Exchange Rates and Monetary Policy

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Abstract

In this paper we confront the data with the financial-market folk wisdom that monetary policy is one of the key drivers of nominal exchange rates. Focusing on measures of conventional and unconventional monetary policy, we find that monetary policy surprises and changes in expectations about future monetary policy can explain a sizable fraction of the variation in exchange rate changes for certain currency pairs. However, our results show that expected excess returns account for most of this variation. We also find that the importance unconventional monetary policy plays for explaining exchange rate changes is larger in the period since the United States hit the zero lower bound in December 2008. In contrast, the importance of conventional monetary policy is lower during this period due to a decrease in the volatility of monetary policy surprises. Meanwhile, the marginal response of exchange rate changes relative to conventional policy surprises actually has strengthened due to a change in the relationship between these surprises and expected excess returns.

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1 Introduction

Monetary policy is an important driver of asset prices—one category being exchange rates. Therefore, it is not surprising that financial markets focus close attention on the link between contemporaneous exchange rate dynamics and changes in short-term policy rates or long-term yields, since long-term yields can contain information about expected future policy rates.

In financial markets, the folk wisdom holds that a country's currency appreciates when its interest rate increases relative to that of other countries. Evidence of this relationship is clearly illustrated in Figure 1 which plots exchange rate levels against both relative short- and long-term yields for the Australian dollar, British pound, Deutschmark/euro, Japanese yen, and Swiss franc against the U.S. dollar. In this graph, an increase in the respective exchange rate represents a depreciation of the country's currency versus the U.S. dollar, while an increase in the relative yield implies an increase in that country's yield relative to the U.S. yield.

Besides financial market participants, central banks have also attempted to exploit this relationship between monetary policy and currency movements in order to manipulate exchange rates and stimulate foreign demand-driven domestic growth. In this context, much emphasis has been placed on conducting conventional monetary policy (that is, manipulating short-term rates). For example, a *Wall Street Journal* report on the Reserve Bank of Australia's February 3, 2015 decision to cut its benchmark interest rate noted that:

> *Nations cutting interest rates or engaging in stimulus measures are hoping to boost domestic growth and employment by weakening their currencies, making exported goods more competitive overseas. Still, currency depreciation also raises the risk of a tit-for-tat race to the bottom, as trading partners seek to outdo one another only to find gains are limited.*[1]

Unlike Australia, which remained relatively unscathed by the global financial crisis, many central bankers of other advanced economies have their hands tied in recent years due to the binding zero lower bound (ZLB) and have turned to unconventional monetary policy tools.[2] The financial press has also made references to the use of unconventional monetary policy for the purpose of devaluing a country's currency as evidenced by the following:

[1] See James Glynn, "Australia Cuts Interest Rates to Record Low; Reserve Bank Eases Policy in Response to Tumbling Oil Price, Deflation Worries," *The Wall Street Journal Online*, February 3, 2015. Available at http://www.wsj.com/article_email/australia-cuts-interest-rates-1422935313-lMyQjAxMTE1MjAwOTIwNDkzWj.

[2] The Federal Reserve effectively hit the zero lower bound in 2008:Q4 while the Bank of England followed soon after in 2009:Q1. The European Central Bank kept their policy rate above zero for a longer period, but began taking unconventional monetary policy actions in 2008. Due to a prolonged recession in the 1990s, Japan has been at the zero lower bound since 1999 and began quantitative easing operations in 2001.

Just as a weaker currency channel was an important part of the Fed's quantitative easing programme, so it is part of the playbook of the European Central Bank for its bond-purchasing drive, though the ECB will not say so publicly.[3]

In this paper, we systematically explore the empirical link between monetary policy and exchange rate fluctuations for ten developed economies vis-à-vis four base currencies: the U.S. dollar, the British pound, the Deutschemark/euro and the Japanese yen. We do this by performing a decomposition of quarterly exchange rate changes that uses two main building blocks: (1) a standard no-arbitrage asset pricing equation[4] and (2) forecasting models of policy rates and exchange rate changes. An advantage of this approach is that we impose few theoretical restrictions; in particular, we do not need to take a stand on whether financial markets are complete or incomplete. For analytical tractability, the motivating theoretical derivations only assume log-normality for the stochastic discount factor, yields, and exchange rate changes.

Our main assumptions center on the estimation of policy rate expectations. We focus on two different forecasting models: a Taylor rule with policy rate smoothing, commonly used in the macro-empirical literature, and a yield factor model of policy rates, which is more widely used in the asset pricing literature. While the Taylor rule is the workhorse equation describing policy rates in macroeconomic models, we consider the yield factor model to be better at describing financial market expectations in a period characterized by binding zero lower bound constraints and unconventional monetary policy actions. Embedding these interest rate forecasting equations into VARs allows us to estimate expectations at all future horizons of policy and exchange rates.

There are two main reasons why we chose to use these two approaches to modeling policy rate expectations rather than using an unrestricted VAR specification with a wider set of variables. First, we compare interest rate surprises from our two specifications with market-based surprises derived from futures data and find correlations as high as 0.84. Second, a literature exists that warns against VAR overidentification and recommends imposing theory-relevant restrictions or priors in VAR estimation (see Giannone, Lenza, and Primiceri (2014) and Rudebusch (1998) among others).

Combining these policy rate expectations with standard no-arbitrage asset pricing conditions allows us to decompose the exchange rate change between periods t and $t+1$ into the period t

[3] Roger Blitz and Claire Jones, "Merkel's 'Strong Euro' Comments Put Currency Investors on Alert," *Financial Times, June 15, 2015*. Available at http://www.ft.com/intl/cms/s/0/8d3ec190-1121-11e5-9bf8-00144feabdc0.html#axzz3r2hkYvOy.

