary policy indicator reverses in a statistically significant way over the 2008:Q4–2012:Q2 period, relative to the periods before and after the Global Recession.[25]

To estimate the responses of exchange rate changes and its components, we use relative one-year forward rates as the policy indicator. That is, we estimate:

$$\Delta s_{t+1} = \alpha_n^s + \beta_n^{\Delta s_{t+1}} \Delta \tilde{f}_{t+1}^n + error_{t+1}, \tag{1}$$

where s_t is the log exchange rate of currency i per US dollar. Thus, an increase in s_t corresponds to a currency i depreciation against the dollar. The expression $\Delta \tilde{f}_{t+1}^n$ represents the relative one-year forward rate n quarters ahead. Throughout the paper, tildes above variable terms denote the relative quantities defined in terms of country i minus the respective US variable. We estimate this regression for forward rate horizons ranging from $n = 8$ to $n = 36$ quarters ahead. The reason why we focus only on the medium and long ends of the yield curve is due to the fact that over this period US short policy rates were constrained by the ZLB and, as result, forward guidance and QE focused primarily on lowering forward rates at horizons equal to and greater than eight quarters. In this first-stage regression, we allow the coefficients to differ by currency pair by interacting the instruments with pair dummies. This allows us to utilize the cross-sectional variation in how foreign forward rates respond to US monetary policy shocks. In this specification, an estimate of $\beta_n^{\Delta s_{t+1}}$ can be interpreted as the contemporaneous response of the exchange rate change to a US monetary policy shock that raises the corresponding relative forward rate by 1 percentage point (see Stock and Watson (2018) for details on the interpretation). A positive estimate of $\beta_n^{\Delta s_{t+1}}$ then indicates that US monetary policy easing led the dollar to appreciate during the Global Recession.

We use *relative* forward rates as policy indicators in the exchange rate regression because the exchange rate is a relative price between two currencies and, therefore, the responses of other central banks to US monetary policy shocks can play an important part in how these shocks are transmitted to exchange rates. Using the relative forward rate as a policy indicator allows us to control for differences across countries in how other central banks respond to US monetary policy.

For variables other than exchange rates, we follow the empirical monetary policy literature and use US interest rates as the policy indicator. We continue to consider one-year forward rates at different horizons, where $n \geq 8$. Thus, our specification for each variable of interest x is:

$$x_{t+1} = \alpha_n^x + \beta_n^{x_{t+1}} \Delta f_{t+1}^{n,US} + error_{t+1}, \tag{2}$$

[25]For the structural break test we consider data from 1991:Q1 to 2015:Q3.

where a negative estimate of $\beta_n^{x_{t+1}}$ indicates that US monetary policy easing led to an increase in the variable x during the Global Recession.

3.3 Exchange Rate Decomposition

In order to shed light on the channels through which monetary policy affects the exchange rate, we will further disentangle $\beta_n^{\Delta s_{t+1}}$ by decomposing exchange rate changes into components with clear economic interpretations. First, we define the expected excess return from investing in one-quarter, risk-free dollar-denominated bonds and taking a simultaneous short position in one-period, risk-free bonds denominated in currency i. We denote the log expected excess return from this trade as:

$$\sigma_t \equiv i_t^{us} - i_t^i + E_t \Delta s_{t+1}. \tag{3}$$

For convenience, we will use the two terms, "expected excess currency return" and "currency risk premia," interchangeably. The expected excess return may capture not only the currency risk premium but also numerous additional frictions, including the inability of traders to borrow at the risk-free government bond rate, counterparty risk, and binding net worth or value-at-risk constraints. In our empirical work, we do not impose any restrictions on σ_t that limit it to only capturing risk premia. In Section 5, we provide one particular model of σ_t as a currency risk premium.

Using equation (3), the actual change in the exchange rate can be written as:

$$\Delta s_{t+1} = \tilde{\imath}_t + \sigma_t + \Delta s_{t+1} - E_t \Delta s_{t+1}. \tag{4}$$

By iterating equation (3) forward we obtain:

$$s_t = -E_t \sum_{k=0}^{\infty} [\tilde{\imath}_{t+k} + \sigma_{t+k}] + E_t \lim_{k \to \infty} s_{t+k}. \tag{5}$$

First-differencing equation (5) and combining the resulting expression with equation (3) then gives us an expression for the expectational error, $\Delta s_{t+1} - E_t \Delta s_{t+1}$, that we combine with equation (4) to obtain:

$$\Delta s_{t+1} = \tilde{\imath}_t - \underbrace{\sum_{k=0}^{\infty}\left(E_{t+1}\tilde{\imath}_{t+k+1} - E_t\tilde{\imath}_{t+k+1}\right)}_{\varphi_{t+1}^{EH}}$$
$$+\, \sigma_t - \underbrace{\sum_{k=0}^{\infty}\left(E_{t+1}\sigma_{t+k+1} - E_t\sigma_{t+k+1}\right)}_{\sigma_{t+1}^{F}} + \underbrace{E_{t+1}\lim_{k\to\infty} s_{t+k} - E_t\lim_{k\to\infty} s_{t+k}}_{s_{t+1,\infty}^{\Delta E}}. \tag{6}$$

Equation (6) expresses the realized exchange rate changes in terms of the time t relative short rate, the time t expected excess return, and the forward-looking variables that reflect changes in expectations in: (i) the path of relative short-term nominal rates, φ_{t+1}^{EH}, (ii) the path of excess returns, σ_{t+1}^{F}, and (iii) long-run nominal exchange rates, $s_{t+1,\infty}^{\Delta E}$. Stavrakeva and Tang (2018b) show that if the real exchange rate is stationary, $s_{t+1,\infty}^{\Delta E}$ reflects changes in expectations over long-run relative price levels, which equals the path of future relative inflation (country i relative to the United States). Therefore, we will refer to $s_{t+1,\infty}^{\Delta E}$ as the inflation component of the exchange rate change decomposition.

We can now use equation (6) to further decompose $\beta_n^{\Delta s_{t+1}}$. An ordinary least squares (OLS) estimate of coefficient $\beta_n^{\Delta s_{t+1}}$ in equation (1) can be rewritten as a ratio of the sample estimates of a covariance and a variance that can be further decomposed in the following way:

$$\begin{aligned}\hat{\beta}_n^{\Delta s_{t+1},OLS} &= \frac{\widehat{Cov}\left(\Delta s_{t+1}, \Delta\tilde{f}_{t+1}^{n}\right)}{\widehat{Var}\left(\Delta\tilde{f}_{t+1}^{n}\right)} \\ &= \frac{\widehat{Cov}\left(\tilde{\imath}_t - \varphi_{t+1}^{EH}, \Delta\tilde{f}_{t+1}^{n}\right)}{\widehat{Var}\left(\Delta\tilde{f}_{t+1}^{n}\right)} + \frac{\widehat{Cov}\left(\sigma_t - \sigma_{t+1}^{F}, \Delta\tilde{f}_{t+1}^{n}\right)}{\widehat{Var}\left(\Delta\tilde{f}_{t+1}^{n}\right)} + \frac{\widehat{Cov}\left(s_{t+1,\infty}^{\Delta E}, \Delta\tilde{f}_{t+1}^{n}\right)}{\widehat{Var}\left(\Delta\tilde{f}_{t+1}^{n}\right)}.\end{aligned}$$

In the case of a 2SLS estimate, the same expression holds, with the change in relative forward rates being replaced with the fitted value of the relative forward rate change from the first-stage regression.

Given that each of the scaled covariances is a univariate regression coefficient obtained from regressing the exchange rate change components in equation (6) on $\Delta\tilde{f}_{t+1}^{n}$, we can write $\hat{\beta}_n$ in terms of the following regression coefficients:

$$\hat{\beta}_n^{\Delta s_{t+1}} = \hat{\beta}_n^{\tilde{\imath}_t - \varphi_{t+1}^{EH}} + \hat{\beta}_n^{\sigma_t - \sigma_{t+1}^{F}} + \hat{\beta}_n^{s_{t+1,\infty}^{\Delta E}}, \tag{7}$$

where the superscripts of $\hat{\beta}$ denote the dependent variables.

Decomposing the relationship between exchange rate changes and monetary policy using equation (7) will help us understand the channels though which US monetary policy easings triggered the dollar to behave in a way that fulfills its "exorbitant duty".

4 The Exorbitant Duty Effects of US Monetary Policy During the Global Recession

In this section, we present evidence that US monetary policy easing during the Global Recession was a shock that, on average, caused the dollar to appreciate, partially due to a flight-to-safety effect, and triggered wealth transfers from the United States to the rest of the world. These features of accommodative US monetary policy over the 2008:Q4–2012:Q2 period are all consistent with the effects of a shock that triggers the dollar's exorbitant duty behavior (see Gourinchas, Rey, and Govillot (2018)). We also present additional related empirical results that we later reconcile with the theory presented in Section 5.

4.1 Effects on the Dollar

First, we estimate an overall response of the exchange rate to US monetary policy by estimating equation (1) as a panel regression with currency pair fixed effects using 2SLS.

Figure 1 plots the slope coefficients and 90 percent confidence intervals from this regression estimated over the Global Recession. The confidence intervals are computed based on Driscoll-Kraay standard errors which are robust to heteroskedasticity, cross-sectional correlation, and up to four lags of autocorrelation in the errors. The estimates are also reported in the first row of Table 1.

During the Global Recession, we see that for forward rates 12 or more quarters ahead, a US monetary policy shock that decreases US medium- and long-term forward rates relative to foreign forward rates causes a statistically significant dollar appreciation.[26] The first-stage regression F-statistics for the Global Recession period are all greater than 39, far exceeding the a rule-of-thumb threshold of 10 commonly used to detect the presence of weak instruments.

We also estimate equation (1) using OLS to obtain the unconditional relationship (see Figure A-3 and the first row of Table A-1 in the Appendix). The results are qualitatively similar to the 2SLS regressions. The fact that the OLS estimates capture the same patterns

[26]In contrast, the same US monetary policy shock has the opposite effect in the pre- and post-Global-Recession samples (see Figure A-1 in the Appendix). An OLS estimate of the same relationship exhibits the same changes in coefficient estimates across these periods (see Figure A-2 in the Appendix).

implies that US monetary policy shocks are potentially an important driver of the overall comovement between exchange rate changes and changes in relative forward rates.

Next, we examine the cross-country heterogeneity in the response of the dollar to US policy surprises during the Global Recession. Heterogeneous responses are likely, given that the pool of currencies that we consider includes the Swiss franc, the Japanese yen and the euro, which are currencies that have been known to have some hedging properties themselves—i.e. to appreciate during bad times.

Figure 2 plots the regression coefficients estimated for each currency pair for $n = 32$ against a measure of whether the dollar serves as a hedge with respect to each currency—the measure we use is motivated by the model that we present in Section 5.[27] More precisely, we measure this hedging property using the covariance between the respective exchange rate change and US real GDP growth.[28] A negative covariance means that, all else equal, the dollar tends to appreciate *unconditionally* against the respective currency when US real GDP growth falls, thus serving as a hedge against low US growth. This figure shows a clear ordering, where the currencies for which the dollar serves as a hedge are those that lost the most value against the dollar in response to US monetary policy easings during the Global Recession. In contrast, the currencies for which the dollar does not serve as a good hedge either lost little value or gained value against the dollar in response to US policy easings.

Unsurprisingly, we find that the dollar does not have hedging properties against the Swiss franc, the Japanese yen and the euro, meaning that the dollar depreciates unconditionally against these currencies when the US real economy is doing badly (the covariances are all positive, but close to zero for the euro). We observe that the dollar depreciated against the Swiss franc and Japanese yen in response to US monetary policy easing, while it appreciated against the euro, though the estimated regression coefficient is rather small.

Given that Figure 2 makes it clear that pooling all nine currencies will not be appropriate, in the remainder of the paper we present estimates based on panel regressions among two country groups, created given the hedging properties of the dollar with respect to each particular currency captured by the covariance shown in Figure 2. We refer to these two groups as "hedge" and "non-hedge" currencies. The hedge currencies include those of Australia, Canada, Norway, New Zealand, Sweden, and the United Kingdom, while the non-hedge currencies are those of the euro area, Japan, and Switzerland.

Figures 3 and 4 plot the 2SLS regression coefficients from equation (1) for the hedge and non-hedge currency groups. The 2SLS estimates of $\hat{\beta}_n^{\Delta s_{t+1}}$, along with their standard errors, are presented in the first rows of Tables 2 and 3 for the two groups of countries. For

[27]The results are robust to using different $n \geq 12$.

[28]We estimate these covariances using data starting in 1990, but the results are robust to using only the period over the Global Recession.

the hedge group, all of the coefficients for forward rate horizons $n \geq 12$ are positive and statistically significantly different from zero during this subperiod. For the non-hedge group, the estimates are negative but are not significantly different from zero for most horizons, which is not surprising given the smaller sample and the fact that the euro appreciated rather than depreciated in response to US policy easings during this period.

4.1.1 Effects on the Components of Exchange Rate Changes

This subsection presents the results from the decomposition of $\hat{\beta}_n^{\Delta s_{t+1}}$ in equation (7) using the terms in the exchange rate change decomposition given by equation (6) that are computed in Stavrakeva and Tang (2018b). To calculate the exchange rate change components, we need the expectations of three-month government bond rates, inflation rates, and exchange rates at all forecast horizons greater than zero. These expectations are obtained by modeling short rates, inflation, and exchange rates using a flexible VAR process. To ensure that the VAR-implied expectations capture private sector expectations well, the VAR is estimated with additional equations that ensure that the VAR-based forecasts remain close to the survey forecasts of exchange rates, interest rates, and inflation obtained from *Blue Chip* and *Consensus Economics*. One can think of this specification as a way to interpolate and extrapolate forecasts at horizons not reported in the surveys. This method has primarily been used to fit bond yields [see Kim and Wright (2005), Kim and Orphanides (2012), Piazzesi, Salomao, and Schneider (2015), and Crump, Eusepi, and Moench (2016)], but to our knowledge, this technique has not been applied to the study of exchange rates.[29]

Using the exchange rate components from equation (6) that are implied by this survey-based VAR, we obtain 2SLS estimates of the regression coefficients in equation (7), which are shown in Figures 5 and 6. These estimates and those for $\hat{\beta}_n^{\Delta s_{t+1}}$ are also presented, along with their standard errors, in Tables 2 and 3. OLS estimates are presented in Figures A-4 and A-5 and Tables A-2 and A-3 in the Appendix.[30]

First consider the hedge group. The most striking result is that the large positive estimates of $\hat{\beta}_n^{\Delta s_{t+1}}$ at the medium and long ends of the forward rate curve in Figure 3 can be entirely explained by two components of the exchange rate change decomposition. The first

[29]Existing papers that use a similar decomposition [for example, Froot and Ramadorai (2005), Engel and West (2005; 2006; 2010), Engel, Mark, and West (2008), Mark (2009), and Engel (2014; 2016)] calculate expectations based on estimating data-generating processes that only use realized macroeconomic data and these studies ask a different set of questions relative to this paper. For more details on the VAR we use, a discussion of the benefits of using survey data, and how well the VAR implied expectations match survey data, see Stavrakeva and Tang (2018b).

[30]Note that even though the dependent variables are estimated, this does not impact the standard error calculation since the regressors are not estimated.

component is the currency risk premia term which gives the response coefficients $\hat{\beta}_n^{\sigma_t - \sigma_{t+1}^F}$.[31] The positive sign of these coefficients implies that a fall in the US forward rate relative to the foreign forward rate in response to a US monetary policy easing led to lower future expected excess returns from being long the dollar-denominated government debt and short the debt denominated in currency i. This result is consistent with the hypothesis in the literature that when conditions cause the dollar to appreciate in accordance with its exorbitant duty, this appreciation occurs due to a flight-to-safety effect. In addition to $\hat{\beta}_n^{\sigma_t - \sigma_{t+1}^F}$, we also see large positive coefficients on the long-run exchange rate expectations term, $\hat{\beta}_n^{s_{t+1,\infty}^{\Delta E}}$. This latter result implies that during the Global Recession, a decrease in the US forward rates relative to the forward rates of other countries, in response to US monetary policy shocks, was associated with a lower expected inflation path in the United States relative to other countries. The inflation channel of dollar appreciation has not been discussed in the literature on exorbitant duty until now. As we will argue in Section 5, this channel contributes to a dollar appreciation only when the policy easing shock originates in the United States.

Looking at the non-hedge currencies, Figure 6 shows that the response of relative inflation expectations to US monetary policy shocks is similar to that of the hedge currencies. Though the currency risk premia term takes the opposite sign relative to the estimates based on the hedge currencies, it tends not to be significantly different from zero, a result that is potentially due to a smaller number of observations in this sample.

The contribution of $\hat{\beta}_n^{\tilde{\imath}_t - \varphi_{t+1}^{EH}}$ to the overall coefficient, $\hat{\beta}_n^{\Delta s_{t+1}}$, is small compared to the other two components discussed above, for both the hedge and non-hedge panels, and pushes the dollar to depreciate for both groups. The negative sign is not surprising given that a main driver of relative forward rates is the expected path of future relative nominal rates, which enters negatively into this nominal rate component of exchange rates.

The OLS results are qualitatively similar, which once again implies that the unconditional relationship between the exchange rate change components and relative forward rates remains consistent with the responses to US monetary policy shocks over this period.

Decomposing the pair-specific results documented in Figure 2, we again find a sorting among the nine currencies where, for currencies where the dollar serves as a better hedge, the expected excess return from being long the dollar and short the other currency fell by more in response to US monetary policy easing (see Figure 7).

[31] We confirm that the behavior of this coefficient is driven primarily by σ_{t+1}^F, which captures changes in expectations over the future path of one-period excess returns from being long the three-month US bond and short the three-month bond of country i. The fact that the lagged expected excess return between periods t and $t+1$, σ_t, does not play an important role is not too surprising given that σ_t is not a function of period $t+1$ variables.

4.2 Further Evidence of the Flight-to-Safety Channel

To further understand how accommodative US monetary policy during the Global Recession led to a flight to safety, we show that US monetary policy easing over the Global Recession led to an increase in the risk aversion of global investors. We consider the percentage change of the aggregate relative risk aversion measure estimated in Bekaert, Engstrom, and Xu (2017) as the dependent variable in equation (2). The measure is highly correlated with the VIX and is estimated using financial variables including equity returns, corporate bond spreads, and term spreads, along with realized variances of a number of asset returns. The results, presented in Table 4, confirm that US policy easings during the Global Recession significantly increased investors' risk aversion. A fall of the US forward rate by 1 percentage point in response to US monetary policy led to an increase in risk aversion between 28 percent and 43 percent, depending on the forward rate horizon.

