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1 Introduction

Monetary policy—one of the most powerful tools used to influence the economy—affects real activity by moving asset prices, including exchange rates. The link between monetary policy and exchange rate movements is particularly relevant when it comes to the effects of US monetary policy on the value of the dollar, given its global dominance in trade invoicing, asset issuance, and official reserve holdings.[1] Therefore, understanding this link is of first-order importance.

A common perception among both financial market participants and policymakers is that lowering a country's interest rates, through either conventional or unconventional monetary policy, depreciates the domestic currency. Former Federal Reserve Chairman Ben Bernanke confirms that this belief was held by many policymakers, including himself, during the global financial crisis. Accommodative US unconventional monetary policy over this period triggered accusations that the Federal Reserve was engaging in "currency wars" by depreciating the dollar (see Bernanke 2017).

However, there is insufficient evidence of whether this widely held perception was true during the Great Recession. This paper examines this issue by systematically studying how conventional and unconventional monetary policy affects the nominal exchange rate. Moreover, we further disentangle the channels through which monetary policy shocks are transmitted to exchange rates and propose a theory to explain the surprising empirical facts that we document.

The first contribution of this paper is empirical. Our main result is that expansionary US monetary policy shocks during the Great Recession—dated from 2008:Q4 to 2012:Q2—in fact, caused the dollar to appreciate, on average, against a basket of currencies, contrary to the conventional wisdom. This result is in contrast to the response of exchange rates to US monetary policy that we document occurred before and after the Great Recession, which is in a direction consistent with the conventional wisdom. We further find important heterogeneity in the dollar's response to US monetary policy shocks. More precisely, a surprise US rate cut induced a larger appreciation of the dollar against currencies that tend to depreciate by more when US real output growth is low (i.e. "non-hedge" currencies from the perspective of the US investor). In contrast, it induced a depreciation of the dollar against currencies that tend to appreciate when US real output growth is low (i.e. "hedge" currencies).

[1] See Goldberg and Tille (2008), Shin (2012), Ivashina, Scharfstein, and Stein (2015), Casas et al. (2017), Gopinath (2016), and Gopinath and Stein (2018).

To disentangle the channels leading to this surprising set of results, we use a novel decomposition of the exchange rate response to US monetary policy. More specifically, based on an accounting identity, we decompose the dollar's response to US monetary policy surprises into components related to expected excess returns (currency risk premia), expected nominal returns, and expected inflation. What is novel is that we measure these components using a VAR approach that disciplines estimates of agents' expectations with survey forecasts (for details, see Stavrakeva and Tang 2019).

We show that US monetary policy easings caused the dollar to appreciate through two channels. The first channel is that 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, whose currencies are not a hedge against low US GDP growth. This led to a stronger dollar against non-hedge currencies, while the opposite was true for hedge currencies. This heterogeneity can account for the cross-currency heterogeneity in the response of exchange rates to US monetary policy over the Great Recession that we observe. The second channel that contributed to the dollar appreciation during the Great Recession is that expansionary US monetary policy also lowered the expected future path of US inflation relative to other countries. This second channel was present for both hedge and non-hedge currencies.

This paper's second main contribution is a model that shows how the signaling or information channel of monetary policy can reconcile all of the empirical facts that we document.[2] We present a stylized model in which the central bank is perceived to have better information about the future economic environment. Therefore, forward guidance that is intended to be accommodative can reveal expected adverse future shocks to US growth. The theory is similar in spirit to Tang (2015), Melosi (2017), Andrade et al. (Forthcoming), but our model additionally 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 forward guidance that promises lower future rates is overshadowed by the signaling effect of this forward guidance, then real GDP growth expectations will decline. This fall in expected growth heightens risk aversion. The higher risk aversion leads to lower expected excess return

[2]Our use of "signaling" is what Campbell et al. (2012) call "Delphic forward guidance" or what Nakamura and Steinsson (2018) call the "information effect." This terminology differs from the use of "signaling" in the quantitative easing (QE) literature to refer to QE actions conveying a commitment to maintain low future policy rates.

from being long US government debt and short foreign debt and generates a flight-to-safety-driven appreciation of the dollar against non-hedge currencies. 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 the dollar's response to US monetary policy shocks during the Great 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 as well as the lack of cross-currency heterogeneity in the inflation expectations channel that we empirically document occurred over the Great Recession.

For this theory to be consistent with our results, the signaling effect had to dominate the direct expansionary effect of promising lower rates during the Great Recession. 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. The model also implies that monetary policy easings over the crisis led to higher risk aversion, which we find is also strongly supported in the data.

To understand why the signaling effect overshadowed the direct effect of monetary policy during the Great Recession, we examine two sufficient conditions that are implied by the model. First, the existence of the signaling channel requires a type of monetary policy that has the potential to convey information about the central bank's forecast for future economic conditions. From 2008:Q4 to 2012:Q2, the Federal Reserve used "calendar-based" forward guidance promising low rates until some later date. Importantly, 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.[3] This type of forward guidance is ripe for interpretation as a sign of a deteriorating economy, thereby weakening expectations for growth.[4]

Second, the model shows that the signaling effect dominates when uncertainty

[3]Del Negro, Giannoni, and Patterson (2015) link the Fed's "Bad GDP language" during the post-crisis period to the information effect.

[4]In additional results available upon request, we attempt to disentangle the effect of forward guidance versus QE on the exchange rate using a higher-frequency approach 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.

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 Great 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. Combined with the calendar-based style of forward guidance, it becomes clear why monetary policy had a particularly strong signaling effect over the 2008:Q4 to 2012:Q2 period, resulting in a dollar appreciation, on average, in response to negative shocks to US rates.

Finally, the dollar appreciation during the Great Recession in response to monetary policy easings, which signaled a worsened economic outlook, is consistent with the literature on the "exorbitant duty" pioneered by Gourinchas, Rey, and Truempler (2012) and Gourinchas, Rey, and Govillot (2018). This literature documents that the dollar appreciates in bad times and the net asset valuations of foreign countries that invest in safe dollar-denominated bonds and borrow in risky foreign-denominated assets increase, resulting in a wealth transfer from the U.S. to the rest of the world. We indeed find that, in addition to a dollar appreciation, US policy easings during the Great Recession also led to a sizable deterioration in the U.S.'s net foreign asset position. As a result, not only did the Federal Reserve not engage in "competitive devaluation" through calendar-based forward guidance during the Great Recession, but it also contributed to a transfer of wealth from the U.S. to the rest of the world.

The paper proceeds as follows. The next subsection reviews the related literature. Section 2 details our empirical strategy. Section 3 presents our results for the effect of US monetary policy on the value of the dollar and its components, GDP growth expectations, risk aversion, and net wealth transfers from the United States to the rest of the world. Section 4 presents the model describing the signaling channel of monetary policy and argues empirically that the theoretical conditions for the signaling channel to dominate the direct channel of monetary were likely met. Section 5 concludes.

1.1 Related Literature

The paper is related to a number of different literatures.

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).[5] Another empirical monetary policy literature our paper contributes to is the literature that studies the signaling effect of monetary policy [see Campbell et al. (2012), Tang (2015), Nakamura and Steinsson (2018), Lunsford (2018), Miranda-Agrippino and Ricco (2018), and Jarociński and Karadi (Forthcoming)]. Consistent with these papers, we find that accommodative monetary policy with a strong signaling component is associated with downward revisions in GDP growth expectations.

Our paper is also related to the literature that examines the link between monetary policy and exchange rates. This literature has studied the effect of conventional monetary policy on exchange rates and finds a relationship consistent with the conventional wisdom (see, for example, Clarida and Gali 1994, Eichenbaum and Evans 1995, Kim and Roubini 2000 and Faust and Rogers 2003). Since the onset of the zero lower bound (ZLB), one other contemporaneous paper has systematically examined the link between unconventional monetary policy and exchange rates at a policy-relevant frequency.[6] Using a monthly VAR combined with high-frequency data to identify monetary policy shocks, Rogers, Scotti, and Wright (2018) study the response of the dollar to unconventional monetary policy easings over the ZLB. They find results consistent with the conventional wisdom, but these results come with an important caveat which is that they use an identification strategy with a sign-restriction that effectively purges monetary policy shocks of their signaling effect.