[4] The core theoretical relationship that we explore in this paper — linking exchange rate appreciation and monetary policy tightening — goes at least as far back as the Dornbusch overshooting model and is also present in the New Keynesian model. (See Frankel (1979) and the literature review in Engel (2014)). In those models, as in our framework, the relationship comes naturally out of a no-arbitrage asset pricing condition. This is quite similar to the Euler equation on safe debt in standard macroeconomic models; however, it does not link stochastic discount factors to preferences or consumption.

interest rate differential and expected excess return plus a period $t+1$ expectational error. The expectational error can be further divided into the following four components: (i) the unanticipated component of period $t+1$ relative policy rates (which we refer to as a "monetary policy surprise"[5]), (ii) the sum of changes in expectations of relative policy rates from period $t+2$ onwards, (iii) the sum of changes in expectations of excess returns from period $t+1$ onwards, and (iv) changes in expectations of the long-run nominal exchange rate level. Other papers that consider a similar decomposition of the exchange rate change include Froot and Ramadorai (2005), Engel and West (2005; 2006), Engel, Mark, and West (2008), Mark (2009), Engel and West (2010), and Engel (2014; Forthcoming).

Our paper focuses on the monetary policy terms—terms (i) and (ii)—which capture the *direct* link between exchange rate changes and contemporaneous unanticipated monetary policy or changes in expectations of the future paths of policy. We further separate out a component of changes in expected future policy rates that is orthogonal to contemporaneous policy surprises. This distinction is conceptually similar to the approach of Gürkaynak, Sack, and Swanson (2005), although the method and underlying data differ. This exercise separates the effect of interest rate persistence from other factors that might affect future rate expectations such as forward guidance and the signaling channel of quantitative easing (QE).

Equipped with this decomposition of the exchange rate change, we next perform a Campbell and Shiller (1988a;b) variance-covariance decomposition to assess the fraction of variation in exchange rate changes that can be explained by variation in the monetary policy terms versus other terms in the decomposition. We find that monetary policy can explain a sizable fraction of the volatility of exchange rate changes for some currency pairs, primarily the ones where the GBP is the base currency. For example, for the GBP-USD, GBP-DEM/EUR and GBP-CHF currency pairs, the fraction of the exchange rate variation explained by monetary policy equals 0.38 (0.31), 0.23 (0.61) and 0.21 (0.29) for the Taylor rule (yield factor) specification, respectively. However, we also find that the expected excess return components account for most of the exchange rate change fluctuations, while changes in expectations over long-run nominal exchange rates and uncovered interest rate parity both play negligible roles in these movements. Engel and West (2010) also perform a similar exercise using a different VAR specification and examine a smaller set of currency pairs, a different time period, and real rather than nominal exchange rates. Froot and Ramadorai (2005) also perform a similar decomposition, but focus on explaining exchange rate movements at a daily frequency using order flows. In contrast to both these papers, we rely on information contained in long-term yields, which is particularly important over the period when central banks were constrained by the zero lower bound.

[5] The use of the term "surprise" here should not be construed as a structural shock. We use this term merely as a more concise way of referring to the unanticipated component of a particular variable.

We then show that the overall importance of monetary policy in explaining exchange rate change volatility decreased in the period after the United States hit the zero lower bound in 2008:Q4. This occurred despite the fact that the correlation between exchange rate changes and relative monetary policy surprises became stronger. A lower variance of relative monetary policy surprises over this period helps us reconcile these two facts. Additionally, we find that our measure of unconventional monetary policy has played a greater role in explaining the volatility of the exchange rate change over the recent period. This finding is consistent with the recent reliance on forward guidance and QE by central bankers.

Another way to assess the relationship between monetary policy and exchange rates is to examine the coefficient obtained by regressing the exchange rate change on our measure of relative monetary policy surprises. This coefficient is similar in spirit to the impact effect of a monetary policy shock on exchange rates that is obtained in impulse responses estimated by the VAR literature (see Eichenbaum and Evans (1995), Rogers (1999), Faust and Rogers (2003), and Scholl and Uhlig (2008)). However, an important distinction is that this literature aims to identify structural monetary policy shocks, something we do not attempt.

While our results are consistent with the findings in the VAR literature, we go a step further and decompose the estimated responses of exchange rate changes to relative monetary policy surprises into the responses of terms in our decomposition of the exchange rate change. This exercise reveals that while the expected excess return components dampened the response of the exchange rate change to a relative monetary policy surprise in the pre-ZLB period, it has since had an amplifying effect. The overall result is that exchange rate changes have responded more to monetary policy surprises in the ZLB period. Taking the USD-DEM/EUR pair as an example, for a 1 percentage point larger monetary policy surprise (on a nonannualized basis) in the eurozone relative to the United States, the euro would have appreciated by a statistically insignificant 3 percent in the pre-ZLB period. In contrast, during the ZLB period, the same situation yields a statistically significant euro appreciation of 43.5 percent.

In addition to the studies referenced above, this paper is related to a number of other literatures. This includes work studying the effect of monetary policy news on exchange rates using high-frequency data, such as Andersen et al. (2002), Faust et al. (2007), and Clarida and Waldman (2008). These papers find that a sizable fraction of the variation in changes in exchange rates during tight windows around the time of policy and macroeconomic announcements can be explained by shocks to monetary policy and macroeconomic fundamentals. Our paper offers a clear theoretical exposition of the link between these shocks and exchange rate changes using minimal assumptions. Furthermore, unlike these other papers using high frequency data, this paper uses quarterly data and addresses a different set of questions.

In the finance literature, our work is closely related to a number of papers that jointly study exchange rates and bond yields. Ang and Chen (2010) and Chen and Tsang (2013) both estimate forecasting equations either of exchange rate changes or excess returns regressed on yield curve variables. The key distinction is that these papers focus on the ability of *lagged* yield curve variables to predict future exchange rate changes or returns, while we examine whether *contemporaneous changes* in yields comove with exchange rate changes in a manner consistent with economic theory.

Another set of related papers links exchange rates and bond yields by structurally estimating two-country term structure models. Examples include Bansal (1997), Backus, Foresi, and Telmer (2001), Ahn (2004), Brennan and Xia (2006), Wu (2007), Sarno, Schneider, and Wagner (2012), and Yung (2014). Our paper makes two main departures from this literature—we impose fewer restrictions in our estimation by not assuming complete markets and we do not assume a particular functional form for the stochastic discount factor. Thus, our approach may be less subject to misspecification. Furthermore, we decompose exchange rate changes in a way that allows us to connect these fluctuations to macroeconomic concepts such as monetary policy surprises and expectations of future interest rate changes.