We conduct a second test of the effect that US monetary policy shocks have on risk aversion that is more closely related to our exchange rate change decomposition and that also emerges as a testable implication from the model that we will present in Section 5. As shown in equation (6), exchange rates depend on the entire path of expected future excess one-period returns. Thus, it is not just the current risk aversion that should matter for determining exchange rates, but the entire path of expected future risk aversion. Therefore, we also estimate how US monetary policy surprises affected the changes in expectations over the future path of the VIX, $\sum_{k=1}^{\infty} (E[VIX_{t+k}|\mathcal{I}_{t+1}] - E[VIX_{t+k}|\mathcal{I}_t])$ by, once again, estimating equation (2).[32] The results, presented in Table 5, show that US policy easings that caused a 1 percentage point fall in a US forward rate significantly lowered the expected path of the VIX—by between 1.6 and 2.7 standard deviations, depending on the forward rate horizon—during the Global Recession.

4.3 Effects on Net Foreign Asset Positions And Valuation Effects

Our evidence so far shows that US monetary policy easings during the Global Recession caused the dollar to appreciate on average against the nine currencies in our sample. The appreciation of the dollar during crisis periods is a key part of the exorbitant duty narrative. Another key component is the fall in US net foreign asset positions and the net transfer of wealth from the United States to the rest of the world.

Historically, the United States has held risky foreign-denominated assets, such as foreign equities, and safe dollar-denominated liabilities, such as US government debt. The literature

[32] Since the VIX is included in the VAR that is used in Stavrakeva and Tang (2018b) to decompose exchange rates, estimates of the path of changes in expectations of the VIX can be obtained from the same VAR in a way that is consistent with the exchange rate decomposition.

has documented that when conditions trigger the dollar to fulfill its exorbitant duty, there is a negative valuation effect on US net foreign assets due to the dollar appreciation, the fall in the price of equities, and the increase in the value of safe dollar-denominated debt. Moreover, an increase in the foreign demand for safe US debt contributes to a negative flow effect. Both the negative valuation effect and the negative flow effect lead to a decrease in US net foreign assets.

It is important to note that a rising dollar does not mechanically imply a negative valuation effect or a decrease in the net US foreign asset position. For example, the dollar appreciation might coincide with an increase in the market value of equities relative to government debt or with a positive flow effect.

We now test whether accommodative US monetary policy did lead to a decrease in the US net foreign asset position and a negative valuation effect by constructing the valuation effect following Gourinchas, Rey, and Truempler (2012). Tables 6 and 7 report the results, where the dependent variable in equation (2) is either the change in nominal US net foreign assets or the nominal valuation effect scaled by the average nominal US GDP over our Global Recession sample.[33] Indeed, we find that 1 percentage point fall of the US forward rate in response to a US monetary policy surprise led to statistically and economically significant decreases of US net foreign asset positions for all forward rate horizons, with the largest decrease being 18 percent of US GDP for $n = 36$. The corresponding number for the net wealth transfer due to the valuation effect is 17 percent of US GDP. These effects are large and very close to the total US net asset valuation losses that Gourinchas, Rey, and Govillot (2018) document occurred during the global financial crisis (14 percent) and during the European debt crisis (16 percent). This finding underscores the importance that US monetary policy can play in terms of triggering sizable wealth transfers between the United States and the rest of the world.

In summary, the US dollar appreciated, on average, in response to US monetary policy easing against a set of advanced economy currencies during the Global Recession. This rise in the dollar was accompanied by an increase in risk aversion and a transfer of wealth from the United States to the rest of the world. The channels that explain the appreciation are consistent with the "flight-to-safety" hypothesis and with the Fed's monetary policy easing engendering relatively lower inflation expectations in the United States. Moreover, in the relationship between exchange rates or future expected currency risk premia and monetary policy, we document cross-country heterogeneity that can be explained by the different hedging properties of the dollar with respect to the major currencies in our sample.

[33] We scale by average GDP to prevent our estimates from being confounded with the response of US GDP to monetary policy shocks.

The next section presents a theory of the signaling channel of monetary policy that reconciles many of these empirical facts.

5 Theoretical Interpretation: Signaling Channel of Monetary Policy

In this section, we present a partial equilibrium model that can reconcile the empirical results presented in the previous sections. We employ a signaling model of monetary policy that is similar in spirit to Tang (2015), Melosi (2017) and Andrade et al. (2018). Our model introduces a number of key deviations from the models in these papers. We model nominal exchange rates and, in the spirit of Campbell and Cochrane (1999), we allow for time-varying currency risk premia due to habit formation. The model is purposefully pared down to clearly illustrate the conditions under which monetary policy exerts a signaling effect that can qualitatively produce the empirical relationships that we document occurred during the Global Recession. This model also helps us understand why the signaling effect of monetary policy may have been especially strong during this 2008:Q4–2012:Q2 period.

5.1 Model

We consider a two-country model, in which the United States is the home country, and study the limiting case where the US economy is approximately closed.[34] This limiting case is a good approximation for large economies such as the United States. We assume that real US GDP growth—equal to the real growth in production of the US tradable good—Δy_t^{us}, and US inflation, π_t^{us}, follow exogenous processes given by:

$$\pi_t^{us} = \alpha \Delta y_t^{us}, \tag{8}$$

$$\Delta y_t^{us} = -\nu \left(i_t^{us} - \pi_t^{us}\right) + \varepsilon_t^y, \tag{9}$$

where $\alpha, \nu > 0$ and i_t^{us} is the US net nominal policy rate. Real GDP growth is assumed to be decreasing in the real interest rate, $i_t^{us} - \pi_t^{us}$, and increasing in the shock, $\varepsilon_t^y \sim N\left(0, \sigma_y^2\right)$. Assuming that $\alpha > 0$ allows us to interpret ε_t^y as a demand shock. The policy rate follows a Taylor rule given by:

$$i_t^{us} = \phi^y \Delta y_t^{us} + \phi^\pi \pi_t^{us} + \varepsilon_t^{mp},$$

[34] Considering an approximately closed US economy makes the model more tractable, but the qualitative implications of this model can be generalized beyond this special case.

where $\phi^y, \phi^\pi > 0$ and $\varepsilon_t^{mp} \sim N\left(0, \sigma_{mp}^2\right)$ is uncorrelated with the demand shock, ε_t^y. We do not impose a ZLB on the interest rate to preserve the model's simplicity. However, Andrade et al. (2018) obtain qualitatively similar results in a setting with a binding ZLB where the central bank's policy tool is an announcement about a future lift-off date from this bound.

Using the three above equations, we can solve for Δy_t^{us}, π_t^{us} and i_t^{us} in terms of the exogenous shocks:

$$\begin{aligned}
\Delta y_t^{us} &= \frac{\varepsilon_t^y - \nu \varepsilon_t^{mp}}{\eta + \nu\kappa}, \\
\pi_t^{us} &= \alpha \frac{\varepsilon_t^y - \nu \varepsilon_t^{mp}}{\eta + \nu\kappa}, \\
i_t^{us} &= \frac{\kappa \varepsilon_t^y + \eta \varepsilon_t^{mp}}{\eta + \nu\kappa},
\end{aligned}$$

where $\kappa \equiv \phi^y + \phi^\pi \alpha > 0$ and we assume that $\eta \equiv 1 - \nu\alpha > 0$, ensuring that a positive interest rate shock increases the equilibrium nominal rate. That is, we assume that the positive monetary policy shock does not cause large enough drops in contemporaneous inflation and real GDP growth to cause the equilibrium nominal interest rate to fall due to the endogenous policy reaction to these two variables.

Next, consider a representative agent located in the United States who has preferences that allow for time-varying currency risk premia due to time-varying risk aversion arising from consumption habits [see Campbell and Cochrane (1999), among other papers]. More specifically, assume that the US representative agent's per-period utility function is given by $u\left(C_t^{us}, X_t^{us}\right) = \frac{\left(C_t^{us} - X_t^{us}\right)^{1-\gamma}}{1-\gamma}$ where C_t^{us} is her consumption of the US tradable good and X_t^{us} is her respective habit reference level of consumption. We assume that the representative agent can invest in risk-free nominal bonds denominated in dollars and in the foreign currency. In the limiting case where the US economy is approximately closed, her Euler equations for bond holdings imply that:[35]

$$E\left[\beta \frac{u_c\left(C_{t+1}^{us}, X_{t+1}^{us}\right)}{u_c\left(C_t^{us}, X_t^{us}\right)} e^{-\pi_{t+1}^{us}} \left(\left(1 + i_t^{us}\right) - \frac{S_t}{S_{t+1}}\left(1 + i_t^i\right)\right) \middle| \mathcal{I}_t\right] = 0, \tag{10}$$

where S_{t+1} is the level of the nominal exchange rate defined as units of a foreign currency per dollar. The net nominal policy rate of country i is i_t^i and $\mathcal{I}_t$ is the representative agent's period t information set, which will be defined below.

The real stochastic discount factor can be expressed as $\beta \frac{u_c\left(C_{t+1}^{us}, X_{t+1}^{us}\right)}{u_c\left(C_t^{us}, X_t^{us}\right)} = \beta e^{\gamma\left(\Delta\rho_{t+1} - \Delta c_{t+1}^{us}\right)}$, where $\rho_t \equiv \ln\left(-\frac{C_t^{us}}{\gamma}\frac{u_{cc}\left(C_t^{us}, X_t^{us}\right)}{u_c\left(C_t^{us}, X_t^{us}\right)}\right) = \ln\left(\frac{C_t^{us}}{C_t^{us} - X_t^{us}}\right)$ is the log of the scaled relative risk aversion coefficient and c_{t+1}^{us} is log US consumption of the US tradable good. Given that we consider

[35] See the Appendix for details on the derivations.

the limiting case of an approximately closed US economy, then $c_t^{us} \approx y_t^{us}$. Moreover, we assume that ρ_t has the following data-generating process:

$$\begin{aligned} \Delta\rho_{t+1} &= -\lambda\bar{\rho}_t \Delta y_{t+1}^{us} \\ \text{with } \bar{\rho}_{t+1} &= \theta\bar{\rho}_t - \lambda\Delta y_{t+1}^{us}, \end{aligned} \tag{11}$$

where $0 < \theta < 1$ and $\lambda > 0$. The assumed functional form in equation (11) makes an implicit assumption regarding the functional form of X_t^{us}, which is unobserved. In the habit formation literature, it is common to specify a data-generating process for ρ_t or $\frac{1}{\rho_t}$ instead of X_t^{us} [see Campbell and Cochrane (1999) and the discussion in Brandt and Wang (2003)]. We consider parametrization such that $\bar{\rho}_t > 0$ for every t, which implies that a decrease in real GDP growth is associated with higher risk aversion for the US representative agent. While most of the model's implications derived below also hold for other data-generating processes for ρ_t, this specific functional form substantially simplifies the analysis.

Given the assumptions made, Δy_{t+k}^{us} and $\bar{\rho}_{t+k}$ are normally distributed, conditional on observing all shocks realized in period t or earlier, for any $k \geq 1$. We conjecture that Δs_{t+1} is also normally distributed, which we later confirm. As a result, equation (10) allows us to express the expected excess return from being long the dollar-denominated bond and short the foreign-currency-denominated bond as:

$$\sigma_t = i_t^{us} - i_t^i + E[\Delta s_{t+1}|\mathcal{I}_t] \tag{12}$$

$$= \frac{Var\left(\Delta s_{t+1}|\mathcal{I}_t\right)}{2} - Cov\left(\Delta s_{t+1}, -\gamma\Delta c_{t+1}^{us} + \gamma\Delta\rho_{t+1} - \pi_{t+1}^{us}|\mathcal{I}_t\right)$$

$$= \frac{Var\left(\Delta s_{t+1}|\mathcal{I}_t\right)}{2} + \left(\gamma + \alpha + \gamma\lambda\bar{\rho}_t\right) Cov\left(\Delta s_{t+1}, \Delta y_{t+1}^{us}|\mathcal{I}_t\right), \tag{13}$$

where the last equality uses the fact that $\Delta c_{t+k}^{us} \approx \Delta y_{t+k}^{us}$ and equations (8) and (11).

We assume for now that $Var\left(\Delta s_{t+1}|\mathcal{I}_t\right)$ and $Cov\left(\Delta s_{t+1}, \Delta y_{t+1}^{us}|\mathcal{I}_t\right)$ are constant and later provide conditions under which this is the case in equilibrium. Then, the only source of time-variation in the expected excess return, σ_t, is the time-varying risk aversion. From equation (12), we see that higher $\bar{\rho}_t$ is associated with a lower expected excess return from holding the dollar between t and $t+1$ against currencies for which the dollar is a good hedge, i.e., $Cov\left(\Delta s_{t+1}, \Delta y_{t+1}^{us}|\mathcal{I}_t\right) < 0$. The opposite is true for non-hedge currencies for which $Cov\left(\Delta s_{t+1}, \Delta y_{t+1}^{us}|\mathcal{I}_t\right) > 0$.

The partial equilibrium nature of the model allows us to avoid taking a stand on whether markets are complete or incomplete and on what drives the hedging properties of the US dollar (i.e. what country-specific properties make the dollar a hedge for some currencies and not for others). As a result, it is easier to identify the properties of the exchange rate and its components that are crucial for the theory to match our empirical findings.

We can iterate equation (12) forward to express the log change in the nominal exchange rate, Δs_{t+1}, in terms of the same components that we used in the empirical sections of the paper:

$$
\begin{aligned}
\Delta s_{t+1} =& i_t^i - i_t^{us} - \underbrace{\sum_{k=0}^{\infty} \left(E[i_{t+k+1}^i - i_{t+k+1}^{us} | \mathcal{I}_{t+1}] - E[i_{t+k+1}^i - i_{t+k+1}^{us} | \mathcal{I}_t] \right)}_{\varphi_{t+1}^{EH}} \\
&+ \sigma_t - \underbrace{\sum_{k=0}^{\infty} \left(E[\sigma_{t+k+1} | \mathcal{I}_{t+1}] - E[\sigma_{t+k+1} | \mathcal{I}_t] \right)}_{\sigma_{t+1}^{F}} \\
&+ \underbrace{\sum_{k=0}^{\infty} \left(E[\pi_{t+k+1}^i - \pi_{t+k+1}^{us} | \mathcal{I}_{t+1}] - E[\pi_{t+k+1}^i - \pi_{t+k+1}^{us} | \mathcal{I}_t] \right)}_{s_{t+1,\infty}^{\Delta E}}. \qquad (14)
\end{aligned}
$$

In doing so, we once again use the assumption that purchasing power parity holds in the long run. Therefore, as before, changes in expectations of the long-run nominal exchange rate level, $s_{t+1,\infty}^{\Delta E}$, equals the changes in expectations over the relative path of future inflation.

From the assumptions that the US demand and monetary policy shocks are i.i.d. normal with constant variances, the US variables and risk premium terms in equation (14) are all conditionally normally distributed with constant variances and constant covariances with Δy_{t+1}^{us}. Thus, the additional assumptions that the nominal interest rate and inflation in the foreign country i, i_{t+k}^i and π_{t+k}^i, are conditionally normally distributed with constant second moments is enough to guarantee that the overall log change in the nominal exchange rate, Δs_{t+1}, will also be conditionally normally distributed with constant second moments.[36]

5.2 The Effect of Forward Guidance

To keep the analysis as simple as possible, we assume that at time $t+1$, the central bank knows future the state of the economy and the monetary policy surprise in period $t+h$ for $h \geq 2$. That is, the central bank can perfectly observe ε_{t+h}^{mp} and ε_{t+h}^{y} at time $t+1$. To obtain our results, it is sufficient for the agents who trade short-term bonds denominated in different currencies to *believe* that the Fed has some additional information about ε_{t+h}^{mp} and ε_{t+h}^{y}. Nakamura and Steinsson (2018) provide a detailed discussion regarding whether the private sector interprets FOMC announcements as a signal about future expectations of economic activity.

[36] See the Appendix for details.

We consider a forward guidance announcement to be the central bank's truthful expectation of i^{us}_{t+h}, based on the Taylor rule. Given that there is no persistence in the variables affecting the policy rate, this forward guidance is equivalent to the central bank announcing the actual policy rate $h-1$ periods from now. Denote the announcement in period $t+1$ as a_{t+1}. Given the assumptions made, $a_{t+1} = i^{us}_{t+h}$. We assume that the agent's time $t+1$ information set contains current and past values of announcements and shocks, i.e., $\mathcal{I}_{t+1} = \{a^{t+1}, \varepsilon^{y,t+1}, \varepsilon^{mp,t+1}\}$. Since shocks are i.i.d., just before the announcement, agents expect i^{us}_{t+h} to be zero, so the entire forward guidance announcement is a surprise.

Assume that the change in the one-period relative forward rate (defined as the non-US forward rate minus the US forward rate) prevailing between periods $t+h$ and $t+h+1$ caused by the announcement a_{t+1} is equal to $-i^{us}_{t+h}$.[37] Then, our estimates of $\hat{\beta}_n^{\sigma_t - \sigma^F_{t+1}}$ and $\hat{\beta}_n^{s^{\Delta E}_{t+1,\infty}}$ in Section 4.1.1 correspond to the derivatives $-\frac{\partial\left(\sigma_t - \sigma^F_{t+1}\right)}{\partial a_{t+1}}$ and $-\frac{\partial s^{\Delta E}_{t+1,\infty}}{\partial a_{t+1}}$ in the model.

First, we derive the effect of the announcement on expected future real GDP growth which, as we will show, is the main driver of the changes in expectations of future currency risk premia and of the relative inflation paths in the two countries.

The agent's expectation of Δy_{t+h} involves a signal extraction problem. Since the future policy rate is a function of both future monetary policy shocks and demand shocks, the forward guidance announcement does not completely reveal the realizations of each shock. However, the agent uses this announcement to extract information about ε^y_{t+h} and ε^{mp}_{t+h}, which then informs her expectation about Δy_{t+h}. Using the posterior expectations of the two shocks in $t+h$, which are presented in the Appendix, one can show that:

$$E\left[\Delta y_{t+h} | \mathcal{I}_{t+1}\right] = K a_{t+1}, \quad \text{where } K \equiv \frac{\kappa \frac{\sigma_y^2}{\sigma_{mp}^2} - \nu\eta}{\kappa^2 \frac{\sigma_y^2}{\sigma_{mp}^2} + \eta^2}. \tag{15}$$

When $\frac{\sigma_y^2}{\sigma_{mp}^2} = 0$, the agent believes that the forward guidance announcement is driven only by a future exogenous monetary policy shock, i.e., $a_{t+1} = i^{us}_{t+h} = \frac{\eta}{\eta+\nu\kappa}\varepsilon^{mp}_{t+h}$. In this case, the effect of the announcement on GDP growth expectations is given by $-\frac{\nu}{\eta} < 0$, which only captures the direct effect of the future interest rate shock on expected real GDP growth, where a negative interest rate surprise improves GDP growth expectations.