Our paper also contributes to the international finance literature which studies the exorbitant duty of the dollar to provide insurance to the rest of the world in bad times in exchange for allowing the U.S. to borrow at low rates during normal times [see Gourinchas, Rey, and Truempler (2012), Gourinchas, Rey, and Govillot (2018), Gourinchas and Rey (2007a;b; 2014), and Gourinchas, Rey, and Sauzet (Forthcoming)]. We are the first paper to document that monetary policy easings with a strong signaling com-

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

[6] Kiley (2013), Glick and Leduc (2015), and Swanson (2017) examine intra-day or daily exchange rate responses. Inoue and Rossi (2019) present impulse responses up to 15 days after a policy surprise.

ponent can deliver the negative economic news that trigger the US dollar's exorbitant duty behavior.[7] Our model's approach is also related to that of 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.[8]

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 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 rebalancing in which savings flow out of equity funds and into money market funds or government bond funds.[9]

2 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 decomposition of exchange rates changes that we will use to disentangle the transmission channels of US monetary policy surprises to exchange rates.

[7] Moreover, we add to the exorbitant duty narrative by documenting that the dollar appreciates not only through a flight-to-safety effect due to higher risk aversion, but also through inflation expectations being relatively lower in the U.S. along with lower growth expectations. We argue that the inflation channel of the dollar's appreciation will only be present for shocks coming from the U.S. Another novel contribution to this literature is the evidence on cross-currency heterogeneity in the exorbitant duty properties of the dollar, which can be explained by the heterogeneous hedging properties of currencies.

[8] Gourinchas, Rey, and Govillot (2018) model a shock that leads to a higher probability of a future rare disaster, while we model a monetary policy shock which is interpreted as a strong signal regarding future economic growth. Maggiori (2017), Farhi and Maggiori (2018) and He, Krishnamurthy, and Milbradt (2019) provide different microfoundations for the flight to safety and the exorbitant duty.

[9] They also find that the Japanese yen and the Swiss franc 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.

2.1 Data Description

We perform our analysis at a quarterly frequency over the 1991:Q1–2015:Q3 period, paying special attention to the Great Recession period—2008:Q4–2012:Q2—and to how responses to US monetary shocks differ during that time relative to the preceding and subsequent periods.[10] We focus primarily on explaining the effects of monetary policy during the Great Recession for a number of reasons. First, it is inherently important to understand the effects of monetary policy during times when it is most needed, such as during severe recessions. Second, and most importantly, we will document that monetary policy has effects over this period that are dramatically different relative to both predictions of standard theories and to other subperiods.[11]

We rely on the methodology developed in the high frequency monetary policy identification literature to identify monetary policy shocks (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 implied yields over a one-hour window ranging from 15 minutes prior to 45 minutes after FOMC statements and QE announcements that were made outside of these regular statements.[12]

Given that our sample of interest includes the ZLB period in the United States, we use prices of futures on underlying securities along the entire yield curve in order to capture unconventional policies that 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.[13] Since the payoffs of these futures contracts depend on the underlying short-term interest rates or Trea-

[10] Our Great Recession sample begins with the contraction in real economic activity that immediately followed the Lehman Brothers collapse and also includes the European debt crisis. The end date coincides both with the end of calendar-based forward guidance in the U.S. and 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.

[11] The fact that this period was special is confirmed by formal structural break tests detailed in the Section C of the Online Appendix.

[12] The list of QE announcements can be found in Online 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 announcements. This is important to note as we will argue later that the effects we find can be attributed to calendar-based forward guidance.

[13] 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.

sury yields that prevail upon settlement of the contracts, changes in these prices can be used to measure changes in interest rate expectations. Measuring these changes over very short time windows occurring around US monetary policy announcements gives changes in expectations that only reflect information about current and future policy actions conveyed by these announcements.[14]

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 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.[15] 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 (2018), and other related papers, we examine responses using forward rates at various horizons as policy indicators to more flexibly and agnostically capture the different dimensions of unconventional monetary policy. The exact empirical specifications are presented in the next subsection.

[14]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, which predominantly vary at lower frequencies. In other words, while we show later on that monetary policy shocks cause changes in risk premia within the same quarter, the changes occurring within a 30-minute window around these policy announcements are negligible.

[15]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.

2.2 Empirical Specification

For the responses of exchange rate changes and its components, we use relative one-year forward rates at various horizons as policy indicators. 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. $\Delta \tilde{f}_{t+1}^n$ represents the relative one-year forward rate n quarters ahead. Throughout the paper, tildes above variables denote 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. We focus on the medium and long ends of the yield curve because US short policy rates were constrained by the ZLB over our main period of interest, the Great Recession, and the policies in place during that time targeted lower yields for maturities of two or more years.

We use *relative* forward rates as policy indicators to estimate the exchange rate response 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. We utilize cross-sectional variation in how foreign forward rates respond to US policy shocks by allowing the first-stage regression coefficients to differ by currency.

In this specification, $\beta_n^{\Delta s_{t+1}}$ is the contemporaneous response of the exchange rate change to a US monetary policy shock that raises the n-quarter-ahead relative forward rate by 1 percentage point (see Stock and Watson 2018 for details). A positive $\beta_n^{\Delta s_{t+1}}$ then indicates that a policy easing led the dollar to appreciate.

For variables other than exchange rates, we follow the empirical monetary policy literature and use US interest rates as the policy indicators though 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:

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

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 Great Recession.

2.3 Exchange Rate Decomposition

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

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

For convenience, we will use the two terms, "expected excess currency return" and "currency risk premia," interchangeably though σ_t may also capture 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. Section 4 provides a 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:

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

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:

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

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^{EH}_{t+1}, (ii) the path of excess returns, σ^F_{t+1}, and (iii) long-run nominal exchange rates, $s^{\Delta E}_{t+1,\infty}$. Stavrakeva and Tang (2019) show that if the real exchange rate is stationary, $s^{\Delta E}_{t+1,\infty}$ reflects changes in expectations over long-run relative price levels, which equals the path of future relative inflation. Therefore, we will refer to $s^{\Delta E}_{t+1,\infty}$ as the inflation component of the exchange rate change decomposition.

We can then 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:

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

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 $\Delta \tilde{f}^n_{t+1}$, we can write $\hat{\beta}_n$ in terms of the following regression coefficients:

$$\hat{\beta}_n^{\Delta s_{t+1}} = \hat{\beta}_n^{\tilde{\imath}_t - \varphi^{EH}_{t+1}} + \hat{\beta}_n^{\sigma_t - \sigma^F_{t+1}} + \hat{\beta}_n^{s^{\Delta E}_{t+1,\infty}}, \tag{7}$$

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

Decomposing the response of exchange rate changes to policy shocks using equation (7) reveals the channels driving the differences in exchange rate responses to US monetary policy during the Great Recession compared to other periods.

For this decomposition, we use estimates of the terms in equation (6) from Stavrakeva and Tang (2019). As noted above, these terms feature expectations of short-term interest rates, exchange rates, and inflation. Since expectations for all future horizons are needed, the estimation in Stavrakeva and Tang (2019) relies upon a flexible vector autoregressive model (a VAR). The key to that estimation is that the VAR coefficients are chosen as those that produce the best match between VAR-implied and survey forecasts of exchange rates, interest rates, and inflation obtained from *Blue Chip* and *Consensus Economics*. Thus, the VAR is a structured way to interpolate and extrapolate survey forecasts to 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.[16]

3 Effects of US Monetary Policy During the Great Recession

3.1 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 Great Recession and the preceding and subsequent periods. We use Driscoll-Kraay standard errors which are robust to heteroskedasticity, cross-sectional correlation, and up to four lags of autocorrelation in the errors.[17] The estimates are also reported in the first row of Table 1.