Lastly, our paper is linked to parts of the exchange rate forecasting literature pioneered by Meese and Rogoff (1983a) and Meese and Rogoff (1983b). Papers studying the forecasting power of Taylor rules include Engel, Mark, and West (2008) and Molodtsova and Papell (2009). Given that we build on a no-arbitrage condition, our paper is linked to the vast literature studying the currency risk premium and the uncovered interest rate parity (UIRP) condition. These variables show up in our decomposition of the exchange rate change, but are not the main focus of this paper. In a recent handbook chapter, Engel (2014) provides a comprehensive review of the UIRP and exchange rate literatures.

Section 2 sets out the model's building blocks and presents our exchange rate change decomposition. Furthermore, it presents results from reduced-form regressions, which confirm the correlations implied by Figure 1. Section 3 presents our VAR methodology and Section 4 contains our empirical results. Section 5 concludes.

2 Model

Consider a trader or financial institution that can invest in any asset denominated in any currency. We will use no-arbitrage conditions seen from the trader's viewpoint to derive our expressions for exchange rate dynamics. We focus on risk-free sovereign bonds of different maturities denominated in the local currency. In both the model and our data, one time period is a quarter. Use $S_t^{i,j}$ to denote the exchange rate of currency i per one unit of currency j. This exchange rate is defined such that an increase of $S_t^{i,j}$ corresponds to a depreciation of currency i against currency j. The one-period return in units of currency i on an n-period bond denominated in currency i is

$$R_{t+1}^{n,i} = \frac{P_{t+1}^{n-1,i}}{P_t^{n,i}},$$

where $P_t^{0,i} = 1$. The log yield-to-maturity of this same bond is defined as

$$y_t^{n,i} = -\frac{1}{n} \ln \left(P_t^{n,i}\right).$$

The yield on a one-period ("short-term") bond, $i_t^i \equiv y_t^{1,i} = -\ln\left(P_t^{1,i}\right)$, will serve as our proxy of the monetary policy rate targeted by the central bank of country i.[6] As is common in the asset pricing literature, we start from no-arbitrage conditions. We assume that the trader's nominal stochastic discount factor (SDF), $M_{t+1}^{\$}$, is denominated in U.S. dollars and we do not take a stand on whether markets are complete or incomplete.[7] Then, the no-arbitrage condition that guarantees the trader's indifference between investing in n-period bonds denominated in U.S. dollars, currency i, or currency j is

$$1 = E_t\left[M_{t+1}^{\$} R_{t+1}^{n,\$}\right] = E_t\left[M_{t+1}^{\$} R_{t+1}^{n,i} \frac{S_t^{i,\$}}{S_{t+1}^{i,\$}}\right] = E_t\left[M_{t+1}^{\$} R_{t+1}^{n,j} \frac{S_t^{j,\$}}{S_{t+1}^{j,\$}}\right] \text{ for every } i, j. \tag{1}$$

For analytical tractability, we assume that the nominal SDF, bond prices, and exchange rate changes are log-normally distributed. First, we focus on one-period risk-free bonds. Denoting logs of variables with lowercase letters, the first equality in equation (1) gives

$$i_t^i = -E_t\left(m_{t+1}^i\right) - \frac{1}{2} Var_t\left(m_{t+1}^i\right), \tag{2}$$

[6] In general, monetary authorities set explicit targets for overnight rates. However, given that we are using quarterly data, the shortest bond maturity in our model is one quarter and our assumption is that the policy-setting behavior of the monetary authorities is the main determinant of the levels and dynamics of the 3-month yields at a quarterly frequency.

[7] This assumption that the trader's nominal SDF is denominated in U.S. dollars can be relaxed. We can assume that the unit of accounting is any currency or basket of currencies, which would be a realistic alternative if one interprets this trader as a global financial institution. More precisely, assume that the SDF of the global bank is denominated in a fictitious currency called ξ. All the derivations that follow will still hold, where the dollar-denominated SDF is related to the $\xi-$denominated SDF by the following accounting identity:

$$M_{t+1}^{\$} = M_{t+1}^{\xi} \frac{S_t^{\$,\xi}}{S_{t+1}^{\$,\xi}}.$$

where

$$m^i_{t+1} = m^\$_{t+1} - \Delta s^{i,\$}_{t+1} \tag{3}$$

is the trader's nominal SDF in terms of currency i and $\Delta s^{\$,\$}_{t+1} = 0$. In our model, equation (3) is simply an accounting identity. Equation (3) resembles the relationship that would hold under an assumption of complete markets linking nominal exchange rate changes to the nominal SDFs of traders in two separate countries—the distinction being that these traders could have different real SDFs. The assumption of complete markets amounts to an assumption about the degree of risk-sharing between the traders in different countries, but it is not necessary for our derivations since these are all made from the perspective of one single trader/institution.

The second equality in equation (1) delivers the following expressions for the expected and realized exchange rate changes. The realized exchange rate change is a function of relative lagged interest rates, currency risk premia, and an expectational error term:

$$\begin{aligned} E_t \Delta s^{i,j}_{t+1} &= i^i_t - i^j_t + \sigma^{i,j}_t && (4) \\ \Delta s^{i,j}_{t+1} &= i^i_t - i^j_t + \sigma^{i,j}_t + \Delta s^{i,j}_{t+1} - E_t \Delta s^{i,j}_{t+1}, && (5) \\ \text{where } \sigma^{i,j}_t &\equiv \frac{1}{2} Var_t \left(\Delta s^{i,\$}_{t+1} \right) - \frac{1}{2} Var_t \left(\Delta s^{j,\$}_{t+1} \right) - Cov_t \left(m^\$_{t+1}, \Delta s^{i,j}_{t+1} \right). \end{aligned}$$

In equation (5), $\sigma^{i,j}_t$ is the currency risk premium plus Jensen's inequality terms. If $\sigma^{i,j}_t$ is assumed to be zero, then equation (4) gives the strong form of the uncovered interest rate parity (UIRP) condition where the currency of country i is expected to depreciate when the interest rate of country i exceeds that of country j. Allowing for nonzero currency risk premia, a negative covariance between the SDF and $\Delta s^{i,j}_{t+1}$ would imply an even larger depreciation in this case, since currency i becomes a good hedge for the trader (that is, currency i gains value when the SDF is high) and the opposite holds if the covariance is positive.