The signaling channel appears when $\frac{\sigma_y^2}{\sigma_{mp}^2} > 0$. Given our parameterization, K is increasing in $\frac{\sigma_y^2}{\sigma_{mp}^2}$. For a sufficiently high $\frac{\sigma_y^2}{\sigma_{mp}^2}$ (i.e., a sufficiently strong signaling channel), K can become positive, meaning that an announcement of a lower future policy rate can *lower* expectations of future real GDP growth.

[37]For this assumption to hold, the sum of the movements of the other country's forward rate and the relative term premia of both forward rates in response to the announcement should be zero. These assumptions are made primarily for tractability in the model and can be relaxed.

More generally, if $\frac{\sigma_y^2}{\sigma_{mp}^2} < \frac{\nu\eta}{\kappa}$, then the direct channel dominates ($K < 0$), and if the opposite is true, the signaling channel dominates ($K > 0$). This result is intuitive, as high prior uncertainty about the demand shock implies that the agent will place more weight on a signal containing information about this demand shock, a_{t+1}, when updating her beliefs about future real GDP growth. In this paper, when we say that the signaling channel is strong, we mean that the signaling channel is strong enough to dominate the direct effect of interest rate movements on real GDP growth, implying that announcing a lower future policy rate causes future real GDP growth expectations to fall. To summarize, in our terminology, the fact that forward guidance has a strong signaling effect corresponds to the case in which $K > 0$.

Indeed, we confirm that the signaling channel of US monetary policy dominated the direct expansionary effects of lower rates during the Global Recession by examining the effect of US policy surprises on real GDP growth expectations in this period. In particular, we estimate equation (2), with the dependent variable being revisions in four-quarter-ahead forecasts of real GDP growth obtained from *Blue Chip Economic Indicators*. The forecast revision is the change between the lagged four-quarter-ahead forecast and the current three-quarter-ahead forecast, thus keeping fixed the forecast quarter. Table 8 presents the results. We find that in response to US monetary policy surprises that lowered the US forward rate by 1 percentage point, GDP growth expectations were revised downwards by a statistically significant and economically meaningful amount that is between 0.71 and 1.03 percentage points, depending on the forward rate horizon. The estimates of these effects imply that $K > 0$ over the Global Recession.

Next, we derive the above-mentioned derivatives of our exchange rate components with respect to the announcement and show that these are tightly linked to $\frac{\partial E[\Delta y_{t+h}|\mathcal{I}_{t+1}]}{\partial a_{t+1}}$. We start with the derivative for $\sigma_t - \sigma^F_{t+1}$. First, note that since σ_t contains information up to only t, $-\frac{\partial\left(\sigma_t - \sigma^F_{t+1}\right)}{\partial a_{t+1}} = \frac{\partial \sigma^F_{t+1}}{\partial a_{t+1}}$. In the Appendix, we show that:

$$
\begin{aligned}
-\frac{\partial\left(\sigma_t - \sigma^F_{t+1}\right)}{\partial a_{t+1}} &= \frac{\partial \sigma^F_{t+1}}{\partial a_{t+1}} = \gamma\lambda\sigma_{s,y}\sum_{k=0}^{\infty}\frac{\partial}{\partial a_{t+1}}\left(E\left[\bar{\rho}_{t+k+1}|\mathcal{I}_{t+1}\right] - E\left[\bar{\rho}_{t+k+1}|\mathcal{I}_{t}\right]\right) \\
&= -\frac{\gamma\lambda^2\sigma_{s,y}}{1-\theta}\underbrace{\frac{\partial E\left[\Delta y_{t+h}|\mathcal{I}_{t+1}\right]}{\partial a_{t+1}}}_{K},
\end{aligned}
$$

where $\sigma_{s,y}$ denotes the constant value of $Cov\left(\Delta s_{t+1}, \Delta y^{us}_{t+1}|\mathcal{I}_t\right)$. If the signaling channel is strong, meaning $K > 0$, as the evidence in Table 8 suggests was true during the Global Recession, a negative forward guidance shock lowers expectations of future real GDP growth. This, in turn, increases expectations of future risk aversion and lowers the expected excess return from being long the dollar bond and short the bond of country i if the dollar is a hedge for currency i, i.e. if $Cov\left(\Delta s_{t+1}, \Delta y^{us}_{t+1}|\mathcal{I}_t\right) < 0$. The opposite is true if the dollar is not a hedge for currency i, i.e. if $Cov\left(\Delta s_{t+1}, \Delta y^{us}_{t+1}|\mathcal{I}_t\right) > 0$. Thus, our empirical findings showing that during the Global Recession, $\hat{\beta}_n^{\sigma_t - \sigma^F_{t+1}} > 0$ for our group of hedge currencies and $\hat{\beta}_n^{\sigma_t - \sigma^F_{t+1}} < 0$ for our group of non-hedge currencies is consistent with the signaling channel being dominant over this crisis period. Moreover, while $\sigma_{s,y}$ can vary across currencies, the rest of the parameters in the expression above are not currency-specific. The more negative $\sigma_{s,y}$ is, the better of a hedge the dollar is, so $-\frac{\partial\left(\sigma_t - \sigma^F_{t+1}\right)}{\partial a_{t+1}}$ will be more positive. This model implication is consistent with the cross-currency heterogeneity in our estimated responses of each currency's risk premium components to US monetary policy surprises, responses that were presented in subsection 4.1.

Finally, we derive the effect of the forward guidance announcement on the long-run nominal exchange rate component, the second exchange rate change component that contributed to the structural break. For simplicity, we assume that US monetary policy shocks do not affect inflation expectations in other countries. Then,

$$
\begin{aligned}
-\frac{\partial s^{\Delta E}_{t+1,\infty}}{\partial a_{t+1}} &= \frac{\partial}{\partial a_{t+1}} \sum_{k=1}^{\infty} \left(E\left[\pi^{us}_{t+k}|\mathcal{I}_{t+1}\right] - E\left[\pi^{us}_{t+k}|\mathcal{I}_t\right]\right) \\
&= \frac{\partial}{\partial a_{t+1}} \left(E\left[\pi^{us}_{t+h}|\mathcal{I}_{t+1}\right] - E\left[\pi^{us}_{t+h}|\mathcal{I}_t\right]\right) = \alpha \underbrace{\frac{\partial E\left[\Delta y_{t+h}|\mathcal{I}_{t+1}\right]}{\partial a_{t+1}}}_{K}.
\end{aligned} \tag{16}
$$

Once again, understanding the effect that the forward guidance surprise has on real GDP growth expectations is sufficient to understand the model's second key derivative. Since we assumed that economic fluctuations are driven primarily by demand shocks, lower real GDP growth is associated with lower inflation. The result is that a negative forward guidance announcement leads to lower expectations of future real GDP growth and inflation when the signaling channel is strong. Thus, conditional on a strong signaling channel during the Global Recession, our model implies that $-\frac{\partial s^{\Delta E}_{t+1,\infty}}{\partial a_{t+1}} > 0$, consistent with the increase in expected future relative inflation in response to US policy easings that we find in the data.

An important observation coming out of the theory is that, during times of crisis, the inflation channel will only push the dollar to appreciate if the initial negative shock to growth expectations originates in the United States—i.e., if the shock leads to lower US growth expectations and, thus, lower expected US inflation relative to countries. By extending the model, one can easily see that if the original shock lowers growth expectations and, thus, inflation expectations, in the other economy by more than in the United States, then the inflation channel will instead push the dollar to depreciate. In contrast, the currency risk premium channel of dollar appreciation will be present, regardless of whether the initial shock stems from the United States or elsewhere in the world, as long as the shock increases risk aversion and as long as the dollar is a good hedge. In summary, while the dollar's exorbitant duty behavior can be triggered by a shock that lowers either expected future US growth or expected future growth in another country, the US shock will cause the dollar to appreciate through both the inflation and flight-to-safety channels while the non-US shock will cause the dollar to appreciate only if the flight-to-safety channel dominates the inflation channel.

To summarize, in this section we showed that US monetary policy sent a sufficiently strong signal about economic conditions during the Global Recession. When joined with preferences featuring habit formation, this strong signaling effect can explain all the responses of exchange rate changes and its components to US monetary policy shocks that we observe in the data. Moreover, the model's predictions are also consistent with other empirical facts that we document, such as accommodative forward guidance policy during the Global Recession leading to downward revisions of US GDP growth forecasts and higher current and expected future risk aversion.

Our results beg the question of why the signaling channel was so much stronger than the direct channel of monetary policy during the Global Recession. We use the model to offer an answer to this question.

First, in the model, the signaling channel is strong when $\frac{\sigma_y^2}{\sigma_{mp}^2}$ is sufficiently high. Evidence that this ratio was particularly high during the Global Recession can be found by examining the average values of macroeconomic and monetary policy uncertainty measures before the Global Recession, during it and after, all of which are presented in Table 9. The two macroeconomic uncertainty measures we consider are the 12-month-ahead macroeconomic uncertainty estimated by Jurado, Ludvigson, and Ng (2015) and the dispersion in four-quarter-ahead US real GDP forecasts obtained from *Blue Chip Economic Indicators*. Not surprisingly, both measures of macroeconomic uncertainty were much higher during the Global Recession subsample versus the other two subsamples. The monetary policy uncertainty measure that we examine is the monetary policy subcomponent of the Baker, Bloom, and Davis (2016) policy uncertainty index.[38] This uncertainty measure actually declines slightly during the Global Recession, most likely due to the ZLB, and declines further still in the time period after 2012:Q2. These results are consistent with $\frac{\sigma_y^2}{\sigma_{mp}^2}$ being particularly high during the Global Recession relative to the periods prior to and after the recession.

Second, during the Global Recession, as in the model, the Federal Reserve used "calendar-

[38]Note that this measure of monetary policy uncertainty could capture uncertainty about both the exogenous monetary policy shock as well as the endogenous responses of monetary policy to economic conditions. However, the divergence of macroeconomic and monetary policy uncertainty during the Global Recession subsample suggests that uncertainty about the monetary policy shock likely declined in this period.

based" forward guidance that promised low rates at least until some future date—a policy that can easily be interpreted as a low assessment of future US growth prospects by the Fed.[39] Monetary policy being conducted in a way that leaves potential for policy actions to be interpreted as signals about the state of the economy is another necessary condition for these policy actions to have a strong signaling effect.

The empirical and theoretical evidence presented in this paper lead us to conclude that periods featuring high fundamental uncertainty relative to monetary policy uncertainty, during which US monetary policy is conducted in such a way that it can be interpreted as signaling information about the economy, are the times when accommodative US monetary policy imposes exorbitant duty effects on the dollar and induces wealth transfers from the United States to the rest of the world.

6 Conclusion

This paper examines the behavior of the dollar over the Global Recession. We find that surprise US monetary policy easings during this period caused the dollar to appreciate, on average, and lowered the value of US net foreign asset positions, thus, pushing the United States toward fulfilling its exorbitant duty of transferring wealth to the rest of the world during crisis times.

Consistent with the narrative in the literature on the "exorbitant duty," we show that a flight-to-safety effect, due to higher risk aversion, is one of the main drivers behind the dollar appreciation in response to US policy easings. We also document a second channel whereby expansionary US monetary policy also increased the value of the dollar over the Global Recession by lowering the expected future path of US inflation relative to other countries.

These two effects of accommodative US monetary policy during the Global Recession are surprising in the context of standard macroeconomic models which predict that US policy easing relative to another country should depreciate the dollar. We present a partial equilibrium model which illustrates that when accommodative monetary policy is interpreted as a signal of worsening future macroeconomic conditions and the signaling channel of monetary policy is strong, as we show is the case over the crisis, then one can explain the surprising empirical results that we document.

Our findings suggest that the type of models used to understand the behavior of the dollar and its link to monetary policy need to be reevaluated, as existing exchange rate models do not consider the signaling effect of monetary policy. As a result, the policy conclusions made

[39] In contrast, the "threshold-based" forward guidance used after 2012:Q2 left less room for interpretation as it mainly communicated the Fed's policy reaction function.

during crises, such as the Global Recession, can be severely flawed.

References

Andrade, Philippe, Gaetano Gaballo, Eric Mengus, and Benoît Mojon. 2018. "Forward Guidance and Heterogeneous Beliefs." SSRN Scholarly Paper ID 3112274. Rochester, NY: Social Science Research Network. Available at https://papers.ssrn.com/abstract=3112274.

Baele, Lieven, Geert Bekaert, and Koen Inghelbrecht. 2010. "The Determinants of Stock and Bond Return Comovements." *The Review of Financial Studies* 23(6): 2374–2428.

Baele, Lieven, Geert Bekaert, Koen Inghelbrecht, and Min Wei. 2018. "Flights to Safety." SSRN Scholarly Paper ID 3204192. Rochester, NY: Social Science Research Network. Available at https://papers.ssrn.com/abstract=3204192.

Bai, Jushan, and Pierre Perron. 1998. "Estimating and Testing Linear Models with Multiple Structural Changes." *Econometrica* 66(1): 47–78.

Baker, Scott R., Nicholas Bloom, and Steven J. Davis. 2016. "Measuring Economic Policy Uncertainty." *Quarterly Journal of Economics* 131(4): 1593–1636.

Beber, Alessandro, Michael W. Brandt, and Kenneth A. Kavajecz. 2009. "Flight-to-Quality or Flight-to-Liquidity? Evidence from the Euro-Area Bond Market." *The Review of Financial Studies* 22(3): 925–957.

Bekaert, Geert, Eric Engstrom, and Nancy Xu. 2017. "The Time Variation in Risk Appetite and Uncertainty." SSRN Scholarly Paper ID 3069078. Rochester, NY: Social Science Research Network. Available at https://papers.ssrn.com/abstract=3069078.

Berkelmans, Leon. 2011. "Imperfect Information, Multiple Shocks, and Policy's Signaling Role." *Journal of Monetary Economics* 58(4): 373–386.

Bernanke, Ben S. 2017. "Federal Reserve Policy in an International Context." *IMF Economic Review* 65(1): 5–36.

Bernanke, Ben S., and Kenneth N. Kuttner. 2005. "What Explains the Stock Market's Reaction to Federal Reserve Policy?" *Journal of Finance* 60(3): 1221–1257.

Brandt, Michael W., and Kevin Q. Wang. 2003. "Time-Varying Risk Aversion and Unexpected Inflation." *Journal of Monetary Economics* 50(7): 1457–1498.

Brunnermeier, Markus K., and Lasse Heje Pedersen. 2009. "Market Liquidity and Funding Liquidity." *The Review of Financial Studies* 22(6): 2201–2238.

Caballero, Ricardo, Emmanuel Farhi, and Pierre-Olivier Gourinchas. 2008. "An Equilibrium Model of Global Imbalances and Low Interest Rates." *American Economic Review* 98(1): 358–393.

Caballero, Ricardo J., and Arvind Krishnamurthy. 2008. "Collective Risk Management in a Flight to Quality Episode." *Journal of Finance* 63(5): 2195–2230.

Campbell, Jeffrey R., Charles L. Evans, Jonas D. M. Fisher, and Alejandro Justiniano. 2012. "Macroeconomic Effects of FOMC Forward Guidance." *Brookings Papers on Economic Activity* 44(1): 1–80. Available at https://www.brookings.edu/wp-content/uploads/2012/03/2012a_Evans.pdf.

Campbell, John, and John Cochrane. 1999. "By Force of Habit: A Consumption-Based Explanation of Aggregate Stock Market Behavior." *Journal of Political Economy* 107(2): 205–251.

Casas, Camila, Federico Diez, Gita Gopinath, and Pierre-Olivier Gourinchas. 2017. "Dominant Currency Paradigm." Mimeo, Harvard. Available at https://scholar.harvard.edu/files/gopinath/files/paper_080217.pdf.

Chahrour, Ryan, and Rosen Valchev. 2017. "International Medium of Exchange: Privilege and Duty." Boston College Working Papers in Economics 934. Boston College Department of Economics. Available at http://fmwww.bc.edu/EC-P/wp934.pdf.

Crump, Richard K., Stefano Eusepi, and Emanuel Moench. 2016. "The Term Structure of Expectations and Bond Yields." Staff Report 775. New York: Federal Reserve Bank of New York. Available at https://www.newyorkfed.org/medialibrary/media/research/staff_reports/sr775.pdf.

Cukierman, Alex, and Allan Meltzer. 1986. "A Theory of Ambiguity, Credibility, and Inflation under Discretion and Asymmetric Information." *Econometrica* 54(5): 1099–1128.

Eichenbaum, Martin, and Charles L. Evans. 1995. "Some Empirical Evidence on the Effects of Shocks to Monetary Policy on Exchange Rates." *Quarterly Journal of Economics* 110(4): 975–1009.

Ellingsen, Tore, and Ulf Söderström. 2001. "Monetary Policy and Market Interest Rates." *American Economic Review* 91(5): 1594–1607.

Engel, Charles. 2014. "Exchange Rates and Interest Parity." In *Handbook of International Economics*, Elhanan Helpman, Gita Gopinath, and Kenneth Rogoff, eds., vol. 4, 453–522. Amsterdam: Elsevier.

Engel, Charles. 2016. "Exchange Rates, Interest Rates, and the Risk Premium." *American Economic Review* 106(2): 436–474.

Engel, Charles, Nelson C. Mark, and Kenneth D. West. 2008. "Exchange Rate Models Are Not as Bad as You Think." In *NBER Macroeconomics Annual 2007*, Daron Acemoglu, Kenneth Rogoff, and Michael Woodford, eds., vol. 22, 381–441. Chicago: University of Chicago Press.

Engel, Charles, and Kenneth D. West. 2005. "Exchange Rates and Fundamentals." *Journal of Political Economy* 113(3): 485–517.

Engel, Charles, and Kenneth D. West. 2006. "Taylor Rules and the Deutschmark: Dollar Real Exchange Rate." *Journal of Money, Credit and Banking* 38(5): 1175–1194.

Engel, Charles, and Kenneth D. West. 2010. "Global Interest Rates, Currency Returns, and the Real Value of the Dollar." *American Economic Review* 100(2): 562–67.