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

During the Great Recession, 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. In contrast, the same US monetary policy shock has the opposite effect in the pre- and post-Global-Recession samples. The first-stage regression F-statistics for the Great Recession period are all greater than 39, far exceeding the rule-of-thumb threshold of 10 commonly used to detect the presence of weak instruments.

[16] 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 data. For more details on the VAR that we use, a discussion of the benefits of using survey data, and how well the VAR-implied expectations match survey forecasts, see Stavrakeva and Tang (2019).

[17] All results are robust to double clustering by currency pair and quarter.

Next, we examine the cross-currency heterogeneity in the response of the dollar to US policy surprises during the Great Recession. Heterogeneous responses are likely, given that the set 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 the hedging properties of each currency with respect to US real GDP growth.[18] The measure we use is motivated by the model that we present in Section 4. This hedging property is measured using the covariance between the respective exchange rate change and US real GDP growth.[19] A currency with a more negative covariance is one that tends to depreciate more *unconditionally* against the dollar when US real GDP growth falls and is not a hedge against low US growth. This figure shows a clear ordering, where the currencies that are a worse hedge also lost the most value against the dollar *conditional* on US monetary policy easings during the Great Recession.

Unsurprisingly, we find that the Swiss franc, the Japanese yen and the euro are hedge currencies for US growth, meaning that they appreciate unconditionally against the dollar when the US real economy is doing badly (the covariance is 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 only slightly against the euro.

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 currency with respect to real US GDP growth, captured by the covariance shown in Figure 2. We refer to these two groups as hedge and non-hedge currencies. The non-hedge currencies include AUD, CAD, NOK, NZD, SEK, and the GBP, while the hedge currencies are the EUR, JPY, and CHF.

Figures 3 and 4 plot the 2SLS regression coefficients from equation (1) for the non-hedge and 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 the non-hedge group, all of the coefficients for forward rate

[18]These results also hold for different $n \geq 12$.

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

horizons $n \geq 12$ are positive and statistically significantly different from zero during this subperiod. For the 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.

3.1.1 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). The 2SLS estimates of the regression coefficients in equation (7) are plotted 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-2 and A-3 and Tables A-2 and A-3 in the Online Appendix.[20]

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

The first is the currency risk premia term which gives the response coefficients $\hat{\beta}_n^{\sigma_t - \sigma_{t+1}^F}$.[21] These positive coefficients indicate that a US monetary policy easing that caused US forward rates to fall relative to foreign forward rates led to lower future expected excess returns from investing in dollar-denominated government debt and shorting debt denominated in currency i.

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 Great Recession, a US monetary policy easing that caused US forward rates to fall relative to foreign forward rates led to a lower expected inflation path in the United States relative to other countries. This is opposite to what would be expected if US policy easings have an expansionary effect on aggregate demand in the U.S.

Looking at the hedge currencies, Figure 6 shows a response of relative inflation expectations to US monetary policy shocks that is similar to that of the non-hedge

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

[21] 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.

currencies, indicating a lack of cross-currency heterogeneity in the strength of this inflation expectations channel. Though the currency risk premia responses are opposite in sign to those in the non-hedge sample, they tend not to be significantly different from zero (potentially due to this sample being smaller).

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 and pushes the dollar to depreciate for both the hedge and non-hedge panels. 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 that are a worse hedge, the expected excess return from being long the dollar and short the other currency fell by more in response to a US monetary policy easing (see Figure 7).

3.2 Why Did the Conventional Wisdom Not Hold Over the Great Recession?: The Signaling Effect of Monetary Policy and Risk Aversion

In this subsection, we attempt to further empirically disentangle what could have driven the surprising results that we find thus far. The hypothesis that we propose and test is that US monetary policy easings over the Great Recession led to lower (rather than higher) future US GDP growth expectations due to a strong signaling channel of monetary policy over this period. If agents expect future lower US GDP growth, and if the economy is driven primarily by demand shocks, they would also lower their expectations of future US inflation, which is consistent with the empirical results documented in the previous subsection. Moreover, based on the habit formation literature in asset pricing, we would expect that as a result of the lowered future GDP growth expectations, investors became more risk averse. This would trigger a flight to safety towards US safe assets and away from assets denominated in currencies that are not a hedge against low US GDP growth, thus explaining our finding that the currency risk premium component of exchange rates is the most important driver of the dollar appreciation against non-hedge currencies.

First, we test whether the signaling channel of US monetary policy was indeed sufficiently strong over the Great Recession. 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 the forecast quarter fixed. Table 4 presents the results. In response to US monetary policy surprises that lowered a US forward rate by 1 percentage point, GDP growth expectations were revised downwards by a statistically and economically significant amount that is between 0.7 and 1 percentage points, depending on the forward rate horizon. Therefore, we confirm that the signaling channel of US monetary policy dominated the direct expansionary effects of lower rates during the Great Recession.

Second, we test how US monetary policy easings over the crisis affected risk aversion. We consider the log 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 5, confirm that US policy easings during the Great Recession significantly increased investors' risk aversion. A fall of a US forward rate by 1 percentage point in response to US monetary policy led to an increase in risk aversion between 28 and 43 percent, depending on the forward rate horizon.

We conduct a second test of US monetary policy's effect 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 4. 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, $\sum_{k=1}^{\infty}\left(E[VIX_{t+k}|\mathcal{I}_{t+1}] - E[VIX_{t+k}|\mathcal{I}_t]\right)$.[22] The results, presented in Table 6, 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

[22]Since the VAR that is used in Stavrakeva and Tang (2019) to decompose exchange rates includes the VIX, 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.

horizon—during the Great Recession.

In conclusion, we find strong empirical support for the hypothesis that US monetary policy easings over the Great Recessions lowered future US GDP growth expectations and led to higher risk aversion.

3.3 Effects on Net Foreign Asset Positions And Valuation Effects

The data strongly supports the hypothesis that unconventional monetary policy did not result in "competitive devaluation" effects. In this subsection, we test whether there were actually wealth transfers from the U.S. to the rest of the world during the Great Recession as a result of US monetary policy easings. If this was indeed the case, it would imply that accommodative US monetary policy did not hurt, but rather potentially helped, other countries weather the crisis.

Such a hypothesis finds its roots in the exorbitant duty literature. According to that literature, the United States predominantly holds risky foreign-denominated assets, such as foreign equities, and borrows in safe dollar-denominated liabilities, such as US government debt. As a result, when the dollar appreciates, there is a mechanical negative valuation effect on US net foreign assets. Moreover, since, according to the exorbitant duty literature, the dollar appreciation is triggered by a flight-to-safety effect in bad times, the value of equities also falls while the value of safe dollar-denominated debt increases, which is the second component of the so-called valuation effect. Finally, an increase in 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 the US net foreign asset position.[23]

We now assess the effect of accommodative US monetary policy during the Great Recession on the US net foreign asset positions and valuation effects as constructed following Gourinchas, Rey, and Truempler (2012). Tables 7 and 8 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 Great Recession sample. Indeed, we find that a US monetary policy surprise that lowered medium- and long-horizon US forward rates by 1 percentage point led to statistically and economically significant decreases of US net foreign asset positions, with the largest decrease being 18 percent of US GDP for $n = 36$. The correspond-

[23] It is important to note that a rising dollar alone does not mechanically imply a negative valuation effect or a fall in the net US foreign asset position since the dollar appreciation could coincide with an increase in the market value of equities relative to government debt or with a positive flow effect.

ing 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.

3.4 Summary of Empirical Results

In summary, we find the novel and surprising result that the US dollar appreciated, on average, against a set of advanced economy currencies in response to US monetary policy easings during the Great Recession. Accommodative monetary policy over the period led to a downward revision of future US GDP growth expectations and an increase in risk aversion, which triggered flight-to-safety effects towards the dollar and away from non-hedge currencies. As a result, non-hedge currencies depreciated against the dollar while the opposite was true for hedge currencies, where the latter is a smaller subset of the sample of currencies we consider and, hence, the overall average effect. The second channel that contributed to a dollar appreciation during the Great Recession, against all currencies considered, is the lowered expectations of future US inflation relative to the inflation of the other countries in response to US monetary policy easings. Finally, we find that US monetary policy easings also resulted in a sizable wealth transfer from the United States to the rest of the world.