In reality, there are numerous frictions that might violate the no-arbitrage condition given by equation (1). These include the inability of traders to borrow at the risk-free government bond rate, counterparty risk, as well as binding "net worth" or value-at-risk constraints. These frictions introduce various wedges in the no-arbitrage condition that are not captured by equation (1). Therefore, when we bring the model to the data, we more generally interpret equation (4) as an identity defining $\sigma^{i,j}_t$ as an expected excess return. Under this interpretation, the currency risk premium is only one of the components of this expected excess return. However, leading up to the results presented in section 4, we use the terms "currency risk premia" and "expected excess return" interchangeably.

In equation (5), $\Delta s^{i,j}_{t+1} - E_t \Delta s^{i,j}_{t+1}$ is an expectational error which is assumed to have a mean of zero and to be uncorrelated with variables in the information set used to form exchange rate expectations in period t. This term then contains only period $t+1$ information which cannot be used for forecasting. However, below we show that this term constitutes a large proportion of the variation in exchange rate changes and therefore, understanding the drivers of this term is important for understanding exchange rate dynamics. In particular, there is a key link between contemporaneous monetary policy actions—both conventional and unconventional—and exchange rate changes which appears through this expectational error term.

To further delve into this expectational error, we iterate equation (4) forward to obtain

$$s_t^{i,j} = -E_t \sum_{k=0}^{\infty} \left[i_{t+k}^i - i_{t+k}^j + \sigma_{t+k}^{i,j} \right] + E_t \lim_{k\to\infty} s_{t+k}^{i,j}. \tag{6}$$

First-differencing equation (6) and combining the resulting expression with equation (4) implies that the expectational error can be expressed as

$$\begin{aligned} \Delta s_{t+1}^{i,j} - E_t \Delta s_{t+1}^{i,j} \quad = \quad & -\underbrace{\left(\tilde{\imath}_{t+1}^{i,j} - E_t \tilde{\imath}_{t+1}^{i,j}\right)}_{\varphi_{t+1}^C} - \underbrace{\sum_{k=1}^{\infty} \left(E_{t+1} \tilde{\imath}_{t+k+1}^{i,j} - E_t \tilde{\imath}_{t+k+1}^{i,j}\right)}_{\varphi_{t+1}^F} \\ & - \underbrace{\sum_{k=0}^{\infty} \left(E_{t+1} \sigma_{t+k+1}^{i,j} - E_t \sigma_{t+k+1}^{i,j}\right)}_{\sigma_{t+1}^F} + \underbrace{E_{t+1} \lim_{k\to\infty} s_{t+k}^{i,j} - E_t \lim_{k\to\infty} s_{t+k}^{i,j}}_{s_{t+1,\infty}^{\Delta E}}, \end{aligned} \tag{7}$$

$$\begin{aligned} \text{where } \tilde{\imath}_t^{i,j} \quad &\equiv \quad i_t^i - i_t^j, \\ \varphi_{t+1}^{C,i} \quad &= \quad \left(i_{t+1}^i - E_t i_{t+1}^i\right), \text{ and } \quad \varphi_{t+1}^{F,i} = \sum_{k=1}^{\infty} \left(E_{t+1} i_{t+k+1}^i - E_t i_{t+k+1}^i\right). \end{aligned}$$

Equation (7) allows us to express the realized exchange rate changes as forward-looking variables which, in addition to the period t interest rate differential and currency premium, also reflect contemporanous surprises in relative policy rates as well as changes in expectations regarding future relative policy rates, currency risk premia, and long-run exchange rate levels. To simplify the notation, we suppress the "i, j" superscript for the labels of the decomposed terms from this point forward. Combining equations (4) and (7) implies that

$$\Delta s_{t+1}^{i,j} = \tilde{\imath}_t + \sigma_t - \varphi_{t+1}^C - \varphi_{t+1}^F - \sigma_{t+1}^F + s_{t+1,\infty}^{\Delta E}. \tag{8}$$

Two of the key terms that enter the expectational error in equation (8) are the unanticipated contemporaneous (period $t+1$) component of monetary policy, φ_{t+1}^C, and changes in expectations of future monetary policy (period $t+2$ and onwards), φ_{t+1}^F. Given that we focus on the link between monetary policy and exchange rates, the analysis will concentrate on the variation in exchange rate changes that can be attributed to φ^C_{t+1} and φ^F_{t+1} and how this quantity is related to the comovements between these two monetary policy terms and the remaining terms in the decomposition.

We start by providing more intuition for the interpretation of φ^C_{t+1} and φ^F_{t+1}. For the sake of brevity, we refer to φ^C_{t+1} as a "monetary policy surprise" from this point forward. In the literature using high frequency data, where periods are days or minutes, φ^C_{t+1} is often interpreted as an exogenous shock.[8] Given that we use quarterly data, φ^C_{t+1} contains not only exogenous monetary policy shocks, but also systematic policy responses to macroeconomic news (for example, new information about inflation or the output gap) that arrived within the quarter.[9] Therefore, our use of the term "monetary policy surprise" should not be confused with what is often referred to as a "monetary policy shock" in the literature.

The φ^F_{t+1} term captures changes in expectations of the path of future monetary policy. These expectational changes can arise from information aquired within the quarter that is useful for forecasting the future state of the economy and future monetary policy responses to those developments. Alternatively, traders might believe that there is inherent gradualism in the rate-setting process so that current monetary policy surprises foreshadow future policy rate changes. For details on monetary policy smoothing, where policymakers are assumed to smooth either the level or change of the policy rate, see Coibion and Gorodnichenko (2012).

Finally, even in the absence of contemporaneous macroeconomic news or monetary policy surprises, expectations regarding future policy may change due to a central bank engaging in unconventional monetary policy actions such as forward guidance or quantitative easing (QE). With forward guidance, policymakers explicitly communicate their plans for future policy and, if credible, these statements will inform the market's forecast of future monetary policy. QE can affect the market's expectations of future policy rates by acting as a signal that the central bank will be committed to expansionary policy for some time. For examples of empirical and theoretical work on the signaling channel of QE, see Bauer and Rudebusch (2014) and Bhattarai, Eggertsson, and Gafarov (2015).