Farhi, Emmanuel, and Matteo Maggiori. 2018. "A Model of the International Monetary System." *Quarterly Journal of Economics* 133(1): 295–355.

Froot, Kenneth A., and Tarun Ramadorai. 2005. "Currency Returns, Intrinsic Value, and Institutional-Investor Flows." *Journal of Finance* 60(3): 1535–1566.

Gertler, Mark, and Peter Karadi. 2015. "Monetary Policy Surprises, Credit Costs, and Economic Activity." *American Economic Journal: Macroeconomics* 7(1): 44–76.

Gilchrist, Simon, Egon Zakrajšek, and Vivian Z. Yue. 2016. "The Response of Sovereign Bond Yields to U.S. Monetary Policy." *Journal Economía Chilena (The Chilean Economy)* 19(2): 102–106.

Goldberg, Linda, and Cédric Tille. 2008. "Vehicle Currency Use in International Trade." *Journal of International Economics* 76(2): 177—192.

Gopinath, Gita. 2016. "The International Price System." *Jackson Hole Symposium Proceedings*. Available at https://www.kansascityfed.org/~/media/files/publicat/sympos/2015/2015gopinath.pdf.

Gopinath, Gita, and Jeremy C Stein. 2018. "Banking, Trade, and the making of a Dominant Currency." Working Paper 24485. Cambridge, MA: National Bureau of Economic Research.

Gourinchas, Pierre-Olivier, and Hélène Rey. 2007a. "From World Banker to World Venture Capitalist: US External Adjustment and the Exorbitant Privilege." In *G-7 Current Account Imbalances: Sustainability and Adjustment*, Richard Clarida, ed., 11–55. Chicago: University of Chicago Press.

Gourinchas, Pierre-Olivier, and Hélène Rey. 2007b. "International Financial Adjustment." *Journal of Political Economy* 115(4): 665–703.

Gourinchas, Pierre-Olivier, and Hélène Rey. 2014. "External Adjustment, Global Imbalances, Valuation Effects." In *Handbook of International Economics*, Kenneth Rogoff Elhanan Helpman and Gita Gopinath, eds., vol. 4, 585–645. Amsterdam: Elsevier.

Gourinchas, Pierre-Olivier, Hélène Rey, and Nicolas Govillot. 2018. "Exorbitant Privilege and Exorbitant Duty." Mimeo.

Gourinchas, Pierre-Olivier, Hélène Rey, and Maxime Sauzet. Forthcoming. "The International Monetary and Financial System." *Annual Review of Economics*.

Gourinchas, Pierre-Olivier, Hélène Rey, and Kai Truempler. 2012. "The Financial Crisis and the Geography of Wealth Transfers." *Journal of International Economics* 88(2): 266–283.

Greenwood, Robin, Samuel G. Hanson, and Dimitri Vayanos. 2016. "Forward Guidance in the Yield Curve: Short Rates versus Bond Supply." In *Monetary Policy through Asset Markets: Lessons from Unconventional Measures and Implications for an Integrated*

World, Elías Albagli, Diego Saravia, and Michael Woodford, eds., vol. 24 of *Central Banking, Analysis, and Economic Policies Book Series*, chap. 2, 011–062. Santiago: Central Bank of Chile.

Gürkaynak, Refet S., Brian Sack, and Eric T. Swanson. 2005. "Do Actions Speak Louder Than Words? The Response of Asset Prices to Monetary Policy Actions and Statements." *International Journal of Central Banking* 1(1): 55–93.

Gürkaynak, Refet S., Brian Sack, and Jonathan H. Wright. 2007. "The U.S. Treasury Yield Curve: 1961 to the Present." *Journal of Monetary Economics* 54(8): 2291–2304.

Hassan, Tarek. 2013. "Country Size, Currency Unions, and International Asset Returns." *Journal of Finance* 68(6): 2269–2308.

Hassan, Tarek A., Thomas M. Mertens, and Tony Zhang. 2016. "Currency Manipulation." SSRN Scholarly Paper ID 2819214. Rochester, NY: Social Science Research Network. Available at https://papers.ssrn.com/abstract=2819214.

He, Zhiguo, Arvind Krishnamurthy, and Konstantin Milbradt. Forthcoming. "A Model of Safe Asset Determination." *American Economic Review*.

Ivashina, Victoria, David S. Scharfstein, and Jeremy C. Stein. 2015. "Dollar Funding and the Lending Behavior of Global Banks." *Quarterly Journal of Economics* 130(3): 1241–1281.

Jurado, Kyle, Sydney C. Ludvigson, and Serena Ng. 2015. "Measuring Uncertainty." *American Economic Review* 105(3): 1177–1216.

Kim, Don H., and Athanasios Orphanides. 2012. "Term Structure Estimation with Survey Data on Interest Rate Forecasts." *Journal of Financial and Quantitative Analysis* 47(1): 241–272.

Kim, Don H., and Jonathan H. Wright. 2005. "An Arbitrage-Free Three-Factor Term Structure Model and the Recent Behavior of Long-Term Yields and Distant-Horizon Forward Rates." Finance and Economics Discussion Series 2005-33. Washington, DC: Board of Governors of the Federal Reserve System. Available at https://www.federalreserve.gov/pubs/feds/2005/200533/200533pap.pdf.

Kuttner, Kenneth. 2001. "Monetary Policy Surprises and Interest Rates: Evidence From the Fed Funds Futures Market." *Journal of Monetary Economics* 47(3): 523–44.

Maggiori, Matteo. 2013. "The U.S. Dollar Safety Premium." Mimeo.

Maggiori, Matteo. 2017. "Financial Intermediation, International Risk Sharing, and Reserve Currencies." *American Economic Review* 107(10): 3038–3071.

Mark, Nelson C. 2009. "Changing Monetary Policy Rules, Learning, and Real Exchange Rate Dynamics." *Journal of Money, Credit and Banking* 41(6): 1047–1070.

Melosi, Leonardo. 2017. "Signaling Effects of Monetary Policy." *Review of Economic Studies* 84(2): 853–884.

Nakamura, Emi, and Jón Steinsson. 2018. "High Frequency Identification of Monetary Non-Neutrality: The Information Effect." *Quarterly Journal of Economics* 133(3): 1283–1330.

Perron, Pierre, and Yohei Yamamoto. 2015. "Using OLS to Estimate and Test for Structural Changes in Models with Endogenous Regressors." *Journal of Applied Econometrics* 30(1): 119–144.

Piazzesi, Monika, Juliana Salomao, and Martin Schneider. 2015. "Trend and Cycle in Bond Premia." Mimeo. Available at http://www.stanford.edu/~piazzesi/trendcycle.pdf.

Piazzesi, Monika, and Eric T. Swanson. 2008. "Futures prices as risk-adjusted forecasts of monetary policy." *Journal of Monetary Economics* 55(4): 677–691.

Rogers, John H., Chiara Scotti, and Jonathan H. Wright. 2014. "Evaluating Asset-Market Effects of Unconventional Monetary Policy: A Multi-Country Review." *Economic Policy* 29(80): 749–799.

Rogers, John H., Chiara Scotti, and Jonathan H. Wright. Forthcoming. "Unconventional Monetary Policy and International Risk Premia." *Journal of Money, Credit and Banking.*

Shin, Hyun Song. 2012. "Global Banking Glut and Loan Risk Premium." *IMF Economic Review* 60(2): 155–192.

Stavrakeva, Vania, and Jenny Tang. 2018a. "The Effect of Unconventional Monetary Policy on Foreign Exchange Markets." Mimeo.

Stavrakeva, Vania, and Jenny Tang. 2018b. "Survey-based Exchange Rate Decomposition: New Methodology and New Facts." Mimeo. Available at https://goo.gl/h53za5.

Stock, James H., and Mark W. Watson. 2018. "Identification and Estimation of Dynamic Causal Effects in Macroeconomics Using External Instruments." Working Paper 24216. Cambridge, MA: National Bureau of Economic Research.

Swanson, Eric T. 2017. "Measuring the Effects of Federal Reserve Forward Guidance and Asset Purchases on Financial Markets." Working Paper 23311. Cambridge, MA: National Bureau of Economic Research.

Tang, Jenny. 2015. "Uncertainty and the Signaling Channel of Monetary Policy." Working Paper 15-8. Boston: Federal Reserve Bank of Boston. Available at https://www.bostonfed.org/-/media/Documents/Workingpapers/PDF/wp1508.pdf.

Vayanos, Dimitri. 2004. "Flight to Quality, Flight to Liquidity, and the Pricing of Risk." SSRN Scholarly Paper ID 509858. Rochester, NY: Social Science Research Network. Available at https://papers.ssrn.com/abstract=509858.

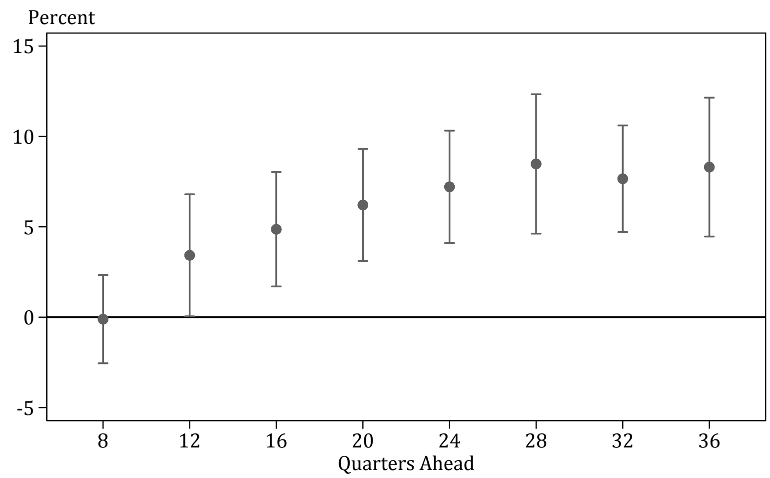
Wright, Jonathan H. 2011. "Term Premia and Inflation Uncertainty: Empirical Evidence from an International Panel Dataset." *American Economic Review* 101(4): 1514–34.

Wright, Jonathan H. 2012. "What does Monetary Policy do to Long-term Interest Rates at the Zero Lower Bound?" *The Economic Journal* 122(564): F447–F466.

Wu, Tao. 2014. "Unconventional Monetary Policy and Long-Term Interest Rates." IMF Working Papers 14/189. Washington, DC: International Monetary Fund. Available at https://www.imf.org/external/pubs/ft/wp/2014/wp14189.pdf.

Figures and Tables

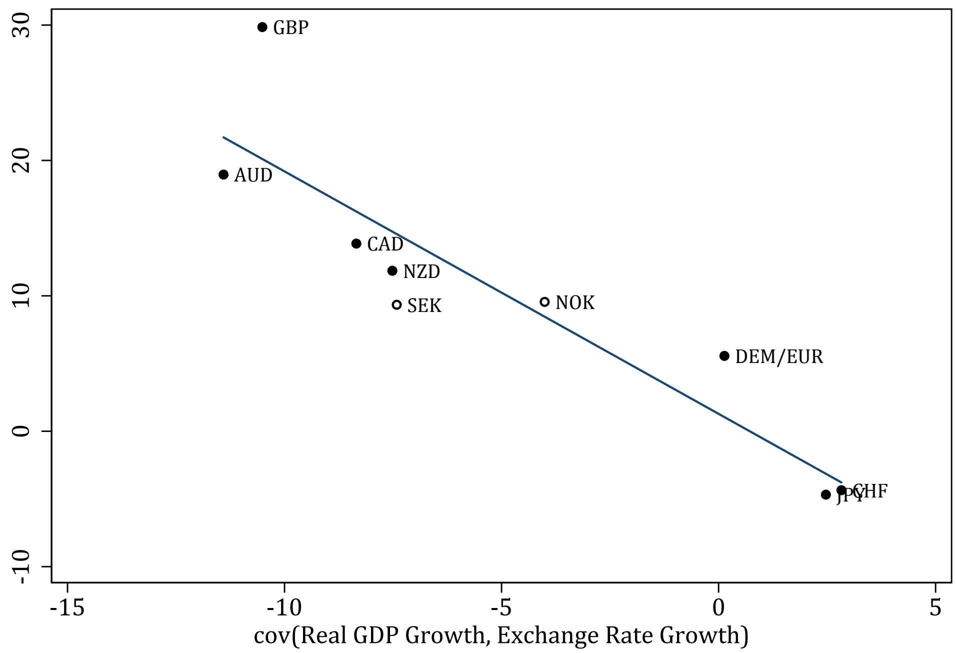
Figure 1: Panel Response of Exchange Rate Changes to US Monetary Policy Surprises for All Currencies (2SLS)



Source: Authors' calculations.

Note: 90 percent confidence intervals based on Driscoll-Kraay standard errors. 2SLS regression of exchange rate change on relative forward rate changes instrumented using yield changes in a one-hour window around FOMC announcements of FF4, ED4, and two- and 10-year Treasury note futures expiring in the current quarter. This sample includes Australia, Canada, euro area, Japan, Norway, New Zealand, Sweden, Switzerland, and the United Kingdom against the US dollar.

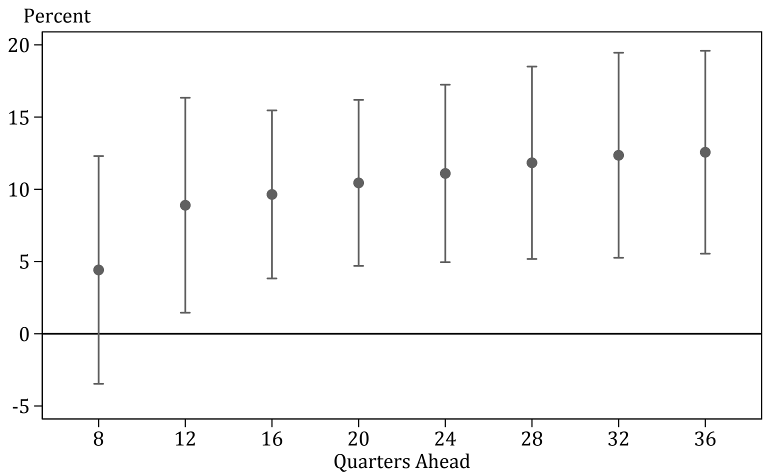
Figure 2: Pair-Specific Response of Exchange Rate Changes to US Monetary Policy Surprises versus Hedging Properties of the Dollar (2SLS)



Source: Authors' calculations.

Note: Filled circles represent significance at 10 percent based on Drisoll-Kraay standard errors. 2SLS regression of exchange rate change on relative forward rate changes instrumented using yield changes in a one-hour window around FOMC announcements of FF4, ED4, and two- and 10-year Treasury note futures expiring in the current quarter. The measure of the dollar's value as a hedge for each currency pair is the covariance between the respective exchange rate change and the US real GDP growth. This covariance is calculated over a longer sample starting in 1991:Q1.

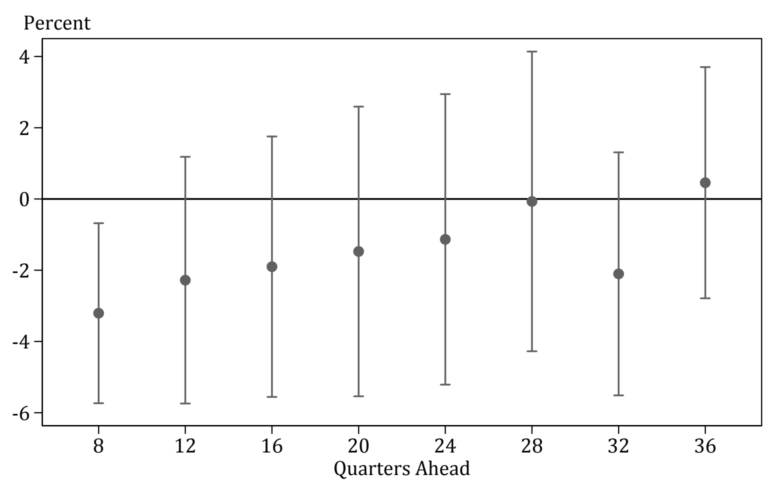
Figure 3: Panel Response of Exchange Rate Changes to US Monetary Policy Surprises for Hedge Currencies (2SLS)



Source: Authors' calculations.

Note: 90 percent confidence intervals based on Driscoll-Kraay standard errors. 2SLS regression of exchange rate change on relative forward rate changes instrumented using yield changes in a one-hour window around FOMC announcements of FF4, ED4, and two- and 10-year Treasury note futures expiring in the current quarter. This sample includes Australia, Canada, Norway, New Zealand, Sweden, and the United Kingdom against the US dollar.

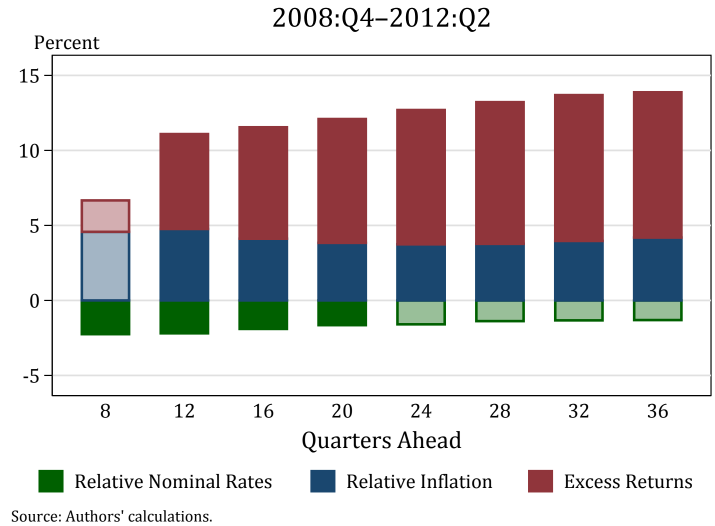
Figure 4: Panel Response of Exchange Rate Changes to US Monetary Policy Surprises for Non-Hedge Currencies (2SLS)



Source: Authors' calculations.