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

4 Theoretical Interpretation: Signaling Effect of Monetary Policy

In this section, we present a stylized model that can reconcile the empirical results presented in the previous sections. We employ a model of monetary policy signaling that is similar in spirit to Tang (2015), Melosi (2017) and Andrade et al. (Forthcoming). 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 Great 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.

4.1 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.[24] 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 nominal policy rate. Real GDP growth is 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:

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

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. (Forthcoming) 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.

The above three equations give the following solutions for Δy_t^{us}, π_t^{us} and i_t^{us}:

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

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 inflation and real GDP growth to result in a lower equilibrium nominal interest rate due to the endogenous policy reaction to these two variables.

Next, consider a representative agent in the United States who has time-varying risk aversion arising from consumption habits (see Campbell and Cochrane 1999, among

[24] This is a reasonable approximation for the United States as it is the least open advanced economy in the world with a gross-trade-to-GDP ratio of only 27 percent in 2017. The qualitative implications of this model can be generalized beyond this special case at the cost of some tractability.

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 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:

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

where S_{t+1} is the level of the nominal exchange rate defined as units of a foreign currency per dollar.[25] 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. Since we consider the limiting case of an approximately closed US economy, $c_t^{us} \approx y_t^{us}$. 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 $\lambda > 0$. This assumption for ρ_t implicitly imposes a functional form assumption on X_t^{us}, which is unobserved.[26]. We consider a 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 functional form substantially simplifies the analysis.

Given the assumptions made, Δy_{t+k}^{us} and $\bar{\rho}_{t+k}$ are normally distributed for any $k \geq 1$, conditional on $\mathcal{I}_t$ so long as $\mathcal{I}_t$ comprises of past shocks and normally distributed signals. We conjecture, and later confirm, that Δs_{t+1} is also normally distributed. As a result, equation (10) allows us to express the expected excess return from being long

[25] See the Online Appendix for details on the derivations.

[26] 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)

the dollar-denominated bond and short the foreign-currency-denominated bond as:

$$\begin{aligned}\sigma_t &= i_t^{us} - i_t^i + E[\Delta s_{t+1}|\mathcal{I}_t] \\ &= \frac{Var\left(\Delta s_{t+1}|\mathcal{I}_t\right)}{2} - Cov\left(\Delta s_{t+1}, -\gamma \Delta c_{t+1}^{us} + \gamma \Delta \rho_{t+1} - \pi_{t+1}^{us}|\mathcal{I}_t\right) \\ &= \frac{Var\left(\Delta s_{t+1}|\mathcal{I}_t\right)}{2} + \left(\gamma + \alpha + \gamma\lambda\bar{\rho}_t\right) Cov\left(\Delta s_{t+1}, \Delta y_{t+1}^{us}|\mathcal{I}_t\right), \end{aligned} \tag{12}$$

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 give conditions under which this is true in equilibrium. Then, the only source of time-variation in the expected excess return, σ_t, is time-varying risk aversion. From equation (12), a higher $\bar{\rho}_t$ lowers the expected excess return from holding the dollar between t and $t+1$ against non-hedge currencies for which $Cov\left(\Delta s_{t+1}, \Delta y_{t+1}^{us}|\mathcal{I}_t\right) < 0$. The opposite is true for hedge currencies for which $Cov\left(\Delta s_{t+1}, \Delta y_{t+1}^{us}|\mathcal{I}_t\right) > 0$.

The stylized nature of the model allows us to avoid taking a stand on whether markets are complete or incomplete and on the determinants of currencies' hedging properties. As a result, this model highlights the properties of the exchange rate and its components that are key for matching 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:

$$\begin{aligned}\Delta s_{t+1} =& i_t^i - i_t^{us} - \underbrace{\sum_{k=0}^{\infty}\left(E[i_{t+k+1}^i - i_{t+k+1}^{us}|\mathcal{I}_{t+1}] - E[i_{t+k+1}^i - i_{t+k+1}^{us}|\mathcal{I}_t]\right)}_{\varphi_{t+1}^{EH}} \\ &+ \sigma_t - \underbrace{\sum_{k=0}^{\infty}\left(E[\sigma_{t+k+1}|\mathcal{I}_{t+1}] - E[\sigma_{t+k+1}|\mathcal{I}_t]\right)}_{\sigma_{t+1}^{F}} \\ &+ \underbrace{\sum_{k=0}^{\infty}\left(E[\pi_{t+k+1}^i - \pi_{t+k+1}^{us}|\mathcal{I}_{t+1}] - E[\pi_{t+k+1}^i - \pi_{t+k+1}^{us}|\mathcal{I}_t]\right)}_{s_{t+1,\infty}^{\Delta E}}. \end{aligned} \tag{13}$$

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, $s^{\Delta E}_{t+1,\infty}$, equal the changes in expectations over the relative path of future inflation.

From the assumptions made on US demand and monetary policy shocks, the US variables and risk premium terms in equation (13) are all conditionally normally distributed with constant variances and constant covariances with Δy^{us}_{t+1}. Thus, additionally assuming that the nominal interest rate and inflation in the foreign country i, i^{i}_{t+k} and π^{i}_{t+k}, are conditionally normally distributed with constant second moments guarantees that the overall log change in the nominal exchange rate, Δs_{t+1}, will also be conditionally normally distributed with constant second moments.[27]

4.2 The Effect of Forward Guidance

To simplify the analysis, we assume that at time $t+1$, the central bank knows 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^{mp}_{t+h} and ε^{y}_{t+h} at time $t+1$.[28]

We consider a forward guidance announcement to be the central bank's truthful expectation of i^{us}_{t+h}, based on the interest rate 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 $-i^{us}_{t+h}$.[29] Then, our estimates of $\hat{\beta}^{\sigma_t - \sigma^F_{t+1}}_n$ and $\hat{\beta}^{s^{\Delta E}_{t+1,\infty}}_n$ in Section 3.1.1 correspond to the derivatives $-\frac{\partial\left(\sigma_t - \sigma^F_{t+1}\right)}{\partial a_{t+1}}$ and

[27] See the Online Appendix for details.

[28] To obtain our results on the responses to monetary policy announcements, 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^{mp}_{t+h} and ε^{y}_{t+h}. 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.

[29] 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.

$-\frac{\partial s^{\Delta E}_{t+1,\infty}}{\partial a_{t+1}}$ 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 Online Appendix, one can show that:

$$E\left[\Delta y_{t+h}|\mathcal{I}_{t+1}\right] = K a_{t+1}, \quad \text{where } K \equiv \frac{\kappa\frac{\sigma_y^2}{\sigma_{mp}^2} - \nu\eta}{\kappa^2\frac{\sigma_y^2}{\sigma_{mp}^2} + \eta^2}. \tag{14}$$

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.

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 $K > 0$. In Section 3.2 we confirmed that $K > 0$ over the Great 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, $-\frac{\partial\left(\sigma_t - \sigma^F_{t+1}\right)}{\partial a_{t+1}} = \frac{\partial \sigma^F_{t+1}}{\partial a_{t+1}}$. Then, one can show:

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

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 4 suggest was the case during the Great 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 currency i is not a hedge, i.e. if $Cov\left(\Delta s_{t+1}, \Delta y^{us}_{t+1}|\mathcal{I}_t\right) < 0$. The opposite is true if currency i is a hedge, 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 Great Recession, $\hat{\beta}_n^{\sigma_t - \sigma^F_{t+1}} > 0$ for our group of non-hedge currencies and $\hat{\beta}_n^{\sigma_t - \sigma^F_{t+1}} < 0$ for our group of 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 worse of a hedge currency i 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 shown in Subsection 3.1.