2.1 Persistence of the Policy Rate

One potential driver of a high correlation between φ^C_{t+1} and φ^F_{t+1} is monetary policy persistence, as discussed above. To give some intuition for this mechanism and how policy persistence can

[8]See the references in the literature review.

[9]In a previous version of this paper we decomposed φ^C, which was calculated using a Taylor rule specification, into a systematic component that captured surprises about inflation and output gap, and a Taylor rule residual. The Taylor rule residual was very highly correlated with φ^C for most currency pairs, which is why in this version we focus solely on φ^C. The results from this decomposition are available upon request.

strengthen the link between contemporaneous monetary policy surprises and exchange rate dynamics, this section considers a parsimonious representation of the policy rate as an AR(1) process,

$$i^j_{t+1} = \rho_j i^j_t + \varepsilon^j_{t+1}, \tag{9}$$

where ε^i_{t+1} has a mean of zero and is a serially uncorrelated random variable. In Appendix Table A2, we present the results from the simple AR(1) regression for a sample starting in 1990:Q1 and the following countries: Australia, Canada, Germany/euro area, Japan, New Zealand, Norway, Sweden, Switzerland, the United Kingdom, and the United States.[10] These countries were selected based on the availability of data pertaining to zero-coupon bond yields in addition to having low default probabilities and highly liquid sovereign debt. We restrict the analysis to the post-1990 period since the importance of global banks and the carry trade have risen substantially since the advent of electronic trading during the late 1980s and the early 1990s, thus creating an environment in which the no-arbitrage conditions are more likely to hold. See Section A.1 in the Appendix for a description of the data. As the AR(1) regressions of the policy rates in Table A2 show, short-term interest rates are highly persistent with estimated ρ's above 0.9 for all but two of these countries. This simple AR(1) process also is a good descriptor of the level of the interest rate, as evidenced by adjusted R^2s of above 0.9 for nine out of the 10 countries.

Under equation (9), our monetary policy terms, φ^C_{t+1} and φ^F_{t+1}, are

$$\begin{aligned} \varphi^C_{t+1} &= \varepsilon^i_{t+1} - \varepsilon^j_{t+1} \\ \varphi^F_{t+1} &= \frac{\rho_i}{1-\rho_i}\varepsilon^i_{t+1} - \frac{\rho_j}{1-\rho_j}\varepsilon^j_{t+1}. \end{aligned} \tag{10}$$

Consider the special case $\rho = \rho_i = \rho_j$, which does not appear to be an unrealistic assumption for our sample of countries based on the results in Table A2. In this case, φ^C_{t+1} and φ^F_{t+1} are perfectly correlated. Holding the variance of $\varepsilon^i_{t+1} - \varepsilon^j_{t+1}$ constant, the variance of φ^F_{t+1}, and likewise the covariance between φ^C_{t+1} and φ^F_{t+1}, is increasing in the persistence of the policy rate. All else equal, a larger ρ implies a larger contribution of monetary policy surprises, φ^C_{t+1}, to variation in the exchange rate change through its effect on future expectations, φ^F_{t+1}.

Combining equations (8) and (10), we can express the exchange rate change as

$$\Delta s^{i,j}_{t+1} = -\frac{1}{1-\rho}\Delta\tilde{i}^{i,j}_{t+1} - \sigma^F_{t+1} + s^{\Delta E}_{t+1,\infty} + \sigma_t, \tag{11}$$

which shows that there is a theoretical link between changes in policy rates in the two countries

[10]Throughout the paper, short-term interest rates and macroeconomic data for Germany are replaced with euro-area data starting in 1999:Q1. For long-term yields, German zero-coupon yields are used throughout the entire sample period since our analysis applies to risk-free sovereign debt.

and exchange rate dynamics. Figure 1 shows that this relationship is also present in the data. If the correlation between the interest rate changes and the remaining terms, $-\sigma^F_{t+1} + s^{\Delta E}_{t+1,\infty} + \sigma_t$, is fairly small and if the AR(1) model is an approximately accurate representation of how traders form expectations, our ρ estimates imply that the β_1 coefficient in the regression,

$$\Delta s^{i,j}_{t+1} = \alpha_1 + \beta_1 \Delta \tilde{i}^{i,j}_{t+1} + \epsilon^1_{t+1}, \tag{12}$$

should be between -8 and -33. Table 1 reports the results of this second regression.

The β_1 coefficients are significant and negative (as suggested by equation (11)) in 18 out of 30 cases and, in most cases, β_1 is within the theoretical range suggested by a simple AR(1). This result is quite remarkable, considering that this regression may have omitted variable bias due to the missing $-\sigma^F_{t+1} + s^{\Delta E}_{t+1,\infty} + \sigma_t$ term. Hence, Table 1 provides suggestive evidence that for many currency pairs, the historical relationship between interest rate changes and exchange rate changes is close to the one implied by a simple AR(1) process for short-term interest rates with no correlation between $-\sigma^F_{t+1} + s^{\Delta E}_{t+1,\infty} + \sigma_t$ and $\Delta \tilde{i}^{i,j}_{t+1}$.

2.2 Exchange Rates and Long-Term Yields

Given that long-term yields directly capture future policy rate expectations, in addition to term premia, it should be interesting to explore the theoretical relationship between long-term yields and exchange rate changes. By iterating the log-normal version of the no-arbitrage condition for long-term bonds forward and combining it with equation (2), one can derive the following expression for expected future policy rates:

$$\frac{1}{n} \sum_{k=0}^{n-1} E_t i^i_{t+k} = y^{n,i}_t - tp^{n,i}_t, \tag{13}$$

where the term premium is given by

$$tp^{n,i}_t = \frac{1}{2n} \left[\sum_{k=1}^{n} E_t Var_{t+k-1} \left(m^i_{t+k} \right) - Var_t \left(m^i_{t,t+n} \right) \right].$$ [11]

Substituting equation (13) into equations (7) and (5) reveals that the quarterly exchange rate change can be expressed as a function of the change in relative n-period bond yields,

$$\Delta s^{i,j}_{t+1} = -n \Delta \tilde{y}^{n,i,j}_{t+1} - \left(\sum_{k=n}^{\infty} E_{t+1} \tilde{i}^{i,j}_{t+k+1} - \sum_{k=n}^{\infty} E_t \tilde{i}^{i,j}_{t+k} \right) + n \Delta \widetilde{tp}^{n,i,j}_{t+1} - \sigma^F_{t+1} + s^{\Delta E}_{t+1,\infty} + \sigma_t, \tag{14}$$

where $\widetilde{tp}_t^{n,i,j} \equiv tp_t^{n,i} - tp_t^{n,j}$ and $\tilde{y}_t^{n,i,j} \equiv y_t^{n,i} - y_t^{n,j}$. The theory predicts that if there is no correlation between the changes in relative long-term yields and the remaining terms in equation (14), then we would expect a coefficient of $\beta_2 = 1$ in the following regression:

$$\Delta s_{t+1}^{i,j} = \alpha_2 + \beta_2 n \Delta \tilde{y}_{t+1}^{n,i,j} + \epsilon_{t+1}^2. \tag{15}$$

The results from this regression are presented in Table 2.