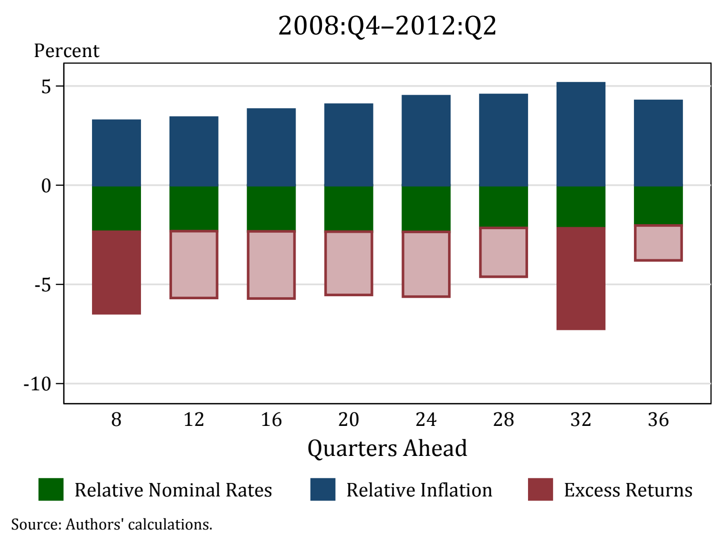
Note: 90 percent confidence intervals based on Driscoll-Kraay standard errors. 2SLS regression of exchange rate change on relative forward rate changes instrumented using yield changes in a one-hour window around FOMC announcements of FF4, ED4, and two- and 10-year Treasury note futures expiring in the current quarter. This sample includes the euro area, Japan, and Switzerland.

Figure 5: Panel Response of Exchange Rate Change Components to US Monetary Policy Surprises for Hedge Currencies (2SLS)



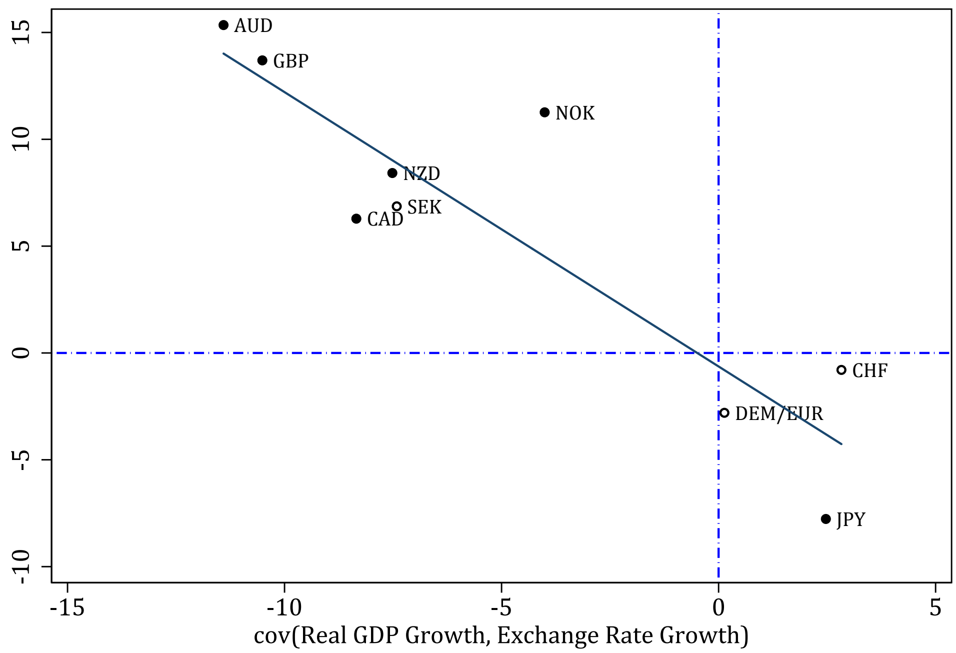
Note: Darker bar areas represent estimates significant at the 10 percent level based on Driscoll-Kraay standard errors. 2SLS regression of exchange rate change on relative forward rate changes instrumented using yield changes in a one-hour window around FOMC announcements of FF4, ED4, and two- and 10-year Treasury note futures expiring in the current quarter. This sample includes Australia, Canada, Norway, New Zealand, Sweden, and the United Kingdom against the US dollar.

Figure 6: Panel Response of Exchange Rate Change Components to US Monetary Policy Surprises for Non-Hedge Currencies (2SLS)



Note: Darker bar areas represent estimates significant at the 10 percent level based on Driscoll-Kraay standard errors. 2SLS regression of exchange rate change on relative forward rate changes instrumented using yield changes in a one-hour window around FOMC announcements of FF4, ED4, and two- and 10-year Treasury note futures expiring in the current quarter. This sample includes the euro area, Japan, and Switzerland against the US dollar.

Figure 7: Pair-Specific Response of Exchange Rate Change Risk Premia Component to US Monetary Policy Surprises versus Hedging Properties of the Dollar (2SLS)



Source: Authors' calculations.

Note: Filled circles represent significance at 10 percent. 2SLS regression of exchange rate change on relative forward rate changes instrumented using yield changes in a one-hour window around FOMC announcements of FF4, ED4, and two- and 10-year Treasury note futures expiring in the current quarter. The measure of the dollar's value as a hedge for each currency pair is the covariance between the respective exchange rate change and the US real GDP growth. This covariance is calculated over a longer sample starting in 1991:Q1.

Table 1: Panel Responses of the Exchange Rate Change and its Components to US Monetary Policy Surprises for All Currencies (2SLS)

Quarters Ahead	8	12	16	20	24	28	32	36
Δs_{t+1}	-0.11	3.42^{*}	4.86^{**}	6.21^{***}	7.21^{***}	8.48^{***}	7.66^{***}	8.30^{***}
	(1.49)	(2.05)	(1.92)	(1.88)	(1.89)	(2.34)	(1.79)	(2.34)
$\tilde{\imath}_t - \varphi_{t+1}^{EH}$	-2.31^{***}	-2.25^{***}	-2.08^{***}	-1.90^{***}	-1.83^{***}	-1.60^{**}	-1.60^{**}	-1.56^{**}
	(0.61)	(0.42)	(0.40)	(0.50)	(0.62)	(0.73)	(0.73)	(0.76)
$s_{t+1,\infty}^{\Delta E}$	3.79^{***}	4.10^{***}	3.99^{***}	3.92^{***}	3.98^{***}	3.99^{***}	4.35^{***}	4.21^{***}
	(0.98)	(0.93)	(0.84)	(0.88)	(0.98)	(1.05)	(1.20)	(1.22)
$\sigma_t - \sigma_{t+1}^{F}$	-1.59	1.58	2.95^{*}	4.19^{***}	5.07^{***}	6.09^{***}	4.92^{***}	5.66^{***}
	(1.31)	(1.62)	(1.56)	(1.46)	(1.36)	(1.52)	(1.35)	(1.72)

Note: Each cell of this table gives the slope coefficient from a 2SLS regression of the variable at the left on the change in the one-year relative forward rate n quarters hence ($\Delta \tilde{f}_{t+1}^{n}$) instrumented using yield changes in a one-hour window around FOMC announcements of FF4, ED4, and two- and 10-year Treasury note futures expiring in the current quarter. Driscoll-Kraay standard errors are in parentheses. Constants are included in the regression, but omitted from this table. This sample includes Australia, Canada, euro area, Japan, Norway, New Zealand, Sweden, Switzerland, and the United Kingdom against the US dollar.

Table 2: Panel Responses of the Exchange Rate Change and its Components to US Monetary Policy Surprises for Hedge Currencies (2SLS)

Quarters Ahead	8	12	16	20	24	28	32	36
Δs_{t+1}	4.42	8.90**	9.65***	10.44***	11.10***	11.84***	12.36***	12.57***
	(4.79)	(4.52)	(3.54)	(3.49)	(3.74)	(4.05)	(4.32)	(4.27)
$\tilde{\imath}_t - \varphi_{t+1}^{EH}$	−2.25**	−2.19***	−1.90***	−1.65*	−1.59	−1.38	−1.33	−1.31
	(1.08)	(0.73)	(0.74)	(0.85)	(0.98)	(1.12)	(1.21)	(1.25)
$s_{t+1,\infty}^{\Delta E}$	4.57	4.76***	4.11***	3.84***	3.74***	3.77***	3.97***	4.19***
	(2.83)	(1.76)	(1.16)	(1.04)	(1.07)	(1.16)	(1.31)	(1.48)
$\sigma_t - \sigma_{t+1}^F$	2.09	6.32**	7.43***	8.25***	8.95***	9.44***	9.72***	9.68***
	(2.02)	(2.56)	(2.15)	(2.07)	(2.21)	(2.39)	(2.53)	(2.40)

Note: Each cell of this table gives the slope coefficient from a 2SLS regression of the variable at the left on the change in the one-year relative forward rate n quarters hence ($\Delta\tilde{f}_{t+1}^n$) instrumented using yield changes in a one-hour window around FOMC announcements of FF4, ED4, and two- and 10-year Treasury note futures expiring in the current quarter. Driscoll-Kraay standard errors are in parentheses. Constants are included in the regression, but omitted from this table. This sample includes Australia, Canada, Norway, New Zealand, Sweden, and the United Kingdom against the US dollar.

Table 3: Panel Responses of the Exchange Rate Change and its Components to US Monetary Policy Surprises for Non-Hedge Currencies (2SLS)

Quarters Ahead	8	12	16	20	24	28	32	36
Δs_{t+1}	−3.21**	−2.28	−1.90	−1.47	−1.13	−0.07	−2.10	0.46
	(1.54)	(2.11)	(2.22)	(2.47)	(2.48)	(2.56)	(2.07)	(1.97)
$\tilde{\imath}_t - \varphi_{t+1}^{EH}$	−2.35***	−2.32***	−2.33***	−2.34***	−2.35***	−2.15***	−2.17***	−2.03***
	(0.44)	(0.42)	(0.42)	(0.50)	(0.59)	(0.74)	(0.74)	(0.71)
$s_{t+1,\infty}^{\Delta E}$	3.25***	3.41***	3.81***	4.06***	4.48***	4.55***	5.14***	4.25***
	(0.28)	(0.55)	(0.67)	(0.75)	(0.81)	(0.85)	(0.88)	(0.86)
$\sigma_t - \sigma_{t+1}^F$	−4.11**	−3.37	−3.39	−3.19	−3.26	−2.46	−5.07*	−1.76
	(1.61)	(2.38)	(2.65)	(3.06)	(3.30)	(3.48)	(3.00)	(2.86)

Note: Each cell of this table gives the slope coefficient from a 2SLS regression of the variable at the left on the change in the one-year relative forward rate n quarters hence ($\Delta\tilde{f}_{t+1}^n$) instrumented using yield changes in a one-hour window around FOMC announcements of FF4, ED4, and two- and 10-year Treasury note futures expiring in the current quarter. Driscoll-Kraay standard errors are in parentheses. Constants are included in the regression, but omitted from this table. This sample includes the euro area, Japan, and Switzerland against the US dollar.

Table 4: Response of Changes in Risk Aversion to US Monetary Policy Surprises (2SLS)

Quarters Ahead	8	12	16	20	24	28	32	36
$\Delta f_{t+1}^{n,US}$	-34.17^{***}	-29.39^{**}	-27.89^{**}	-28.69^{**}	-31.19^{**}	-34.94^{**}	-39.34^{**}	-43.42^{**}
	(11.24)	(12.75)	(13.33)	(14.09)	(15.27)	(16.65)	(17.62)	(17.26)
# of Observations	15	15	15	15	15	15	15	15

Note: Each column of this table gives the coefficients from regressing the estimated risk aversion series from Bekaert, Engstrom, and Xu (2017) on the change in the one-year US forward rate n quarters hence (Δf_{t+1}^{n}) instrumented using yield changes in a one-hour window around FOMC announcements of FF4, ED4, and two- and 10-year Treasury note futures expiring in the current quarter. Newey-West standard errors are in parentheses. Constants are included in the regression, but omitted from this table.

Table 5: Response of Changes in Expectations over the VIX Path to US Monetary Policy Surprises (2SLS)

Quarters Ahead	8	12	16	20	24	28	32	36
$\Delta f_{t+1}^{n,US}$	-2.13^{***}	-1.76^{***}	-1.63^{**}	-1.66^{**}	-1.80^{**}	-2.05^{**}	-2.35^{***}	-2.67^{***}
	(0.61)	(0.68)	(0.69)	(0.72)	(0.77)	(0.82)	(0.85)	(0.81)
# of Observations	15	15	15	15	15	15	15	15

Note: Each column of this table gives the coefficients from regressing $\sum_{k=1}^{\infty}(E[VIX_{t+k}|\mathcal{I}_{t+1}] - E[VIX_{t+k}|\mathcal{I}_t])$ (normalized to have a mean of 0 and a standard deviation of 1) on the change in the one-year US forward rate n quarters hence (Δf_{t+1}^{n}) instrumented using yield changes in a one-hour window around FOMC announcements of FF4, ED4, and two- and 10-year Treasury note futures expiring in the current quarter. Newey-West standard errors are in parentheses. Constants are included in the regression, but omitted from this table.

Table 6: Response of Changes in US Net Foreign Asset Position to US Monetary Policy Surprises (2SLS)

Quarters Ahead	8	12	16	20	24	28	32	36
$\Delta f_{t+1}^{n,US}$	10.11^{***}	9.90^{***}	10.28^{***}	11.25^{***}	12.74^{***}	14.61^{***}	16.57^{***}	18.14^{***}
	(3.82)	(3.73)	(3.70)	(3.76)	(3.97)	(4.35)	(4.84)	(5.27)
# of Observations	15	15	15	15	15	15	15	15

Note: Each column of this table gives the coefficients from regressing the change in the US net international investment position as a percent of the average US GDP over the Global Recession on the change in the one-year US forward rate n quarters hence (Δf_{t+1}^{n}) instrumented using yield changes in a one-hour window around FOMC announcements of FF4, ED4, and two- and 10-year Treasury note futures expiring in the current quarter. Newey-West standard errors are in parentheses. Constants are included in the regression, but omitted from this table.

Table 7: Response of US External Valuation Gain to US Monetary Policy Surprises (2SLS)

Quarters Ahead	8	12	16	20	24	28	32	36
$\Delta f_{t+1}^{n,US}$	9.41**	9.30**	9.71***	10.67***	12.11***	13.90***	15.76***	17.24***
	(3.76)	(3.63)	(3.59)	(3.65)	(3.84)	(4.20)	(4.68)	(5.12)
# of Observations	15	15	15	15	15	15	15	15

Note: Each column of this table gives the coefficients from regressing the US external valuation gain on the change in the one-year US forward rate n quarters hence (Δf_{t+1}^{n}) instrumented using yield changes in a one-hour window around FOMC announcements of FF4, ED4, and two- and 10-year Treasury note futures expiring in the current quarter. This external valuation gain is computed as the change in the US net international investment position minus the current account balance. This gain is scaled as a percent of the average US GDP over the Global Recession. Newey-West standard errors are in parentheses. Constants are included in the regression, but omitted from this table.

Table 8: Response of US GDP Forecast Revisions to US Monetary Policy Surprises (2SLS)

Quarters Ahead	8	12	16	20	24	28	32	36
$\Delta f_{t+1}^{n,US}$	0.81***	0.73***	0.71***	0.74***	0.80***	0.88***	0.96***	1.03***
	(0.09)	(0.12)	(0.13)	(0.15)	(0.16)	(0.18)	(0.18)	(0.17)
# of Observations	15	15	15	15	15	15	15	15

Note: Each column of this table gives the coefficients from regressing the revision in the *Blue Chip Financial Forecasts* four-quarter-ahead GDP forecast on the change in the one-year US forward rate n quarters hence (Δf_{t+1}^{n}) instrumented using yield changes in a one-hour window around FOMC announcements of FF4, ED4, and two- and 10-year Treasury note futures expiring in the current quarter. Newey-West standard errors are in parentheses. Constants are included in the regression, but omitted from this table.

Table 9: Subsample Means of Uncertainty Measures

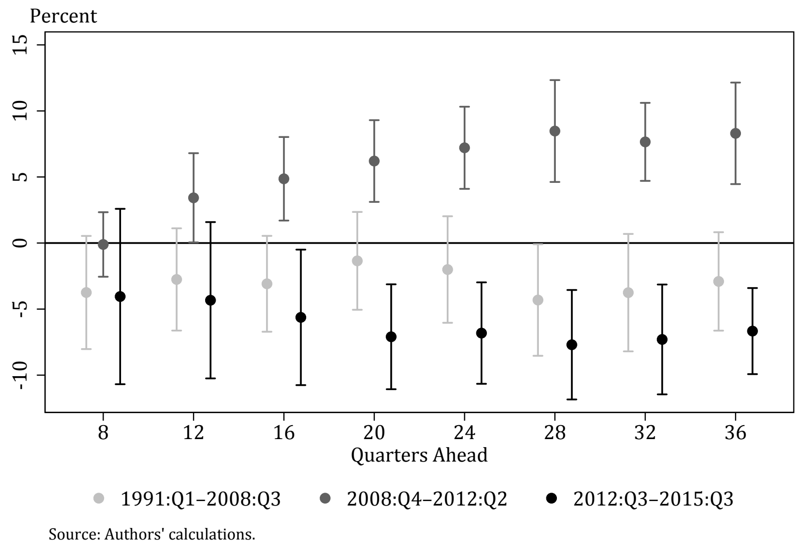
	1990:Q3-2008:Q3	2008:Q4-2012:Q2	2012:Q3-2015:Q3
JLN Macro Uncertainty	-0.04	0.80	-0.67
GDP Forecast Dispersion	0.04	0.88	-1.24
BBD Monetary Policy Uncertainty	0.12	-0.06	-0.59

Note: The JLN macro uncertainty measure is 12-month ahead macroeconomic uncertainty estimated by Jurado, Ludvigson, and Ng (2015). GDP forecast dispersion is the 25th–75th percentile range of four-quarter-ahead US real GDP forecasts from *Blue Chip Financial Forecasts*. BBD monetary policy uncertainty is the monetary policy subcomponent of the Baker, Bloom, and Davis (2016) policy uncertainty index. All three measures are standardized over the full 1991:Q1–2015:Q3 sample to facilitate interpretation.