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,

$$
\begin{aligned}
-\frac{\partial s^{\Delta E}_{t+1,\infty}}{\partial a_{t+1}} &= \frac{\partial}{\partial a_{t+1}} \sum_{k=1}^{\infty} \left(E\left[\pi^{us}_{t+k} | \mathcal{I}_{t+1}\right] - E\left[\pi^{us}_{t+k} | \mathcal{I}_{t}\right] \right) \\
&= \frac{\partial}{\partial a_{t+1}} \left(E\left[\pi^{us}_{t+h} | \mathcal{I}_{t+1}\right] - E\left[\pi^{us}_{t+h} | \mathcal{I}_{t}\right] \right) = \alpha \underbrace{\frac{\partial E\left[\Delta y_{t+h} | \mathcal{I}_{t+1}\right]}{\partial a_{t+1}}}_{K}.
\end{aligned} \tag{15}
$$

Once again, understanding the effect that the forward guidance surprises have 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 US GDP growth is associated with lower US inflation. Thus, conditional on a strong signaling channel during the Great Recession, our model implies that $-\frac{\partial s^{\Delta E}_{t+1,\infty}}{\partial a_{t+1}} > 0$, which is consistent with the increase in expected future relative inflation (defined as foreign minus the U.S.) in response to US policy easings that we find in the data. Note also that the model implies no cross-currency heterogeneity in the response of this inflation component, as we find empirically.

An important observation coming out of the theory is that, during times of crisis, the inflation expectations 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 other 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 expectations 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 another country, as long as the shock increases risk aversion and as long as currency i is not a hedge for US growth. In summary, while the dollar's exorbitant duty behavior can be triggered by a shock that lowers expected future growth in either the U.S. or another country, the US shock will cause the dollar to appreciate through both the inflation expectations and flight-to-safety channels while a non-US shock will cause the dollar to appreciate only if the flight-to-safety effect dominates the inflation expectations effect.

To summarize, we showed that US monetary policy sent a sufficiently strong signal about economic conditions during the Great 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 Great Recession leading to downward revisions of US GDP growth forecasts and higher current and expected future risk aversion.

Our results raise the question of why the signaling channel was so much stronger than the direct channel of monetary policy during the Great 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 Great Recession can be seen in the average values of macroeconomic and monetary policy uncertainty measures before the Great 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 Great Recession subsample versus the other two subsamples. In contrast, the monetary policy subcomponent of the Baker, Bloom, and Davis (2016) policy uncertainty index actually declines slightly during the Great Recession and declines further still in the time period after 2012:Q2.[30] These results are consistent with $\frac{\sigma_y^2}{\sigma_{mp}^2}$ being particularly high during the Great Recession relative to the periods prior to and after the recession.

Second, during the Great Recession, as in the model, the Federal Reserve used calendar-based forward guidance that promised low rates at least until some future date—a policy that can easily be interpreted as a negative assessment of future US growth prospects by the Fed.[31] Policy actions and communication that leave room 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.

[30]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.

[31]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.

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 can impose exorbitant duty effects on the dollar and induces wealth transfers from the United States to the rest of the world.

5 Conclusion

In this paper, we revisit an old but important question—what is the effect of monetary policy on the nominal exchange rate? We study this effect, examining whether the relationship changes with the nature of monetary policy and further disentangling the transmission channels.

We find that, surprisingly, expansionary US monetary policy shocks during the Great Recession—dated from 2008:Q4 to 2012:Q2—caused the dollar to appreciate against a basket of advanced economy currencies, contrary to conventional wisdom. This is dramatically different from the effect of policy shocks in the periods before and after the crisis. 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 that an inflation expectations effect, whereby expansionary US monetary policy lowered the expected future path of US inflation relative to other countries, also increased the value of the dollar over the Great Recession.

These effects of accommodative US monetary policy during the Great Recession are surprising in the context of standard macroeconomic models which predict that, all else equal, a US policy easing should depreciate the dollar. We present a stylized model which illustrates that all of the empirical results that we document can arise when financial market participants interpret accommodative monetary policy as a signal of worsening future macroeconomic conditions, as we show was the case over the crisis.

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 re-evaluated, as existing exchange rate models do not consider the signaling effect of monetary policy. The accusations, informed by these models, that the U.S. was conducting "currency wars" during the Great Recession are shown to be unfounded. Instead, the negative economic news revealed by the Fed's forward guidance resulted in an appreciation of the dollar and a transfer of wealth from the U.S. to the rest of the world.

References

Andrade, P., G. Gaballo, E. Mengus, and B. Mojon. Forthcoming. "Forward Guidance and Heterogeneous Beliefs." *American Economic Journal: Macroeconomics*.

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

Baele, L., G. Bekaert, K. Inghelbrecht, and M. Wei. 2018. "Flights to Safety." SSRN Scholarly Paper ID 3204192. Rochester, NY: Social Science Research Network.

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

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

Beber, A., M. W. Brandt, and K. 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, G., E. Engstrom, and N. Xu. 2017. "The Time Variation in Risk Appetite and Uncertainty." SSRN Scholarly Paper ID 3069078. Rochester, NY: Social Science Research Network.

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

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

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

Campbell, J. R., C. L. Evans, J. D. M. Fisher, and A. Justiniano. 2012. "Macroeconomic Effects of FOMC Forward Guidance." *Brookings Papers on Economic Activity* 44(1): 1–80.

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

Casas, C., F. Diez, G. Gopinath, and P.-O. Gourinchas. 2017. "Dominant Currency Paradigm." Mimeo, Harvard.

Clarida, R., and J. Gali. 1994. "Sources of real exchange-rate fluctuations: How important are nominal shocks?" *Carnegie-Rochester Conference Series on Public Policy* 41: 1–56.

Crump, R. K., S. Eusepi, and E. Moench. 2016. "The Term Structure of Expectations and Bond Yields." Staff Report 775. New York: Federal Reserve Bank of New York.

Del Negro, M., M. Giannoni, and C. Patterson. 2015. "The Forward Guidance Puzzle." Working paper. Federal Reserve Bank of New York.

Eichenbaum, M., and C. 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.

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

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

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

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

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

Faust, J., and J. Rogers. 2003. "Monetary policy's role in exchange rate behavior." *Journal of Monetary Economics* 50(7): 1403–1424.

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

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

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

Glick, R., and S. Leduc. 2015. "Unconventional Monetary Policy and the Dollar: Conventional Signs, Unconventional Magnitudes." Working Paper Series 2015-18. San Francisco: Federal Reserve Bank of San Francisco.

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

Gopinath, G. 2016. "The International Price System." *Jackson Hole Symposium Proceedings*.

Gopinath, G., and J. C. Stein. 2018. “Banking, Trade, and the making of a Dominant Currency.” Working Paper 24485. Cambridge, MA: National Bureau of Economic Research.

Gourinchas, P.-O., and H. 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, P.-O., and H. Rey. 2007b. “International Financial Adjustment.” *Journal of Political Economy* 115(4): 665–703.

Gourinchas, P.-O., and H. 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, P.-O., H. Rey, and N. Govillot. 2018. “Exorbitant Privilege and Exorbitant Duty.” Mimeo.

Gourinchas, P.-O., H. Rey, and M. Sauzet. Forthcoming. “The International Monetary and Financial System.” *Annual Review of Economics*.

Gourinchas, P.-O., H. Rey, and K. Truempler. 2012. “The Financial Crisis and the Geography of Wealth Transfers.” *Journal of International Economics* 88(2): 266–283.

Greenwood, R., S. G. Hanson, and D. 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, R. S., B. Sack, and E. 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, R. S., B. Sack, and J. H. Wright. 2007. “The U.S. Treasury Yield Curve: 1961 to the Present.” *Journal of Monetary Economics* 54(8): 2291–2304.

He, Z., A. Krishnamurthy, and K. Milbradt. 2019. “A Model of Safe Asset Determination.” *American Economic Review* 109(4): 1230–1262.