The estimates of β_2 are indeed negative and significant in 19 out of 30 cases. In 18 of those 19 cases, we cannot reject the null hypothesis that β_2 is equal to -1 at standard confidence levels.

To summarize, reduced-form regressions tell us that the theory-implied empirical relationships between exchange rate dynamics and changes in relative short-term interest rates or long-term yields are not completely overpowered by omitted variable bias. This conclusion stands in contrast to standard uncovered interest rate parity regressions, where the comovement between $\tilde{\imath}_t$ and σ_t creates an omitted variable bias that is strong enough to lead to a rejection of the theory-implied coefficient. Appendix Table A3 presents the results from the UIRP regression—the coefficients are not only insignificant and of the wrong magnitude, but also of the wrong sign for some countries.[12]

Though the reduced-form regressions are highly suggestive, they give us only part of the picture. Many questions remain, such as what fraction of the variance of the exchange rate change can be attributed to monetary policy relative to the other terms in the decomposition? Additionally, do the correlations between monetary policy terms and the remaining terms in the decomposition amplify or dampen the reduced-form relationships between monetary policy and exchange rate dynamics? If so, which of the other terms in the decomposition contribute the most to amplification or dampening?

In order to be able to answer these questions, the next section describes how we can calculate the various terms in equation (8) using a VAR approach.

3 VAR Approach

We use VAR-based expectations of interest rates and changes in exchange rates to compute the terms in equation (8). We will consider two specifications that differ in the sets of variables used to describe market participants' expectations. In the first specification, we follow the macroeconomics literature in assuming that policy rate expectations are based on an estimated Taylor rule. In the second specification, we assume that policy rate expectations are formed using information

[12]The latter is known as the forward premium puzzle, which indicates omitted variable bias driven by a strong correlation between the relative policy rates and the period t currency risk premium.

obtained from the yield curve, a specification that is commonly used in the asset pricing literature. One caveat is that we do not impose a zero lower bound in the VAR. However, we observe very few negative policy rate forecasts and the implied zero lower bound violations are very small. Therefore, given the computational complexity that introducing a zero lower bound would entail, we abstract from this constraint.

We assume that exchange rates and short-term interest rates can be described by the following structural VAR(p) process:

$$\begin{aligned} F_{t+1} &= \tilde{\mathbf{F}} + \tilde{\boldsymbol{\gamma}}_0 F_{t+1} + \tilde{\boldsymbol{\gamma}}_1 F_t + ... + \tilde{\boldsymbol{\gamma}}_p F_{t-p+1} + \tilde{\varepsilon}_{F,t+1}, \\ \text{where } F_{t+1} &\equiv [q^{i,j}_{t+1}, i^i_{t+1}, z^i_{t+1}, i^j_{t+1}, z^j_{t+1}, i^{US}_{t+1}, z^{US}_{t+1}]' \end{aligned} \tag{16}$$

and the diagonal elements in $\tilde{\boldsymbol{\gamma}}_0$ are normalized to zero. Here, $q^{i,j}_{t+1}$ is the level of the real exchange rate between currencies i and j. By including the real exchange rate in levels, we are estimating a specifiation where a stable estimate of the VAR implies that long-run purchasing power parity holds and VAR-based expectations of the long-run real exchange rate are constant. The vector z^c_{t+1} for $c \in \{i, j, US\}$ represents other variables which may be useful for forecasting either short-term interest rates or changes in the exchange rate. Importantly, we always include a quarterly inflation rate (measured using the GDP deflator) in z^c_{t+1}. This allows us to compute VAR-based expectations of nominal exchange rate changes from our estimates of the real exchange rate and inflation equations. In our Taylor rule specification, z^c_{t+1} also includes the GDP gap and the current account-to-GDP ratio. In our second specification, the z^i_{t+1} vector also includes the empirical term structure slope and curvature factors, which are defined below.

The corresponding reduced-form representation of this VAR is

$$F_{t+1} = \underbrace{(\mathbf{I} - \tilde{\boldsymbol{\gamma}}_0)^{-1} \tilde{\mathbf{F}}}_{\bar{\mathbf{F}}} + \underbrace{(\mathbf{I} - \tilde{\boldsymbol{\gamma}}_0)^{-1} \tilde{\boldsymbol{\gamma}}_1}_{\boldsymbol{\gamma}_1} F_t + ... + \underbrace{(\mathbf{I} - \tilde{\boldsymbol{\gamma}}_0)^{-1} \tilde{\boldsymbol{\gamma}}_p}_{\boldsymbol{\gamma}_p} F_{t-p+1} + \underbrace{(\mathbf{I} - \tilde{\boldsymbol{\gamma}}_0)^{-1} \tilde{\varepsilon}_{F,t+1}}_{\varepsilon_{F,t+1}}. \tag{17}$$

This reduced-form VAR(p) can be transformed into a VAR(1) by stacking lags of F into a single vector $\mathbf{X}$ as follows:

$$\underbrace{\begin{bmatrix} F_{t+1} \\ \vdots \\ F_{t-p+2} \end{bmatrix}}_{\mathbf{X}_{t+1}} = \underbrace{\begin{bmatrix} \bar{\mathbf{F}} \\ 0 \\ 0 \end{bmatrix}}_{\bar{\mathbf{X}}} + \underbrace{\begin{bmatrix} \boldsymbol{\gamma}_1 & \boldsymbol{\gamma}_2 & \cdots & \boldsymbol{\gamma}_p \\ & \mathbf{I} & & \mathbf{0} \end{bmatrix}}_{\boldsymbol{\Gamma}} \underbrace{\begin{bmatrix} F_t \\ \vdots \\ F_{t-p+1} \end{bmatrix}}_{\mathbf{X}_t} + \underbrace{\begin{bmatrix} \varepsilon_{F,t+1} \\ 0 \\ \vdots \end{bmatrix}}_{\boldsymbol{\Xi}_{t+1}}.$$