Appendix

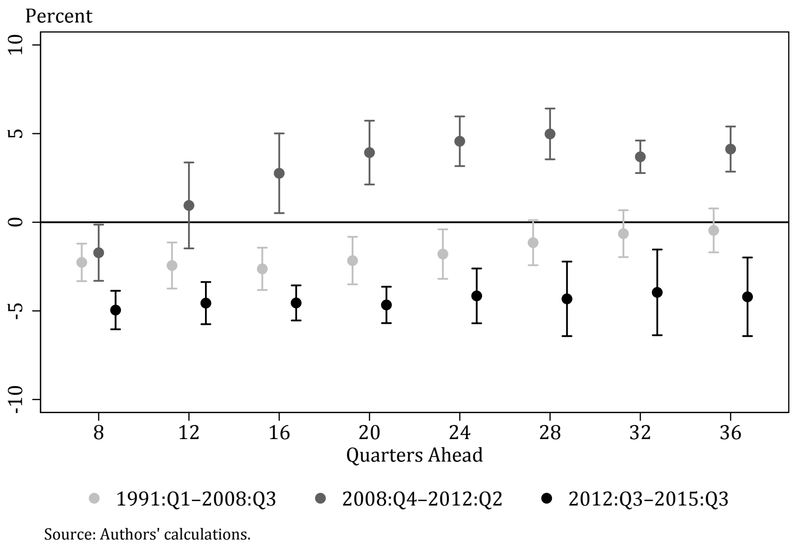
A Additional Figures and Tables

Figure A-1: Panel Response of Exchange Rate Changes to US Monetary Policy Surprises for All Currencies (2SLS)



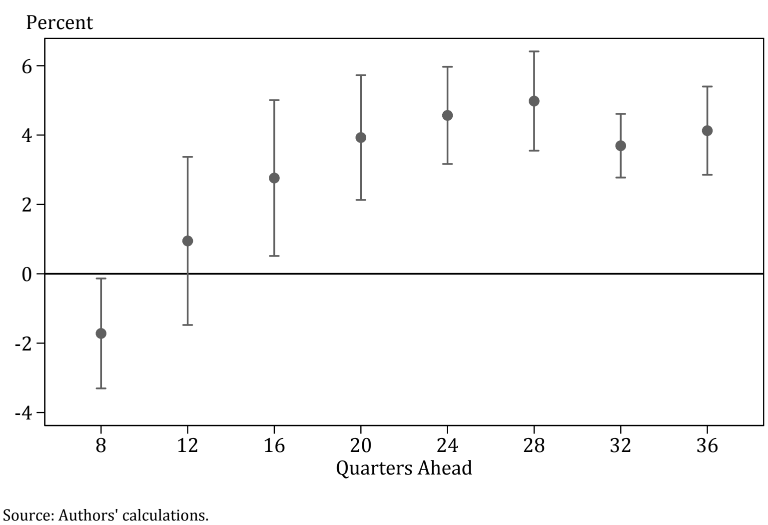
Note: 90 percent confidence intervals based on Driscoll-Kraay standard errors. 2SLS regression of exchange rate change on relative forward rate changes instrumented using yield changes in a one-hour window around FOMC announcements of FF4, ED4, and two- and 10-year Treasury note futures expiring in the current quarter. This sample includes Australia, Canada, euro area, Japan, Norway, New Zealand, Sweden, Switzerland, and the United Kingdom against the US dollar.

Figure A-2: Panel Regression of Exchange Rate Changes on Relative Forward Rate Changes for All Currencies (OLS)



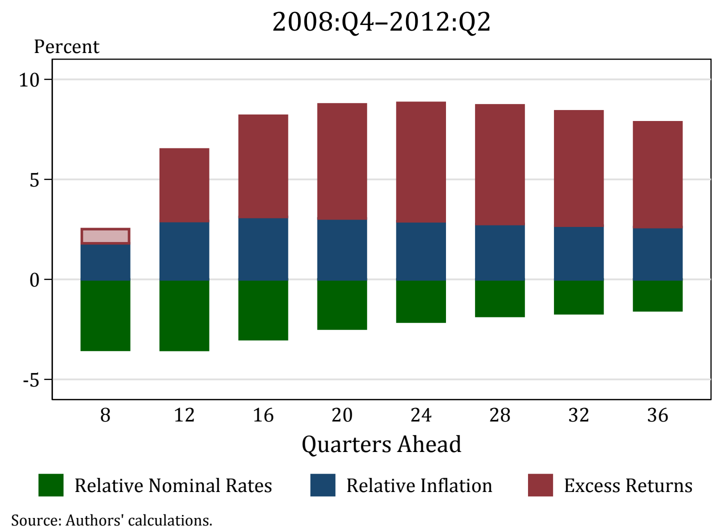
Note: 90 percent confidence intervals based on Driscoll-Kraay standard errors. OLS regression of exchange rate change on relative forward rate changes. This sample includes Australia, Canada, euro area, Japan, Norway, New Zealand, Sweden, Switzerland, and the United Kingdom against the US dollar.

Figure A-3: Panel Regression of Exchange Rate Changes on Relative Forward Rate Changes for All Currencies (OLS)



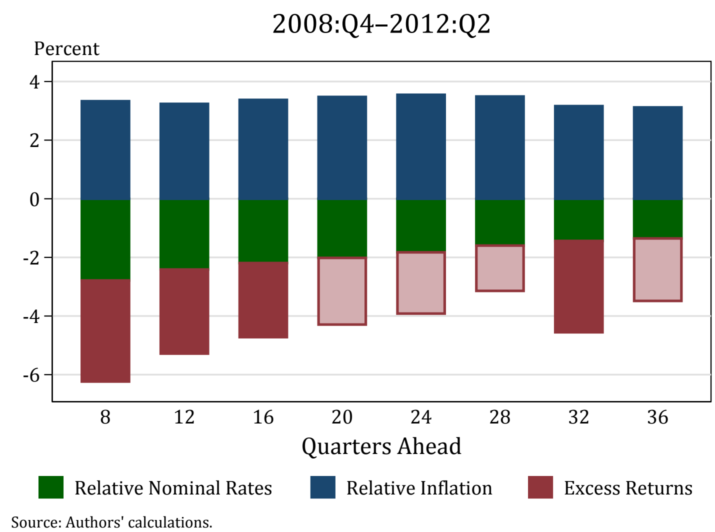
Note: 90 percent confidence intervals based on Driscoll-Kraay standard errors. OLS regression of exchange rate change on relative forward rate changes. This sample includes Australia, Canada, euro area, Japan, Norway, New Zealand, Sweden, Switzerland, and the United Kingdom against the US dollar.

Figure A-4: Panel Regression of Exchange Rate Changes on Relative Forward Rate Changes for Hedge Currencies (OLS)



Note: Darker bar areas represent estimates significant at the 10 percent level based on Driscoll-Kraay standard errors. OLS regression of exchange rate change on relative forward rate changes. This sample includes Australia, Canada, Norway, New Zealand, Sweden, and the United Kingdom against the US dollar.

Figure A-5: Panel Regression of Exchange Rate Changes on Relative Forward Rate Changes for Non-Hedge Currencies (OLS)



Note: Darker bar areas represent estimates significant at the 10 percent level based on Driscoll-Kraay standard errors. OLS regression of exchange rate change on relative forward rate changes. This sample includes the euro area, Japan, and Switzerland against the US dollar.

Table A-1: Panel Regression of Exchange Rate Change and its Components on Relative Forward Rate Changes for All Currencies (OLS)

Quarters Ahead	8	12	16	20	24	28	32	36
Δs_{t+1}	-1.72^{*}	0.95	2.76^{**}	3.93^{***}	4.57^{***}	4.98^{***}	3.69^{***}	4.13^{***}
	(0.96)	(1.47)	(1.37)	(1.09)	(0.85)	(0.87)	(0.56)	(0.77)
$\tilde{\imath}_t - \varphi^{EH}_{t+1}$	-3.25^{***}	-3.08^{***}	-2.69^{***}	-2.31^{***}	-2.03^{***}	-1.76^{***}	-1.60^{***}	-1.48^{***}
	(0.44)	(0.39)	(0.42)	(0.48)	(0.50)	(0.48)	(0.39)	(0.37)
$s^{\Delta E}_{t+1,\infty}$	2.37^{***}	3.05^{***}	3.22^{***}	3.19^{***}	3.10^{***}	2.98^{***}	2.86^{***}	2.78^{***}
	(0.69)	(0.60)	(0.53)	(0.54)	(0.58)	(0.56)	(0.62)	(0.59)
$\sigma_t - \sigma^F_{t+1}$	-0.85	0.98	2.24^{**}	3.05^{***}	3.49^{***}	3.76^{***}	2.43^{***}	2.83^{***}
	(0.95)	(1.07)	(0.97)	(0.68)	(0.45)	(0.39)	(0.40)	(0.34)

Note: Each cell of this table gives the slope coefficient from an OLS regression of the variable at the left on the change in the one-year relative forward rate n quarters hence ($\Delta \tilde{f}^n_{t+1}$). Driscoll-Kraay standard errors are in parentheses. Constants are included in the regression, but omitted from this table. This sample includes Australia, Canada, the euro area, Japan, Norway, New Zealand, Sweden, Switzerland, and the United Kingdom against the US dollar.

Table A-2: Panel Regression of Exchange Rate Change and its Components on Relative Forward Rate Changes for Hedge Currencies (OLS)

Quarters Ahead	8	12	16	20	24	28	32	36
Δs_{t+1}	-1.01	2.97	5.19^{**}	6.29^{***}	6.71^{***}	6.88^{***}	6.71^{***}	6.31^{***}
	(1.61)	(2.83)	(2.53)	(2.10)	(1.84)	(1.72)	(1.59)	(1.41)
$\tilde{\imath}_t - \varphi^{EH}_{t+1}$	-3.52^{***}	-3.53^{***}	-2.99^{***}	-2.46^{***}	-2.11^{***}	-1.83^{***}	-1.69^{***}	-1.55^{***}
	(0.34)	(0.38)	(0.50)	(0.54)	(0.51)	(0.46)	(0.41)	(0.36)
$s^{\Delta E}_{t+1,\infty}$	1.80^{**}	2.92^{***}	3.12^{***}	3.05^{***}	2.90^{***}	2.77^{***}	2.69^{***}	2.62^{***}
	(0.82)	(0.81)	(0.58)	(0.51)	(0.51)	(0.51)	(0.53)	(0.56)
$\sigma_t - \sigma^F_{t+1}$	0.71	3.58^{*}	5.06^{***}	5.70^{***}	5.92^{***}	5.93^{***}	5.72^{***}	5.24^{***}
	(1.24)	(2.12)	(1.87)	(1.48)	(1.27)	(1.16)	(1.02)	(0.83)

Note: Each cell of this table gives the slope coefficient from an OLS regression of the variable at the left on the change in the one-year relative forward rate n quarters hence ($\Delta \tilde{f}^n_{t+1}$). Driscoll-Kraay standard errors are in parentheses. Constants are included in the regression, but omitted from this table. This sample includes Australia, Canada, Norway, New Zealand, Sweden, and the United Kingdom against the US dollar.

Table A-3: Panel Regression of Exchange Rate Change and its Components on Relative Forward Rate Changes for Non-Hedge Currencies (OLS)

Quarters Ahead	8	12	16	20	24	28	32	36
Δs_{t+1}	-2.91^{***}	-2.04^{**}	-1.35	-0.82	-0.37	0.35	-1.39	-0.37
	(0.99)	(0.95)	(0.89)	(0.98)	(1.27)	(1.22)	(1.47)	(1.26)
$\tilde{\imath}_t - \varphi_{t+1}^{EH}$	-2.79^{***}	-2.41^{***}	-2.19^{***}	-2.02^{***}	-1.83^{***}	-1.60^{**}	-1.44^{***}	-1.35^{**}
	(0.58)	(0.36)	(0.40)	(0.52)	(0.66)	(0.73)	(0.53)	(0.55)
$s_{t+1,\infty}^{\Delta E}$	3.33^{***}	3.24^{***}	3.37^{***}	3.47^{***}	3.55^{***}	3.49^{***}	3.16^{***}	3.12^{***}
	(0.20)	(0.33)	(0.46)	(0.58)	(0.74)	(0.64)	(0.77)	(0.66)
$\sigma_t - \sigma_{t+1}^F$	-3.45^{***}	-2.87^{**}	-2.53^{*}	-2.27	-2.09	-1.55	-3.12^{*}	-2.13
	(1.15)	(1.23)	(1.39)	(1.56)	(1.88)	(1.59)	(1.85)	(1.52)

Note: Each cell of this table gives the slope coefficient from an OLS regression of the variable at the left on the change in the one-year relative forward rate n quarters hence ($\Delta\tilde{f}_{t+1}^{n}$). Driscoll-Kraay standard errors are in parentheses. Constants are included in the regression, but omitted from this table. This sample includes euro area, Japan, and Switzerland against the US dollar.

B Data Description

B.1 Macroeconomic and Financial Variables

- *Exchange rates*: End-of-quarter exchange rates are computed using daily data from Global Financial Data.
- *Short-term rates*: End-of-quarter three-month bill rates are obtained from the following sources:
  - Australia, Canada, New Zealand, Norway, Sweden, Switzerland, United Kingdom, and United States: Central bank data obtained through Haver Analytics.
  - Germany: Reuters data obtained through Haver Analytics. German three-month bill rates are replaced with three-month EONIA OIS swap rates starting in 1999:Q1.
  - Japan: Bloomberg.
- *Zero-coupon yields:* End-of-quarter zero-coupon yields are obtained from the following sources:
  - Canada, Germany, Sweden, Switzerland, and United Kingdom: Central banks. German zero-coupon bond yields are replaced with estimates of zero-coupon yields on AAA-rated euro area sovereign debt provided by the European Central Bank (ECB).
  - Norway: Data from Wright (2011) extended with data from the BIS.
  - Australia, New Zealand: Data from Wright (2011) extended with data from central banks.
  - Japan: Bloomberg.
  - United States: Gürkaynak, Sack, and Wright (2007).
- *US real GDP growth forecasts*: Consensus (mean) forecasts from *Blue Chip Financial Forecasts.*

Though this paper focuses mainly on period of the Global Recession, which we define as 2008:Q4–2012:Q2, our full data sample (which is used to estimate the exchange rate decomposition and hedging properties of the dollar) is as follows for each currency pair. Note that we exclude periods of fixed exchange rates:

Data Sample Ranges

Australia	1989:Q4 – 2015:Q4
Canada	1992:Q2 – 2015:Q4
Germany	1991:Q2 – 2015:Q4
Japan	1992:Q3 – 2015:Q4
New Zealand	1990:Q1 – 2015:Q4
Norway	1989:Q4 – 2015:Q4
Sweden	1992:Q4 – 2015:Q4
Switzerland	1992:Q1 – 2011:Q2
United Kingdom	1992:Q4 – 2015:Q4
United States	1989:Q4 – 2015:Q4

B.2 US Policy Surprises

To capture policy surprises, we use changes in yields implied by futures prices over one-hour windows starting 15 minutes before and ending 45 minutes after FOMC policy announcements. The set of events that we consider are announcements made after scheduled FOMC meetings, announcements after unscheduled FOMC meetings in which a policy target change was made (most relevant for the pre-1994 period when statements were not released following FOMC meetings), and important QE announcements identified by the literature (see Section B.3). In keeping with the rest of the literature, we exclude the September 17, 2001 statement accompanying a conference call held in response to the September 11 attacks.

The times of the FOMC announcements are obtained from the data appendix of Gürkaynak, Sack, and Swanson (2005) and are updated using the press release and meeting calendars available at https://www.federalreserve.gov/monetarypolicy.htm.[40]

The futures that we consider are the 30-day federal funds futures contract expiring three months hence (FF4), the eurodollar futures contract expiring three quarters hence (ED4), the two-year Treasury futures contract expiring in the current quarter, and the 10-year Treasury futures contract expiring in the current quarter. For FF4 and ED4, the implied yield is simply 100 minus the futures price. For the Treasury futures, the implied yield is computed as the yield to maturity implied by the futures price, which is based on the delivery of Treasury securities with the designated maturity and a 6 percent per annum semiannual coupon. The calculation also takes into account that prior to January 1, 2000, futures prices were based on an 8 percent per annum semiannual coupon. This conversion of prices to yields is important due to this change in the notional coupon rate of the securities

[40]For the January 22, 2008 FOMC statement, the time of the announcement is not available from the Federal Reserve's website. Multiple news sources state that this announcement was made in the morning, prior to the opening of US stock markets. Therefore, we use a time window of 7:00am to 9:30am EST to measure the surprise for this event.

underlying Treasury futures contracts. The same change in the *yield* would correspond to a smaller change in Treasury futures *prices* prior to January 1, 2000 than after this date, due simply to the higher notional coupon rate embedded in the futures contract.

The data for FF4 and ED4, up to June 2012, were generously provided to us by Refet Gürkaynak. We extended his data past June 2012 and with additional QE announcements using intra-day data from Tick Data. Data on FF4 for the November 25, 2008 and December 1, 2008 QE announcements were obtained from the Chicago Mercantile Exchange (CME), since Tick Data's coverage of this security does not begin until January 2010.

For Treasury futures, we use the front contract from Tick Data, as in Wright (2012).

All intra-day surprises are summed over quarters for the empirical exercises at a quarterly frequency.

B.3 QE Announcement Dates

The following list of QE dates are collected from a number of papers including Rogers, Scotti, and Wright (2014), Wu (2014), and Swanson (2017).

Table A-4: QE Announcements

Date	Description
11/25/2008	Initial large-scale-asset-purchase (LSAP) announcement.
12/1/2008	Bernanke states Treasuries may be purchased.
12/16/2008	The FOMC indicated that "it stands ready to expand its purchases of agency debt and mortgage-backed securities as conditions warrant. The Committee is also evaluating the potential benefits of purchasing longer-term Treasury securities."
1/28/2009	FOMC Statement.
3/18/2009	FOMC announces it will purchase \$750B of mortgage-backed securities, \$300B of longer-term Treasuries, and \$100B of agency debt (a.k.a. "QE1").
8/12/2009	The FOMC eliminated the "up to" phrase in its intended purchase amount of Treasury securities. It also stated that it would "slow the pace of these transactions and anticipates that the full amount will be purchased by the end of October.
9/23/2009	The FOMC eliminated the "up to" phrase in its intended purchase amount of the MBS, as well as its plan to "slow the pace of these purchases in order to promote a smooth transition in markets and anticipates that they will be executed by the end of the first quarter of 2010."