Inoue, A., and B. Rossi. 2019. “The Effects of Conventional and Unconventional Monetary Policy on Exchange Rates.” *Journal of International Economics* 118: 419–447.

Ivashina, V., D. S. Scharfstein, and J. C. Stein. 2015. “Dollar Funding and the Lending Behavior of Global Banks.” *Quarterly Journal of Economics* 130(3): 1241–1281.

Jarociński, M., and P. Karadi. Forthcoming. "Deconstructing monetary policy surprises: the role of information shocks." *American Economic Journal: Macroeconomics*.

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

Kiley, M. 2013. "Exchange Rates, Monetary Policy Statements, and Uncovered Interest Parity: Before and After the Zero Lower Bound." Finance and Economics Discussion Series 2013-17. Washington, DC: Board of Governors of the Federal Reserve System.

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

Kim, D. H., and J. 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.

Kim, S., and N. Roubini. 2000. "Exchange Rate Anomalies in the Industrial a Solution With a Structural VAR Approach." *Journal of Monetary Economics* 45(3): 561–586.

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

Lunsford, K. G. 2018. "Understanding the Aspects of Federal Reserve Forward Guidance." Working Papers (Old Series) 1815. Federal Reserve Bank of Cleveland.

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

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

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

Miranda-Agrippino, S., and G. Ricco. 2018. "The Transmission of Monetary Policy Shocks." CEPR Discussion Papers 13396. C.E.P.R. Discussion Papers.

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

Perron, P., and Y. 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, M., J. Salomao, and M. Schneider. 2015. "Trend and Cycle in Bond Premia." Mimeo.

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

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

Rogers, J. H., C. Scotti, and J. H. Wright. 2018. "Unconventional Monetary Policy and International Risk Premia." *Journal of Money, Credit and Banking* 50(8): 1827–1850.

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

Stavrakeva, V., and J. Tang. 2019. "Survey-based Exchange Rate Decomposition: New Methodology and New Facts." Mimeo.

Stock, J. H., and M. 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, E. 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, J. 2015. "Uncertainty and the Signaling Channel of Monetary Policy." Working Paper 15-8. Boston: Federal Reserve Bank of Boston.

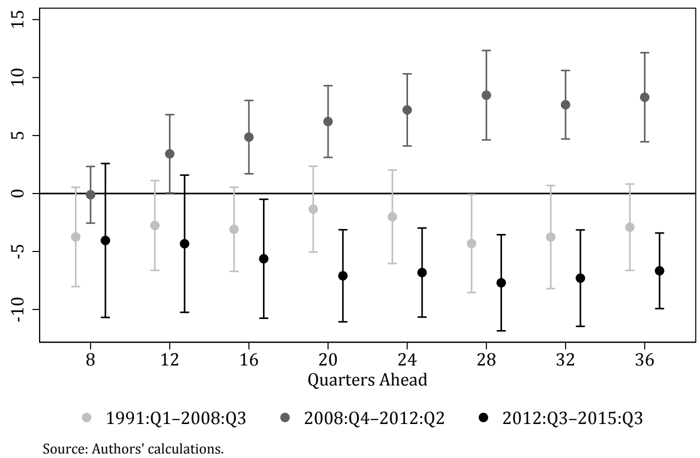
Wright, J. H. 2011. "Term Premia and Inflation Uncertainty: Empirical Evidence from an International Panel Dataset." *American Economic Review* 101(4): 1514–34.

Wright, J. 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, T. 2014. "Unconventional Monetary Policy and Long-Term Interest Rates." IMF Working Papers 14/189. Washington, DC: International Monetary Fund.

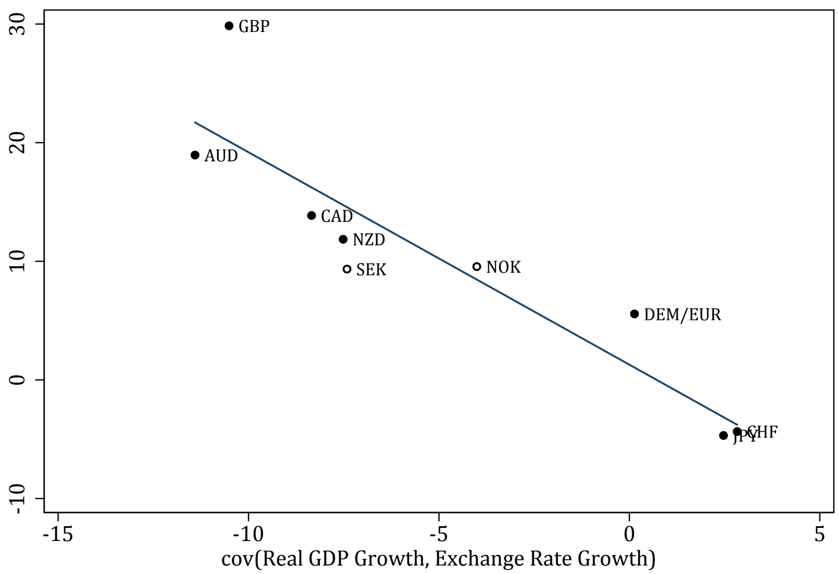
Figures and Tables

Figure 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.

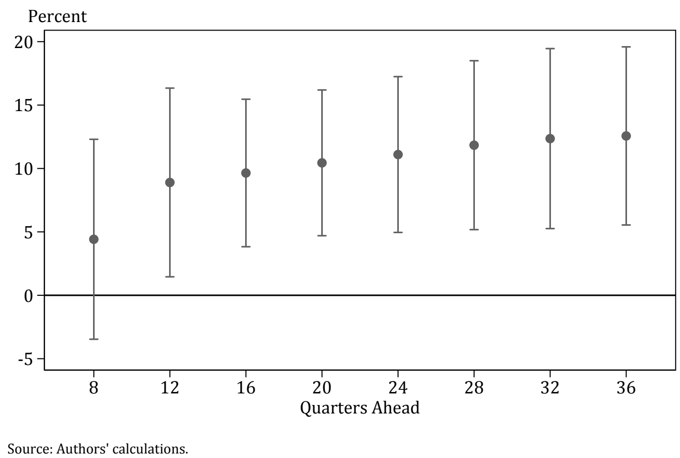
Figure 2: Pair-Specific Response of Exchange Rate Changes to US Monetary Policy Surprises versus Hedging Properties of the Dollar (2SLS)



Source: Authors' calculations.

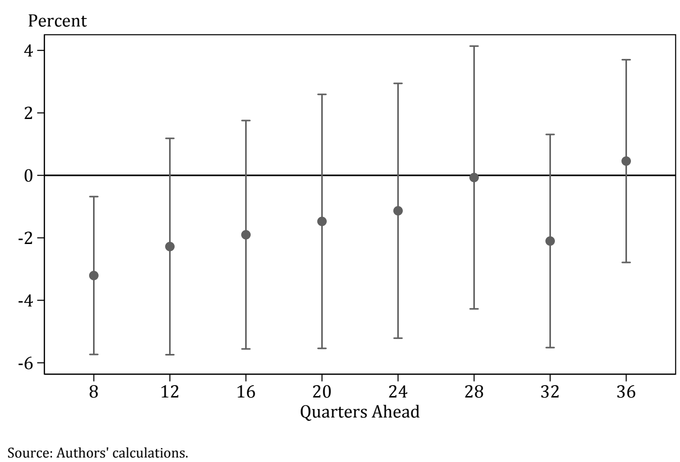
Note: Filled circles represent significance at 10 percent based on Drisoll-Kraay standard errors. 2SLS regression of exchange rate change on relative forward rate changes instrumented using yield changes in a one-hour window around FOMC announcements of FF4, ED4, and two- and 10-year Treasury note futures expiring in the current quarter. The measure of the hedging properties of a currency 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 Non-Hedge Currencies (2SLS)