We estimate two different specifications of this VAR by OLS, choosing a lag length of $p = 2$ in both specifications. In the Taylor rule specification, we impose the following restrictions:

Taylor Rule Specification (*TR*)

1. We restrict the short-term interest rate equation for country i to follow a Taylor rule of the form,

$$\begin{aligned} i^i_{t+1} &= \tilde{\imath}^i + \rho^i_1 i^i_t + \rho^i_2 i^i_{t-1} + \psi^i_{\pi,0}\pi^i_{t+1} + \psi^i_{g,0} g^i_{t+1} \\ &\quad + \psi^i_{\pi,1}\pi^i_t + \psi^i_{g,1} g^i_t + \psi^i_{\pi,2}\pi_{t-1} + \psi^i_{g,2} g^i_{t-1} + \tilde{\varepsilon}^i_{MP,t+1}, \end{aligned} \tag{18}$$

   and use an analogous functional form for i^j_{t+1}. This means that in our model, each country's short-term interest rate is assumed to depend only on two lags of itself as well as contemporaneous realizations plus two lags each of the domestic inflation rate and the output gap.

2. z^i_{t+1} depends only on $\left\{i^i_{t-l}, z^i_{t-l}, i^{US}_{t-l}, z^{US}_{t-l}\right\}^1_{l=0}$ and z^j_{t+1} depends only on $\left\{i^j_{t-l}, z^j_{t-l}, i^{US}_{t-l}, z^{US}_{t-l}\right\}^1_{l=0}$.
3. z^{US}_{t+1} depends only on $\left\{i^{US}_{t-l}, z^{US}_{t-l}\right\}^1_{l=0}$.
4. $\left\{q^{i,j}_{t-l}\right\}^1_{l=0}$ enters only in the $q^{i,j}_{t+1}$ equation and not the equations for $\left\{i^i_{t+1}, z^i_{t+1}, i^j_{t+1}, z^j_{t+1}, i^{US}_{t+1}, z^{US}_{t+1}\right\}$.

The Taylor rule restrictions are made in the spirit of Leeper, Sims, and Zha (1996) and Sims and Zha (2006a;b) as we are restricting the set of variables that each country's central bank responds to when setting the policy rate. However, note that the precise restrictions differ since our baseline specification does allow the interest rate to respond contemporaneously to inflation and the output gap. This approach seems seems plausible to use at the quarterly frequency since some contemporaneous statistical measures of these variables are made available within the quarter.[13] In addition, restricting the short-term interest rate equation in this way addresses some of the criticisms presented in Rudebusch (1998).

In order to implement the Taylor rule restrictions, we estimate a hybrid structural and reduced-form specification by allowing the appropriate entries of $\tilde{\gamma}_0$ in the rows corresponding to the short-term interest rate equations to be estimated while the remaining entries are constrained to be zero. We then transform the entire system into the full reduced form to proceed with the computation of our exchange rate components.[14]

Note that while we are not explicitly estimating the structural contemporaneous relationships in equations other than our short-term interest rate equations, this estimation procedure does not preclude the existence of these relationships. More specifically, our method can accommodate

[13] Measures of industrial production as well as consumer and producer price inflation are released with a one-month lag, so some information about real activity and inflation in the first two months of the quarter is available by the end of the quarter.

[14] An alternate method of obtaining the reduced-form VAR estimates would be to estimate a constrained system where i_{t+1} depends on only lagged $\{i_{t-l}, z_{t-l}\}^1_{l=0}$, but where the coefficients on lags of the current account-to-GDP ratio are constrained to be consistent with the way in which they enter into inflation and the output gap.

the following contemporaneous relationships without explicitly estimating them: (i) dependence of $\Delta s^{i,j}_{t+1}$ on all other variables in F_{t+1}, (ii) dependence of z^i_{t+1} on $\left\{i^i_{t+1}, z^i_{t+1}, i^{US}_{t+1}, z^{US}_{t+1}\right\}$ (and analogously for z^j_{t+1}), and (iii) dependence of z^{US}_{t+1} on $\left\{i^{US}_{t+1}, z^{US}_{t+1}\right\}$.

In this specification of the short-term interest rate, we choose to use Taylor rules since they have been widely used to approximate monetary policy. A main feature that helps produce a better fit of the policy rate is to allow the interest rate to be persistent in *changes* rather than just levels. This implies not only that policymakers do not want to deviate too much from the previous policy rate, but also that they adjust gradually to the new target rate over a few consecutive quarters. This specification can arise in a variety of optimal policy frameworks—see Brainard (1967) and Woodford (2003)—and has been used successfully in empirical work by, among others, Judd and Rudebusch (1998) and Clarida and Waldman (2008). Coibion and Gorodnichenko (2012) also show that this specification is especially important for fitting U.S. monetary policy. According to a model featuring second-order partial adjustment, the policy rate evolves as follows:

$$\begin{aligned} \triangle i_{t+1} &= \left(1-\rho^{\Delta}_1\right)\left(i^*_{t+1}-i_t\right)+\rho^{\Delta}_2 \triangle i_t+\tilde{\varepsilon}_{MP,t+1}, \\ \text{where } i^*_{t+1} &= r^*+\pi^*+\psi^*_{\pi,0}\pi_{t+1}+\psi^*_{g,0}g_{t+1} \\ &\quad +\psi^*_{\pi,1}\pi_t+\psi^*_{g,1}g_t+\psi^*_{\pi,2}\pi_{t-1}+\psi^*_{g,2}g_{t-1}. \end{aligned} \tag{19}$$

Note that $\rho^{\Delta}_2 = 0$ would correspond to a Taylor rule with only persistence in levels:

$$i_{t+1} = \rho_1 i_t + (1-\rho_1)\, i^*_{t+1} + \tilde{\varepsilon}_{MP,t+1}.$$

Rearranging equation (19) gives the form that we have in equation (18) above with:

$$\begin{aligned} \tilde{\imath} &= \left(1-\rho^{\Delta}_1\right)\left(r^*+\pi^*\right),\ \psi = \left(1-\rho^{\Delta}_1\right)\psi^*, \\ \rho_1 &= \rho^{\Delta}_1+\rho^{\Delta}_2,\ \text{and}\ \rho_2 = -\rho^{\Delta}_2. \end{aligned}$$

Assumptions 2 and 3 require further explanation. We assume that changing economic conditions in the United States affect expectations about future monetary policy in other countries through spillovers from the United States into the macroeconomies of these other countries. Commodity-producing countries seem to be the most sensitive to the lagged U.S. output gap. This result is not surprising given the high sensitivity of growth in these economies to world growth. See Miranda-Agrippino and Rey (2015) for VAR-based evidence of such spillovers.