11/4/2009	The FOMC clarified that the intended purchase amount of agency debt is \$175 billion, instead of "up to \$200 billion", as previously announced.
8/10/2010	The FOMC announced that it "will keep constant the Federal Reserve's holdings of securities at their current level by reinvesting principal payments from agency debt and agency mortgage-backed securities in longer-term Treasury securities. The Committee will continue to roll over the Federal Reserve's holdings of Treasury securities as they mature."
8/27/2010	Bernanke Speech at Jackson Hole.
9/21/2010	FOMC Statement.
10/15/2010	Bernanke Speech at the Boston Fed.
11/3/2010	FOMC announces it will purchase an additional \$600B of longer-term Treasuries (a.k.a. "QE2").
8/26/2011	Bernanke Speech at Jackson Hole.
9/21/2011	FOMC announces it will sell \$400B of short-term Treasuries and use the proceeds to buy \$400B of long-term Treasuries (a.k.a. "Operation Twist").
6/20/2012	The FOMC announced its intention "to continue through the end of the year its program to extend the average maturity of its holdings of securities."
9/13/2012	FOMC announces it will purchase \$40B of mortgage-backed securities per month for the indefinite future.
12/12/2012	FOMC announces it will purchase \$45B of longer-term Treasuries per month for the indefinite future.
5/22/2013	Bernanke Congressional Testimony ("Taper Tantrum").
6/19/2013	FOMC Statement.
12/18/2013	FOMC announces it will start to taper purchases of longer-term Treasuries and mortgage-backed securities to \$40B and \$35B per month, respectively.

C Break Date Estimation

To estimate break dates, we follow the procedure of Bai and Perron (1998) using OLS estimation of equation (1). Though our main interest is in the two-stage least squares estimate, Perron and Yamamoto (2015) argue that estimating break dates using OLS is generally more precise.

The procedure involves searching over a grid of possible break dates, for a predefined number of breaks, to find the set that minimizes the regression's sum of squared residuals (SSR). We do this for one, two, and three breaks. We search for breaks using a sample from 1991:Q1 to 2015:Q3 and set a minimum subsample length of 10 quarters, which corresponds to about 10 percent of our sample. Table A-5 presents the optimal break dates for each forward rate horizon considered, while the dashed lines in Figure A-6 plot the resulting SSRs as a ratio of the SSRs achieved by not allowing for a break in the estimated coefficients.

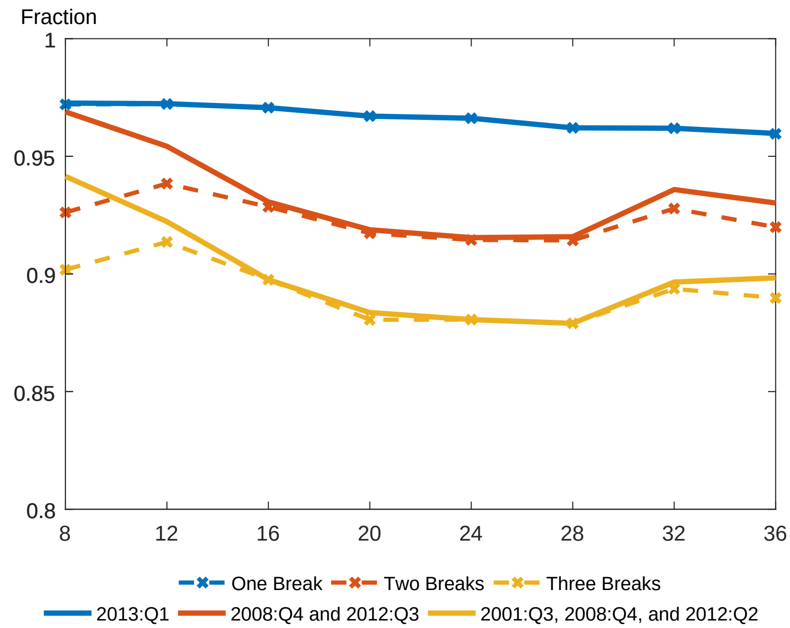
Table A-5: Break Dates that Minimize Sum of Squared Residuals

Quarters Ahead	One Break	Two Breaks	Three Breaks
8	2012:Q4	2002:Q2, 2005:Q1	1995:Q2, 2001:Q3, 2005:Q1
12	2012:Q4	2002:Q2, 2005:Q1	2002:Q2, 2005:Q1, 2012:Q4
16	2013:Q1	2008:Q4, 2012:Q2	2001:Q3, 2008:Q4, 2012:Q2
20	2013:Q1	2008:Q4, 2012:Q2	2006:Q2, 2008:Q4, 2012:Q2
24	2013:Q1	2008:Q4, 2012:Q2	2001:Q3, 2008:Q4, 2012:Q2
28	2013:Q1	2008:Q4, 2012:Q2	2001:Q3, 2008:Q4, 2012:Q2
32	2013:Q1	2001:Q3, 2012:Q4	2001:Q3, 2008:Q4, 2012:Q4
36	2012:Q4	2002:Q1, 2012:Q4	1997:Q1, 2001:Q3, 2012:Q4

Note: Break dates given are the start dates of subsamples.

Note that for most horizons, including the longer ones which we are mainly interested in, the largest incremental improvement in SSRs is achieved when we move from one to two breaks (as opposed to moving from zero to one break or two to three breaks). The set of two break dates that occurs most commonly, particularly for longer horizons, is 2008:Q4 and 2012:Q2. For the longest horizons, 2012:Q4 also occurs as a break date. These results show that the relationship between exchange rate changes and changes in relative forward rates over our chosen sample of 2008:Q4–2012:Q2 is indeed different than the behavior in other time periods. The red solid line in Figure A-6 plots the relative SSRs obtained when we apply these two break dates to all horizons. Note that for horizons equal to or above 12 quarters, the SSR achieved using these two break dates is very close to the ones achieved using the

Figure A-6: Sums of Squared Residuals Relative to No Break Case



Source: Authors' calculations.

Note: Dashed lines are the SSRs relative to the case of no breaks for the optimal one, two, or three break dates for each horizon (shown in Table A-5). Solid lines are the relative SSRs for each horizon at the break dates shown in the legend.

optimal horizon-specific breaks shown in Table A-5. Figure A-6 also plots the relative SSRs for the most commonly found single break and set of three breaks across horizons.

When we allow 2001:Q3 to serve as a third break date in our regressions, the coefficient estimates from the first two subsamples are very similar, particularly for longer horizons.

D Summary of Survey-Based VAR Specification

We use the exchange rate decomposition in Stavrakeva and Tang (2018b) to decompose the estimated coefficients describing the relationship between exchange rate changes and yields, conditional on US monetary policy shocks. We briefly summarize the survey-based VAR specification here, but interested readers should see Stavrakeva and Tang (2018b) for more details.

For each country i, we estimate a VAR with two lags and the following companion form:

$$X_{t+1} = \bar{X} + \Gamma X_t + \Xi_{t+1}.$$

The vector X_{t+1} contains the real exchange rate level in addition to the following variables both for country i and the United States: three-month bill rates, the slopes and curvatures of the yield curve, defined as

$$\begin{aligned} slope^i &= y^{40,i} - i^i \\ curve^i &= 2y^{8,i} - (y^{40,i} + i^i), \end{aligned}$$

as well as CPI inflation, GDP gap, and current-account-to-GDP ratio. We additionally include the US VIX, TED spread (spread between three-month US LIBOR and treasury bill rates), and an exponential moving average of US inflation computed as

$$\pi_{t+1}^{avg,US} = \rho \pi_t^{avg,US} + (1-\rho)\pi_{t-1}^{US}.$$

This VAR is estimated subject to the following constraint

$$Y_t^S = H_t(\bar{X}, \Gamma) X_t + H_t^Z Z_t + \Xi_{h,t}^S, \tag{A-1}$$

where Y_t^S are survey forecasts and the right-hand-side are VAR-implied forecasts plus a measurement error term, $\Xi_{h,t}^S$.

A number of restrictions are placed on the coefficient matrix Γ, including that each country's financial variables follow a small VAR. This can be interpreted as a version of country-specific three-factor affine term structure models. We further assume that US variables are not affected by other countries, though changes in conditions in the United States can spill over into the macroeconomies of other countries. Lastly, we restrict real exchange rates to enter as a lag only in its own equation.

D.1 Additional Macroeconomic Data Used in the VAR

- *Output gap and current account-to-GDP ratio:* All macro data are from the OECD Main Economic Indicators and Economic Outlook databases. The GDP gap is computed using the OECD's annual estimates of potential GDP, which were log-linearly interpolated to the quarterly frequency.
- *CPI inflation:* Government statistical agencies.
- *VIX and three-month US TED spread:* Haver Analytics. The three-month US TED spread is calculated as the three-month US libor rate minus the three-month US bill rate.
- Macroeconomic data for Germany is replaced with Eurozone data starting in 1999:Q1.
- The US net foreign asset positions is from FRED: IIPUSNETIQ – U.S. Net International Investment Position, Millions of Dollars, Quarterly, Not Seasonally Adjusted

D.2 Survey Data

In the VAR, we include the following survey data for three-month interest rates, CPI inflation, and exchange rates:

Blue Chip Economic Indicators

- Countries: Australia, Canada, Germany/Eurozone, Japan, United Kingdom, United States.
- Date range: 1993:Q3–2015:Q4.
- Non-US variables: Current, one-, and two-years ahead forecasts of three-month interest rates, CPI inflation, and exchange rates.
- US variables: 7–11 year ahead average three-month bill rate (starting in 1990:Q1).
- Other details: Forecasts for German three-month interest rates and CPI inflation are replaced with Eurozone forecasts starting in January 2000, when they become available.

Blue Chip Financial Forecasts

- Countries: Australia, Canada, Germany/Eurozone, Japan, Switzerland, United Kingdom, United States.
- Date range: 1993:Q1–2015:Q4.
- Variables: Three-, six- and 12-month ahead three-month interest rate, 10-year yield, and exchange rate forecasts.

- Other details: Forecasts for German three-month interest rates and exchange rates are replaced with Eurozone forecasts starting in January 1999. Forecasts for the German 10-year yield are used throughout the sample since forecasts for AAA-rated euro area 10-year yields are not available.

Consensus Economics

- Country coverage: Australia, Canada, Germany/Eurozone, Japan, Norway, New Zealand, Sweden, Switzerland, United Kingdom, United States.
- Date range: 1990:Q1–2015:Q4.
- Variables: Current, one-, and two-years ahead and 6-10 year ahead average for CPI inflation; three- and 12-month ahead for three-month interest rates and 10-year yields; three-, 12-, and 24-month ahead for exchange rates. Forecasts for 6–10 year ahead average GDP growth rates are used to impute long-horizon non-US three-month bill rate forecasts, as described in Stavrakeva and Tang (2018b), but are not directly included in the VAR estimation.
- Other details: Forecasts for Germany are replaced with Eurozone forecasts as these become available. Short-horizon CPI inflation and three-month interest rate forecasts switch from Germany to Eurozone in December 2002 and January 2005, respectively. Long-horizon CPI inflation and GDP growth forecasts switch from Germany to Eurozone in April 2003.

Other details:

- All inflation forecasts are for price index changes calculated on an annual-average over annual-average basis. Annual interest rate and exchange rate forecasts are for end-of-year values. Months-ahead forecasts are for end-of-month values.
- Surveys are usually published within the first two weeks of the month and contain responses from survey participants from the end of the prior month. To map the survey data to our model, we backdate the survey variables (for example, a January publication is mapped to model-implied forecasts as of the end of Q4).
- CPI forecasts for the United Kingdom begin in 2004:Q2 in all databases. Previous inflation forecasts for the United Kingdom were for the retail price index.
- Three-month interest rate forecasts, for certain countries, are explicitly for interbank rates rather than bill rates. There are also cases where the survey does not specify the particular rate forecast by the respondents. To account for this, we allow data-source-specific constants in the rows of equation (A-1) that correspond to three-month interest rate forecast data. Though sometimes statistically significant, the estimated constants

are small and consistent with the average spreads between interbank and bill rates. When assessing the model's fit, we include this additional constant in the model-implied counterpart to forecasts of the *surveyed variable*. However, this additional constant is not considered to be part of the model-implied three-month *bill rate* forecasts.

E Model: Additional Derivations

E.1 Deriving the Euler Equation

Consider a two-country world and assume that there are exogenous endowments of the home country (the United States in our framework) and foreign tradable goods, Y_t^{us} and Y_t^i, and that there are no non-tradable goods. Small letters will denote logs of the variables. Consider a cashless economy, where the dollar prices of the tradable good in the United States and the foreign-currency-denominated prices of the foreign tradable good in the United States are P_t^{us} and P_t^i, respectively. The nominal exchange rate, given by S_t, is the relative price of one unit of currency i per one dollar.

The representative agent in the United States maximizes:

$$\max_{C_t^{us}, C_t^{us,i}, B_t^{us}, B_t^{us,i}} E\left[\sum_{l=0}^{\infty} \beta^{t+l}\left((1-\tau)\,u\left(C_{t+l}^{us,i}, X_{t+l}^{us,i}\right)+\tau u\left(C_{t+l}^{us}, X_{t+l}^{us}\right)\right)\middle|\,\mathcal{I}_t\right],$$

where C_t^{us} and $C_t^{us,i}$ represent her consumption of the US tradable good and the tradable good of country i, while X_t^{us} and $X_t^{us,i}$ are the respective habit reference levels of consumption. The degree of home bias is $0 \leq \tau \leq 1$. We consider the limiting case where $\tau \to 1$ and the US economy becomes approximately closed. The representative agent's optimization problem is subject to the standard budget constraint:

$$C_t^{us,i}\frac{P_t^i}{S_t}+\frac{B_t^{us,i}}{S_t}+C_t^{us}P_t^{us}+B_t^{us} \leq P_t^{us}Y_t^{us}+\left(1+i_{t-1}^{us}\right)B_{t-1}^{us}+\left(1+i_{t-1}^{i}\right)\frac{B_{t-1}^{us,i}}{S_t} \quad [\mu_t],$$

where B_t^{us} and $B_t^{us,i}$ are the savings in the dollar and foreign-currency denominated bonds. The Lagrangian can be expressed as:

$$\max_{C_t^{us}, C_t^{us,i}, B_t^{us}, B_t^{us,i}} E\sum_{l=0}^{\infty}\beta^{t+l}\left[\begin{array}{c} (1-\tau)\,u\left(C_{t+l}^{us,i}, X_{t+l}^{us,i}\right)+\tau u\left(C_{t+l}^{us}, X_{t+l}^{us}\right)+ \\ \mu_t\left(\begin{array}{c} P_{t+l}^{us}Y_{t+l}^{us}+\left(1+i_{t+l-1}^{us}\right)B_{t+l-1}^{us}+\left(1+i_{t+l-1}^{i}\right)\frac{B_{t+l-1}^{us,i}}{S_{t+l}} \\ -C_{t+l}^{us,i}\frac{P_{t+l}^i}{S_{t+l}}-\frac{B_{t+l}^{us,i}}{S_{t+l}}-C_{t+l}^{us}P_{t+l}^{us}-B_{t+l}^{us}\end{array}\right)\end{array}\middle|\,\mathcal{I}_t\right].$$

The first-order conditions are given by:

$$
\begin{aligned}
C_t^{us} &: \tau u_c\left(C_t^{us}, X_t^{us}\right) = \mu_t P_t^{us}, \\
C_t^{us,i} &: (1-\tau)\, u_c\left(C_t^{us,i}, X_t^{us,i}\right) = \mu_t \frac{P_t^i}{S_t}, \\
B_t^{us} &: \mu_t = E\left[\mu_{t+1}\beta\left(1+i_t^{us}\right)\middle|\, \mathcal{I}_t\right], \\
B_t^{us,i} &: \mu_t = E\left[\mu_{t+1}\beta\left(1+i_t^{i}\right)\frac{S_t}{S_{t+1}}\middle|\, \mathcal{I}_t\right],
\end{aligned}
$$

and can be re-written as follows:

$$
E\left[\frac{\mu_{t+1}}{\mu_t}\beta\left(\left(1+i_t^{us}\right)-\left(1+i_t^{i}\right)\frac{S_t}{S_{t+1}}\right)\middle|\, \mathcal{I}_t\right] = 0. \tag{A-2}
$$

In our limiting case where the US economy is approximately closed, i.e. $\tau \to 1$, the stochastic discount factor is given by:

$$
\begin{aligned}
\frac{\mu_{t+1}}{\mu_t} &= \frac{u_c\left(C_{t+1}^{us}, X_{t+1}^{us}\right)}{u_c\left(C_t^{us}, X_t^{us}\right)}\frac{P_t^{us}}{P_{t+1}^{us}} \\
&= \frac{u_c\left(C_{t+1}^{us}, X_{t+1}^{us}\right)}{u_c\left(C_t^{us}, X_t^{us}\right)} e^{-\pi_{t+1}^{us}}.
\end{aligned} \tag{A-3}
$$

Combining equations A-2 and A-3 gives equation 10 in the main text.

The optimization problem of the foreign consumer is purposefully left unspecified and does not have to be symmetric. We also assume that, in the long run, the weak form of purchasing power parity holds; i.e. $\lim_{k\to\infty} \frac{\tilde{P}_{t+k}^{us,i}}{P_{t+k}^{us} S_{t+k}} = c$ where $c > 0$ is some constant and $\tilde{P}_{t+k}^{us,i}$ is the price of the US tradable good in units of currency i in country i. We also define foreign inflation to be import price inflation as follows: $\pi_{t+k}^i = \Delta\tilde{p}_{t+k}^{us,i}$.