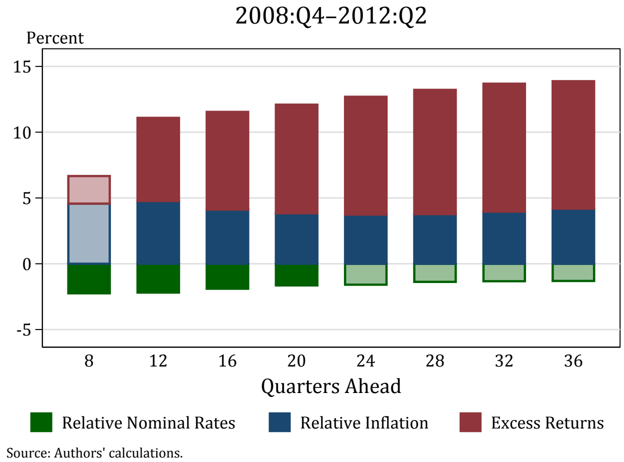
Note: 90 percent confidence intervals based on Driscoll-Kraay standard errors. 2SLS regression of exchange rate change on relative forward rate changes instrumented using yield changes in a one-hour window around FOMC announcements of FF4, ED4, and two- and 10-year Treasury note futures expiring in the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



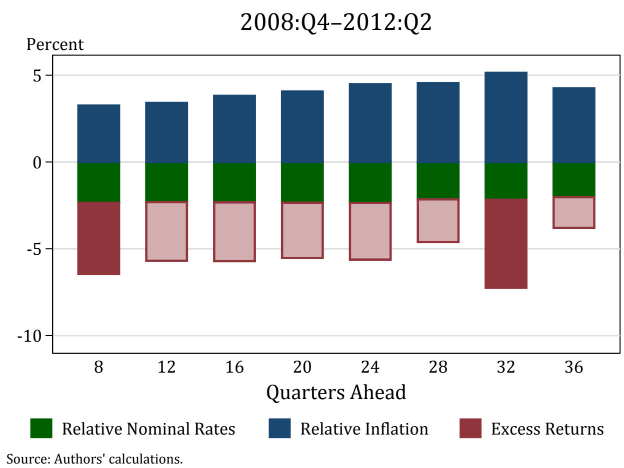
Note: 90 percent confidence intervals based on Driscoll-Kraay standard errors. 2SLS regression of exchange rate change on relative forward rate changes instrumented using yield changes in a one-hour window around FOMC announcements of FF4, ED4, and two- and 10-year Treasury note futures expiring in the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

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



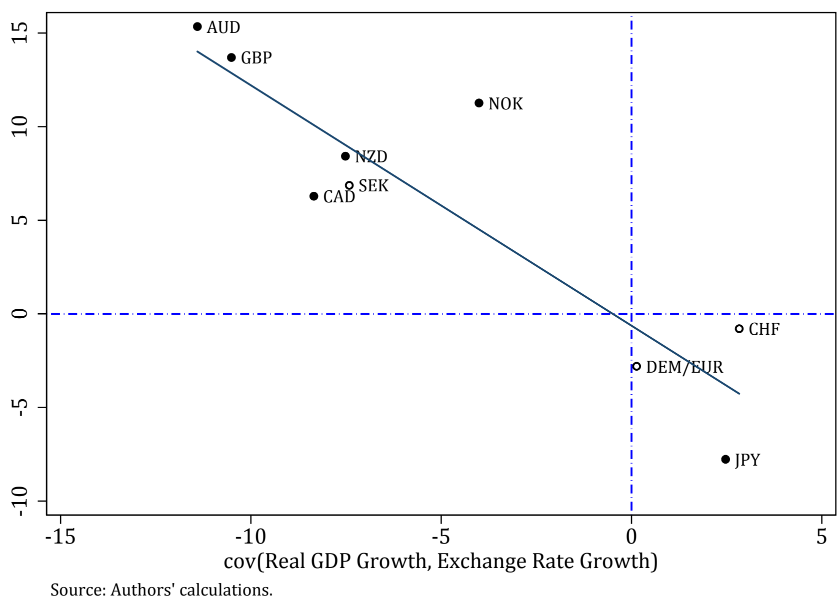
Note: Darker bar areas represent estimates significant at the 10 percent level based on Driscoll-Kraay standard errors. 2SLS regression of exchange rate change on relative forward rate changes instrumented using yield changes in a one-hour window around FOMC announcements of FF4, ED4, and two- and 10-year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



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 hedging properties of a currency 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)
$\tilde{\imath}_t - \varphi^{EH}_{t+1}$	-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)
$s^{\Delta E}_{t+1,\infty}$	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)
$\sigma_t - \sigma^F_{t+1}$	-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 ($\Delta\tilde{f}^n_{t+1}$) 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 Non-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)
$\tilde{\imath}_t - \varphi^{EH}_{t+1}$	-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)
$s^{\Delta E}_{t+1,\infty}$	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)
$\sigma_t - \sigma^F_{t+1}$	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 ($\Delta\tilde{f}^n_{t+1}$) 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 currencies of 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 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)
$\tilde{\imath}_t - \varphi^{EH}_{t+1}$	-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)
$s^{\Delta E}_{t+1,\infty}$	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)
$\sigma_t - \sigma^F_{t+1}$	-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 ($\Delta \tilde{f}^n_{t+1}$) 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 currencies of the euro area, Japan, and Switzerland against the US dollar.

Table 4: Response of US GDP Forecast Revisions to US Monetary Policy Surprises (2SLS)

Quarters Ahead	8	12	16	20	24	28	32	36
$\Delta f^{n,US}_{t+1}$	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^n_{t+1}) 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 Risk Aversion to US Monetary Policy Surprises (2SLS)

Quarters Ahead	8	12	16	20	24	28	32	36
$\Delta f^{n,US}_{t+1}$	-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^n_{t+1}) 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 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 7: 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 8: 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 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

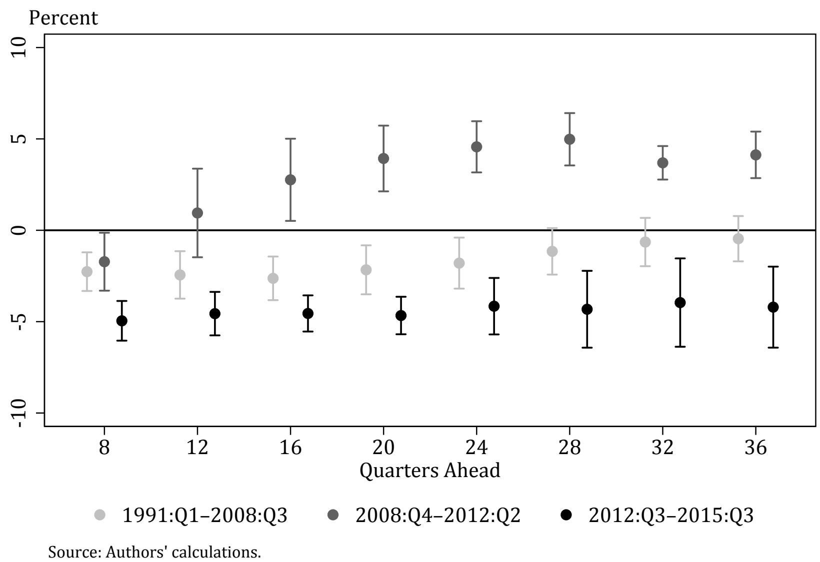
Note: The JLN macro uncertainty measure is 12-month ahead macroeconomic uncertainty estimated by Jurado, Ludvigson, and Ng (2015). GDP forecast dispersion is the 25th–75th percentile range of four-quarter-ahead US real GDP forecasts from *Blue Chip Financial Forecasts*. BBD monetary policy uncertainty 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

For Online Publication

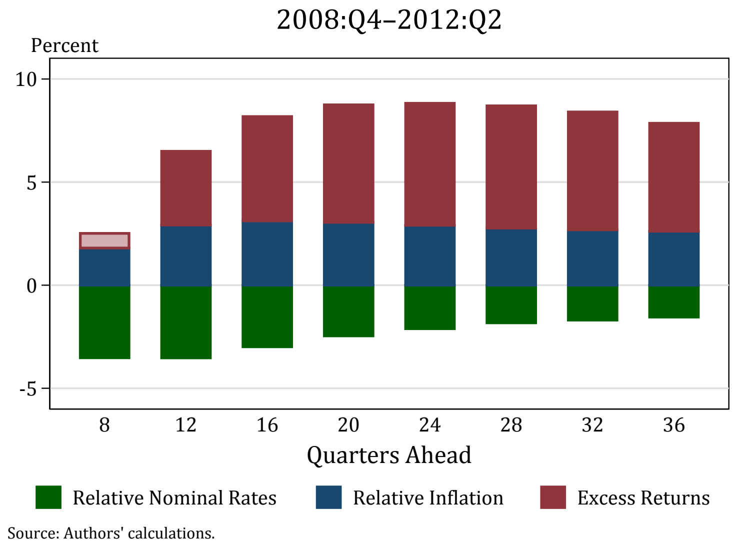
A Additional Figures and Tables

Figure A-1: Panel Regression of Exchange Rate Changes on Relative Forward Rate Changes for All Currencies (OLS)



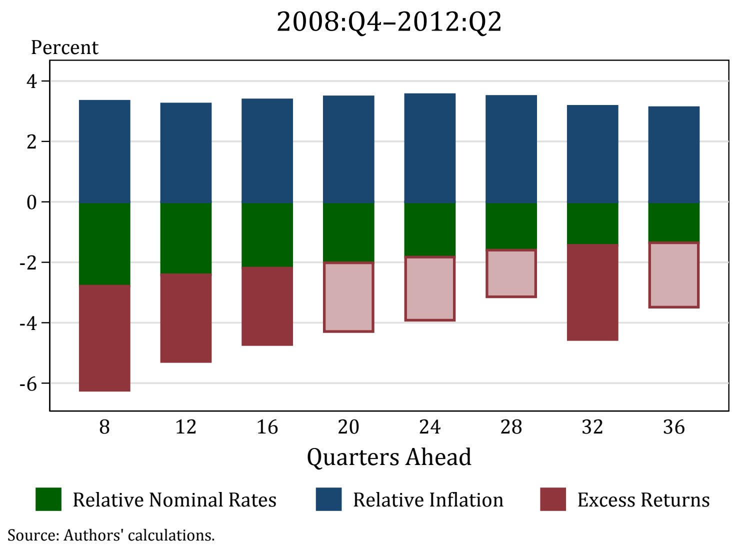
Note: 90 percent confidence intervals based on Driscoll-Kraay standard errors. OLS regression of exchange rate change on relative forward rate changes. This sample includes the currencies of Australia, Canada, euro area, Japan, Norway, New Zealand, Sweden, Switzerland, and the United Kingdom against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

Figure A-2: Panel Regression of Exchange Rate Changes on Relative Forward Rate Changes for 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 currencies of Australia, Canada, Norway, New Zealand, Sweden, and the United Kingdom against the US dollar.

Figure A-3: 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 the currencies of 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 Non-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 the currencies of 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 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 the currencies of the 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 Great 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.[32]

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 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).

[32]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.

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-4 plot the resulting SSRs as a ratio of the SSRs achieved by not allowing for a break in the estimated coefficients.

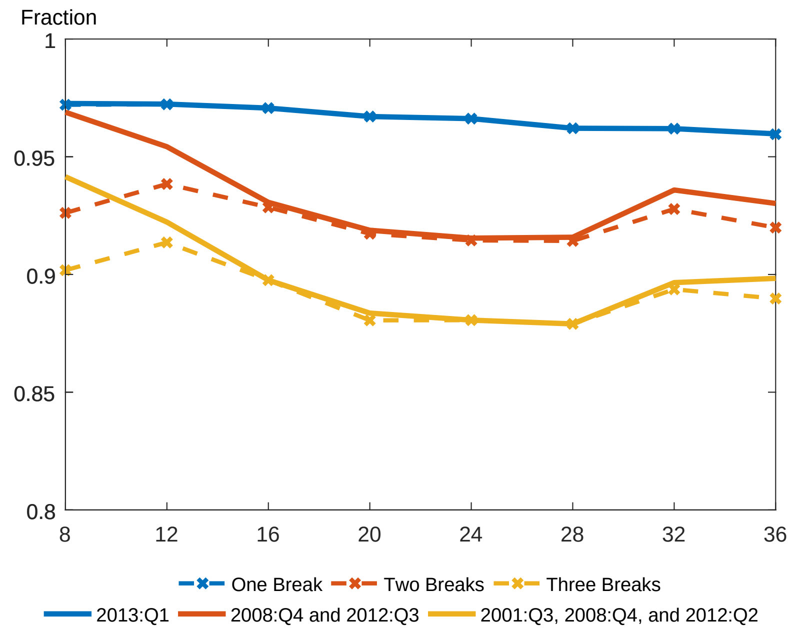
Table A-5: Break Dates that Minimize Sum of Squared Residuals

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-4

Figure A-4: 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.

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 optimal horizon-specific breaks shown in Table A-5. Figure A-4 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 Model: Additional Derivations

D.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:

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

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

$$\begin{aligned}\frac{\mu_{t+1}}{\mu_t} &= \frac{u_c\left(C_{t+1}^{us}, X_{t+1}^{us}\right)}{u_c\left(C_t^{us}, X_t^{us}\right)} \frac{P_t^{us}}{P_{t+1}^{us}} \\ &= \frac{u_c\left(C_{t+1}^{us}, X_{t+1}^{us}\right)}{u_c\left(C_t^{us}, X_t^{us}\right)} e^{-\pi_{t+1}^{us}}. \end{aligned} \tag{A-2}$$

Combining equations A-1 and A-2 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: $\pi_{t+k}^i = \Delta \tilde{p}_{t+k}^{us,i}$.

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

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

Similarly,

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

The posterior distribution of GDP growth is then given by:

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

D.3 Second moments of Δs_{t+1}

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

$$
Cov\left(\Delta s_{t+1}, \Delta y_{t+1}^{us} | \mathcal{I}_t\right) = -Cov\left(\varphi_{t+1}^{EH}, \Delta y_{t+1}^{us} | \mathcal{I}_t\right) - Cov\left(\sigma_{t+1}^{F}, \Delta y_{t+1}^{us} | \mathcal{I}_t\right) + Cov\left(s_{t+1,\infty}^{\Delta E}, \Delta y_{t+1}^{us} | \mathcal{I}_t\right).
$$

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 $\left\{\varepsilon_{t+1}^{y}, \varepsilon_{t+1}^{mp}\right\}$) and $t+h$ variables (based on the observation of the announcement $a_{t+1} = i_{t+h}^{us}$). Beliefs about all other future observations are not updated.

For the nominal rate path term, we have:

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

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

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

which implies that

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

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

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

where

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

Then, we have:

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

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:

$$\begin{aligned}&E\left[\bar{\rho}_{t+k+1}|\mathcal{I}_{t+1}\right]-E\left[\bar{\rho}_{t+k+1}|\mathcal{I}_t\right]\\ =&\begin{cases}-\lambda\theta^k\left(\Delta y^{us}_{t+1}-E\left[\Delta y^{us}_{t+1}|\mathcal{I}_t\right]\right) & \text{if } k<h-1\\ -\lambda\theta^k\left(\Delta y^{us}_{t+1}-E\left[\Delta y^{us}_{t+1}|\mathcal{I}_t\right]\right)-\lambda\theta^{k-(h-1)}E\left[\Delta y^{us}_{t+h}|\mathcal{I}_{t+1}\right] & \text{if } k\geq h-1,\end{cases}\end{aligned}\tag{A-3}$$

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,

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

Putting together all three covariance terms gives us:

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

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 currency is a non-hedge currency i ($\sigma_{s,y} < 0$) or a hedge currency ($\sigma_{s,y} \geq 0$).

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.

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

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

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-3):

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

Then, we have:

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

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

K for $\sigma_{s,y} > 0$.