E.2 Agent's Signal Processing Problem

The US central bank's signal can be decomposed as:

$$
a_{t+1} = i_{t+h}^{us} = a_{t+h}^{y} + a_{t+h}^{mp},
$$

$$
\text{where } a_{t+h}^{y} \equiv \frac{\kappa\varepsilon_{t+h}^{y}}{\eta+\nu\kappa} \quad \text{and} \quad a_{t+h}^{mp} \equiv \frac{\eta\varepsilon_{t+h}^{mp}}{\eta+\nu\kappa}.
$$

Note that a^y_{t+h} and a^{mp}_{t+h} are both mean zero and i.i.d. normal. Thus, the posterior means of the two shocks are given by:

$$\begin{aligned}
E\left[\varepsilon^y_{t+h}|\mathcal{I}_{t+1}\right] &= \frac{\eta+\nu\kappa}{\kappa}E\left[a^y_{t+h}|a^{t+1},\varepsilon^{y,t+1},\varepsilon^{mp,t+1}\right] \\
&= \frac{\eta+\nu\kappa}{\kappa}E\left[a^y_{t+h}|a_{t+1}\right] \text{ since } a^y_{t+h} \text{ is i.i.d.} \\
&= \frac{\eta+\nu\kappa}{\kappa}\frac{Var\left(a^y_{t+h}|a_{t+1}\right)}{Var\left(a^y_{t+h}|a_{t+1}\right)+Var\left(a^{mp}_{t+h}|a_{t+1}\right)}a_{t+1} \\
&= \frac{\kappa\left(\eta+\nu\kappa\right)\sigma_y^2}{\kappa^2\sigma_y^2+\eta^2\sigma_{mp}^2}a_{t+1}.
\end{aligned}$$

Similarly,

$$\begin{aligned}
E\left[\varepsilon^{mp}_{t+h}|\mathcal{I}_{t+1}\right] &= \frac{\eta+\nu\kappa}{\eta}E\left[a^{mp}_{t+h}|a_{t+1}\right] \\
&= \frac{\eta+\nu\kappa}{\eta}\frac{Var\left(a^{mp}_{t+1}|a_{t+1}\right)}{Var\left(a^y_{t+1}|a_{t+1}\right)+Var\left(a^{mp}_{t+1}|a_{t+1}\right)}a_{t+1} \\
&= \frac{\eta\left(\eta+\nu\kappa\right)\sigma_{mp}^2}{\kappa^2\sigma_y^2+\eta^2\sigma_{mp}^2}a_{t+1}.
\end{aligned}$$

The posterior distribution of GDP growth is then given by:

$$\begin{aligned}
E\left[\Delta y^{us}_{t+h}|\mathcal{I}_{t+1}\right] &= \frac{E\left[\varepsilon^y_{t+h}|\mathcal{I}_{t+1}\right]-\nu E\left[\varepsilon^{mp}_{t+h}|\mathcal{I}_{t+1}\right]}{\eta+\nu\kappa} \\
&= Ka_{t+1}, \\
\text{where } K &= \frac{\kappa\frac{\sigma_y^2}{\sigma_{mp}^2}-\nu\eta}{\kappa^2\frac{\sigma_y^2}{\sigma_{mp}^2}+\eta^2}.
\end{aligned}$$

E.3 Second moments of Δs_{t+1}

$Cov\left(\Delta s_{t+1},\Delta y^{us}_{t+1}|\mathcal{I}_t\right)$ can be derived from the conditional covariance between US GDP growth and each component of the exchange rate change. That is,

$$Cov\left(\Delta s_{t+1},\Delta y^{us}_{t+1}|\mathcal{I}_t\right) = -Cov\left(\varphi^{EH}_{t+1},\Delta y^{us}_{t+1}|\mathcal{I}_t\right)-Cov\left(\sigma^F_{t+1},\Delta y^{us}_{t+1}|\mathcal{I}_t\right)+Cov\left(s^{\Delta E}_{t+1,\infty},\Delta y^{us}_{t+1}|\mathcal{I}_t\right).$$

Throughout the derivations below, we will use the fact that the information structure and the i.i.d. nature of shocks simplify the belief updating process from time t to $t+1$ to only updates in beliefs about $t+1$ variables (based on the observation of the shocks $\left\{\varepsilon^y_{t+1},\varepsilon^{mp}_{t+1}\right\}$) and $t+h$ variables (based on the observation of the announcement $a_{t+1}=i^{us}_{t+h}$). Beliefs about all other future observations are not updated.

For the nominal rate path term, we have:

$$\begin{aligned}Cov\left(\varphi^{EH}_{t+1},\Delta y^{us}_{t+1}|\mathcal{I}_t\right) &= Cov\left(\sum_{k=0}^{\infty}\left(E[i^{i}_{t+k+1}|\mathcal{I}_{t+1}]-E[i^{i}_{t+k+1}|\mathcal{I}_t]\right),\Delta y^{us}_{t+1}\middle|\,\mathcal{I}_t\right)\\ &\quad -Cov\left(\sum_{k=0}^{\infty}\left(E[i^{us}_{t+k+1}|\mathcal{I}_{t+1}]-E[i^{us}_{t+k+1}|\mathcal{I}_t]\right),\Delta y^{us}_{t+1}\middle|\,\mathcal{I}_t\right).\end{aligned}$$

Note that i^{us}_{t+1} is perfectly revealed by $a_{t-h+2}\in\mathcal{I}_t$, i^{us}_{t+h} is perfectly revealed by a_{t+1}, and $Cov\left(i^{us}_{t+h},\Delta y^{us}_{t+1}|\mathcal{I}_t\right)=0$ because shocks are i.i.d. Then,

$$\begin{aligned}&Cov\left(\sum_{k=0}^{\infty}\left(E[i^{us}_{t+k+1}|\mathcal{I}_{t+1}]-E[i^{us}_{t+k+1}|\mathcal{I}_t]\right),\Delta y^{us}_{t+1}\middle|\,\mathcal{I}_t\right)\\ =\;& Cov\left(i^{us}_{t+1}-E[i^{us}_{t+1}|\mathcal{I}_t]+E[i^{us}_{t+h}|\mathcal{I}_{t+1}]-E[i^{us}_{t+h}|\mathcal{I}_t],\Delta y^{us}_{t+1}\middle|\,\mathcal{I}_t\right)\\ =\;& 0,\end{aligned}$$

which implies that

$$Cov\left(\varphi^{EH}_{t+1},\Delta y^{us}_{t+1}|\mathcal{I}_t\right)=Cov\left(\sum_{k=0}^{\infty}\left(E[i^{i}_{t+k+1}|\mathcal{I}_{t+1}]-E[i^{i}_{t+k+1}|\mathcal{I}_t]\right),\Delta y^{us}_{t+1}\middle|\,\mathcal{I}_t\right).$$

For expectations of long-run exchange rate levels, we have:

$$\begin{aligned}Cov\left(s^{\Delta E}_{t+1,\infty},\Delta y^{us}_{t+1}|\mathcal{I}_t\right) &= Cov\left(\sum_{k=0}^{\infty}\left(E[\pi^{i}_{t+k+1}|\mathcal{I}_{t+1}]-E[\pi^{i}_{t+k+1}|\mathcal{I}_t]\right),\Delta y^{us}_{t+1}\middle|\,\mathcal{I}_t\right)\\ &\quad -Cov\left(\sum_{k=0}^{\infty}\left(E[\pi^{us}_{t+k+1}|\mathcal{I}_{t+1}]-E[\pi^{us}_{t+k+1}|\mathcal{I}_t]\right),\Delta y^{us}_{t+1}\middle|\,\mathcal{I}_t\right),\end{aligned}$$

where

$$\begin{aligned}&Cov\left(\sum_{k=0}^{\infty}\left(E[\pi^{us}_{t+k+1}|\mathcal{I}_{t+1}]-E[\pi^{us}_{t+k+1}|\mathcal{I}_t]\right),\Delta y^{us}_{t+1}\middle|\,\mathcal{I}_t\right)\\ =\;& Cov\left(\pi^{us}_{t+1}-E[\pi^{us}_{t+1}|\mathcal{I}_t]+E[\pi^{us}_{t+h}|\mathcal{I}_{t+1}]-E[\pi^{us}_{t+h}|\mathcal{I}_t],\Delta y^{us}_{t+1}\middle|\,\mathcal{I}_t\right)\\ =\;& \alpha Cov\left(\Delta y^{us}_{t+1}+E[\Delta y^{us}_{t+h}|\mathcal{I}_{t+1}],\Delta y^{us}_{t+1}\middle|\,\mathcal{I}_t\right)\\ =\;& \alpha Var\left(\Delta y^{us}_{t+1}\middle|\,\mathcal{I}_t\right)\\ =\;& \alpha Var\left(\frac{\varepsilon^{y}_{t+1}-\nu\varepsilon^{mp}_{t+1}}{\eta+\nu\kappa}\middle|\,\mathcal{I}_t\right)\\ =\;& \alpha\frac{\sigma^2_y+\nu^2\sigma^2_{mp}}{(\eta+\nu\kappa)^2}.\end{aligned}$$

Then, we have:

$$
\begin{aligned}
Cov\left(s^{\Delta E}_{t+1,\infty}, \Delta y^{us}_{t+1} | \mathcal{I}_t\right) &= Cov\left(\sum_{k=0}^{\infty}\left(E[\pi^i_{t+k+1}|\mathcal{I}_{t+1}] - E[\pi^i_{t+k+1}|\mathcal{I}_t]\right), \Delta y^{us}_{t+1} \middle| \mathcal{I}_t\right) \\
&\quad -\alpha\frac{\sigma_y^2 + \nu^2\sigma_{mp}^2}{(\eta+\nu\kappa)^2}.
\end{aligned}
$$

Lastly, for σ^F_{t+1}, we have:

$$
\begin{aligned}
\sigma^F_{t+1} &= \sum_{k=0}^{\infty}\left(E\left(\sigma_{t+k+1}|\mathcal{I}_{t+1}\right) - E\left(\sigma_{t+k+1}|\mathcal{I}_t\right)\right) \\
&= \gamma\lambda\sigma_{s,y}\sum_{k=0}^{\infty}\left(E\left[\bar{\rho}_{t+k+1}|\mathcal{I}_{t+1}\right] - E\left[\bar{\rho}_{t+k+1}|\mathcal{I}_t\right]\right).
\end{aligned}
$$

Since $\bar{\rho}_{t+1} = \theta\bar{\rho}_t - \lambda\Delta y^{us}_{t+1}$, this means that

$$
\begin{aligned}
\bar{\rho}_{t+k+1} &= \theta\bar{\rho}_{t+k} - \lambda\Delta y^{us}_{t+k+1} = \theta^3\bar{\rho}_{t+k-2} - \theta^2\lambda\Delta y^{us}_{t+k-1} - \theta\lambda\Delta y^{us}_{t+k} - \lambda\Delta y^{us}_{t+k+1} \\
&= \theta^{k+1}\bar{\rho}_t - \lambda\sum_{i=0}^{k}\theta^i\Delta y^{us}_{t+k+1-i},
\end{aligned}
$$

so that

$$
E\left[\bar{\rho}_{t+k+1}|\mathcal{I}_{t+1}\right] = \theta^{k+1}\bar{\rho}_t - \lambda\sum_{i=0}^{k}\theta^i E\left[\Delta y^{us}_{t+k+1-i}|\mathcal{I}_{t+1}\right],
$$

due to Δy^{us}_{t+i} being i.i.d.

The update in expectations of $\bar{\rho}_{t+k+1}$ between time t and $t+1$ is:

$$
\begin{aligned}
&E\left[\bar{\rho}_{t+k+1}|\mathcal{I}_{t+1}\right] - E\left[\bar{\rho}_{t+k+1}|\mathcal{I}_t\right] \\
= &\begin{cases} -\lambda\theta^k\left(\Delta y^{us}_{t+1} - E\left[\Delta y^{us}_{t+1}|\mathcal{I}_t\right]\right) & \text{if } k < h-1 \\ -\lambda\theta^k\left(\Delta y^{us}_{t+1} - E\left[\Delta y^{us}_{t+1}|\mathcal{I}_t\right]\right) - \lambda\theta^{k-(h-1)}E\left[\Delta y^{us}_{t+h}|\mathcal{I}_{t+1}\right] & \text{if } k \geq h-1, \end{cases}
\end{aligned} \tag{A-4}
$$

which implies

$$
\begin{aligned}
\sigma^F_{t+1} &= \gamma\lambda\sigma_{s,y}\sum_{k=0}^{\infty}\left(E\left[\bar{\rho}_{t+k+1}|\mathcal{I}_{t+1}\right] - E\left[\bar{\rho}_{t+k+1}|\mathcal{I}_t\right]\right) \\
&= -\gamma\lambda\sigma_{s,y}\sum_{k=0}^{\infty}\lambda\theta^k\left(\Delta y^{us}_{t+1} - E\left[\Delta y^{us}_{t+1}|\mathcal{I}_t\right]\right) - \gamma\lambda\sigma_{s,y}\sum_{k=h-1}^{\infty}\lambda\theta^{k-(h-1)}E\left[\Delta y^{us}_{t+h}|\mathcal{I}_{t+1}\right].
\end{aligned}
$$

Then,

$$
\begin{aligned}
Cov\left(\sigma_{t+1}^{F}, \Delta y_{t+1}^{us} | \mathcal{I}_t\right) &= \gamma\lambda\sigma_{s,y} Cov\left(\sum_{k=0}^{\infty}\left(E\left[\bar{\rho}_{t+k+1}|\mathcal{I}_{t+1}\right] - E\left[\bar{\rho}_{t+k+1}|\mathcal{I}_t\right]\right), \Delta y_{t+1}^{us}\middle| \mathcal{I}_t\right) \\
&= -\frac{\gamma\lambda^2\sigma_{s,y}}{1-\theta} Cov\left(\Delta y_{t+1}^{us} + E\left[\Delta y_{t+h}^{us}|\mathcal{I}_{t+1}\right], \Delta y_{t+1}^{us}\middle| \mathcal{I}_t\right) \\
&= -\frac{\gamma\lambda^2\sigma_{s,y}}{1-\theta} Var\left(\Delta y_{t+1}^{us}\middle| \mathcal{I}_t\right) \\
&= -\frac{\gamma\lambda^2\sigma_{s,y}}{1-\theta} \frac{\sigma_y^2 + \nu^2\sigma_{mp}^2}{(\eta+\nu\kappa)^2}.
\end{aligned}
$$

Putting together all three covariance terms gives us:

$$
\begin{aligned}
\sigma_{s,y} &\equiv Cov\left(\Delta s_{t+1}, \Delta y_{t+1}^{us}|\mathcal{I}_t\right) \\
&= -Cov\left(\varphi_{t+1}^{EH}, \Delta y_{t+1}^{us}|\mathcal{I}_t\right) - Cov\left(\sigma_{t+1}^{F}, \Delta y_{t+1}^{us}|\mathcal{I}_t\right) + Cov\left(s_{t+1,\infty}^{\Delta E}, \Delta y_{t+1}^{us}|\mathcal{I}_t\right) \\
&= Cov\left(\sum_{k=0}^{\infty}\left(E[\pi_{t+k+1}^{i}|\mathcal{I}_{t+1}] - E[\pi_{t+k+1}^{i}|\mathcal{I}_t]\right), \Delta y_{t+1}^{us}|\mathcal{I}_t\right) \\
&\quad -Cov\left(\sum_{k=0}^{\infty}\left(E[i_{t+k+1}^{i}|\mathcal{I}_{t+1}] - E[i_{t+k+1}^{i}|\mathcal{I}_t]\right), \Delta y_{t+1}^{us}|\mathcal{I}_t\right) \\
&\quad +\left(\frac{\gamma\lambda^2\sigma_{s,y}}{1-\theta} - \alpha\right) \frac{\sigma_y^2 + \nu^2\sigma_{mp}^2}{(\eta+\nu\kappa)^2}.
\end{aligned}
$$

Solving this for $\sigma_{s,y}$ gives:

$$
\begin{aligned}
\sigma_{s,y} = &-\frac{1}{1-\frac{\gamma\lambda^2}{1-\theta}\frac{\sigma_y^2+\nu^2\sigma_{mp}^2}{(\eta+\nu\kappa)^2}} Cov\left(\sum_{k=0}^{\infty}\left(E[i_{t+k+1}^{i} - \pi_{t+k+1}^{i}|\mathcal{I}_{t+1}] - E[i_{t+k+1}^{i} - \pi_{t+k+1}^{i}|\mathcal{I}_t]\right), \Delta y_{t+1}^{us}\middle| \mathcal{I}_t\right) \\
&-\frac{\alpha}{\frac{(\eta+\nu\kappa)^2}{\sigma_y^2+\nu^2\sigma_{mp}^2} - \frac{\gamma\lambda^2}{1-\theta}},
\end{aligned}
$$

which is constant as we have assumed that $\{i_{t+k}^{i}, \pi_{t+k}^{i}\}$ have constant second moments. This expression also makes clear under what conditions the dollar is a hedge for currency i ($\sigma_{s,y} < 0$) or not a hedge ($\sigma_{s,y} \geq 0$).

A similar approach can be used to show that $Var\left(\Delta s_{t+1}|\mathcal{I}_t\right)$ is constant under the same set of assumptions.

E.4 Model Prediction for Excess Return Component Response

We now relate the terms from our exchange rate decomposition to the GDP growth expectation.

First, note that, since $Cov\left(\Delta s_{t+1}, \Delta y^{us}_{t+1} | \mathcal{I}_t\right)$ and $Var\left(\Delta s_{t+1} | \mathcal{I}_t\right)$ are constant in our model, the expected excess returns term is as follows:

$$\begin{aligned}
\sigma^F_{t+1} &= \sum_{k=0}^{\infty} \left(E\left(\sigma_{t+k+1} | \mathcal{I}_{t+1}\right) - E\left(\sigma_{t+k+1} | \mathcal{I}_t\right)\right) \\
&= \gamma \lambda \sigma_{s,y} \sum_{k=0}^{\infty} \left(E\left[\bar{\rho}_{t+k+1} | \mathcal{I}_{t+1}\right] - E\left[\bar{\rho}_{t+k+1} | \mathcal{I}_t\right]\right).
\end{aligned}$$

Next, we relate the effect of the announcement on the update in expectations regarding $\bar{\rho}_{t+k+1}$ to the effect of the announcement on beliefs about GDP growth. From equation (A-4):

$$\frac{\partial}{\partial a_{t+1}} \left(E\left[\bar{\rho}_{t+k+1} | \mathcal{I}_{t+1}\right] - E\left[\bar{\rho}_{t+k+1} | \mathcal{I}_t\right]\right) = \begin{cases} 0 & \text{if } k < h-1 \\ -\lambda \theta^{k-(h-1)} \frac{\partial E\left[\Delta y^{us}_{t+h} | \mathcal{I}_{t+1}\right]}{\partial a_{t+1}} & \text{if } k \geq h-1 \end{cases}.$$

Then, we have:

$$\begin{aligned}
\frac{\partial \sigma^F_{t+1}}{\partial a_{t+1}} &= \gamma \lambda \sigma_{s,y} \sum_{k=0}^{\infty} \frac{\partial}{\partial a_{t+1}} \left(E\left[\bar{\rho}_{t+k+1} | \mathcal{I}_{t+1}\right] - E\left[\bar{\rho}_{t+k+1} | \mathcal{I}_t\right]\right) \\
&= -\gamma \lambda^2 \sigma_{s,y} \sum_{k=h-1}^{\infty} \theta^{k+(h-1)} \frac{\partial E\left[\Delta y^{us}_{t+h} | \mathcal{I}_{t+1}\right]}{\partial a_{t+1}} \\
&= -\frac{\gamma \lambda^2 \sigma_{s,y}}{1-\theta} \frac{\partial E\left[\Delta y^{us}_{t+h} | \mathcal{I}_{t+1}\right]}{\partial a_{t+1}} \\
&= -\frac{\gamma \lambda^2 \sigma_{s,y}}{1-\theta} K.
\end{aligned}$$

This derivative is always positively proportional to K for $\sigma_{s,y} < 0$ and negatively proportional to K for $\sigma_{s,y} > 0$.