When modeling the policy rate with a Taylor rule over a sample that includes the zero lower bound period, one criticism is that it provides poor forecasts of policy rates over longer horizons. Furthermore, if forward guidance was indeed successful, long-horizon policy rate expectations implied by a Taylor rule will be flawed in the recent period for advanced economies and since the

2000s for Japan. While below we will present evidence that our VAR-implied interest rate expectations accord well with futures market data, the use of futures market data for this comparison is problematic for longer horizon forecasts due to term premia in future prices. In order to address this concern and to provide a more agnostic forecasting model for policy rates, following common practice in the asset pricing literature, our second specification assumes that policy rates are driven by lags of themselves in addition to the slope and curvature of the yield curve. This specification uses information from long-term yields that can capture the effects of forward guidance, among other things, and might therefore bring our VAR-based long-horizon policy rate forecasts closer to expectations held by financial markets.

For our second specification, we define the empirical slope and curvature factors as follows:

$$\begin{aligned} sl_{t+1}^{i} &= y_t^{40,i} - i_t^i \\ c_{t+1}^{i} &= 2y_t^{8,i} - \left(y_t^{40,i} + i_t^i\right). \end{aligned}$$

In this specification, the z_{t+1}^i vector contains sl_{t+1}^i and c_{t+1}^i in addition to inflation, the output gap, and the current account-to-GDP ratio.

Yield Factors Specification (*YF*)

1. $\left\{i_{t+1}^i, sl_{t+1}^i, c_{t+1}^i\right\}$ depends only on $\left\{i_{t-l}^i, sl_{t-l}^i, c_{t-l}^i\right\}_{l=0}^1$ and $\left\{i_{t+1}^j, sl_{t+1}^j, c_{t+1}^j\right\}$ depends only on $\left\{i_{t-l}^j, sl_{t-l}^j, c_{t-l}^j\right\}_{l=0}^1$.
2. z_{t+1}^i depends only on $\left\{i_{t-l}^i, z_{t-l}^i, i_{t-l}^{US}, z_{t-l}^{US}\right\}_{l=0}^1$ and z_{t+1}^j depends only on $\left\{i_{t-l}^j, z_{t-l}^j, i_{t-l}^{US}, z_{t-l}^{US}\right\}_{l=0}^1$.
3. z_{t+1}^{US} depends only on $\left\{i_{t-l}^{US}, z_{t-l}^{US}\right\}_{l=0}^1$.
4. $\left\{q_{t-l}^{i,j}\right\}_{l=0}^1$ enters only in the $q_{t+1}^{i,j}$ equation and not the equations for $\left\{i_{t+1}^i, z_{t+1}^i, i_{t+1}^j, z_{t+1}^j, i_{t+1}^{US}, z_{t+1}^{US}\right\}$.

In this specification, we do not account explicitly for any contemporaneous relationship and estimate the VAR in its reduced form.

As a check for external validity, we can compare our model-implied interest rate expectations with market-based measures of short-term interest rate surprises computed using futures prices by adapting the method used in Bernanke and Kuttner (2005) to a quarterly frequency. In the *TR* specification, the reduced-form expected U.S. short-term interest rate change from our model, $\hat{E}_t\left[i_{t+1}^{US}\right] - i_t^{US}$, has a correlation of 62 percent (over the full sample) with the quarterly expected component of U.S. 3-month eurodollar interest rate changes. This correlation increases to 65 percent over the pre-ZLB period ending in 2008:Q4. Furthermore, our model-implied quarterly interest rate surprise, $i_{t+1}^{US} - \hat{E}_t\left[i_{t+1}^{US}\right]$, has a correlation of 69 percent with the market-based 3-month eurodollar interest rate surprise measure. This correlation is 71 percent over the pre-ZLB period ending in 2008:Q4. For the *YF* specification, $\hat{E}_t\left[i_{t+1}^{US}\right] - i_t^{US}$, has a correlation of 75 percent with the expected component of U.S. 3-month eurodollar interest rate change over the full sample and 83 percent over the pre-ZLB period. The *YF* model-implied quarterly interest rate surprise, $i_{t+1}^{US} - \hat{E}_t\left[i_{t+1}^{US}\right]$, has a correlation of 77 percent with the market-based federal funds rate surprise measure over the full sample and 79 percent over the pre-ZLB period. Tables 3 and 4 show these correlations over the full sample for a number of additional countries. The high correlations for a large majority of the countries we consider are evidence that our VAR-based interest rate expectations accord reasonably well with expectations held by financial market participants.[15]

With the estimated VARs, we can now decompose exchange rates into the six terms listed in equation (8). First, using the notation of the VAR, the exchange rate change and relative policy rates can be expressed as

$$
\begin{aligned}
\Delta s_{t+1}^{i,j} &\equiv \Delta q_{t+1}^{i,j} + \pi_{t+1}^{i} - \pi_{t+1}^{j} = \left(e_q + e_{\pi,i} - e_{\pi,j}\right)\mathbf{X}_{t+1} - e_q\mathbf{X}_t \\
\tilde{\imath}_{t+1}^{i,j} &= \left(e_i^i - e_i^j\right)\mathbf{X}_{t+1},
\end{aligned}
$$

where e_q is a row vector that selects $q_{t+1}^{i,j}$ from $\mathbf{X}_{t+1}$. That is, it has the same number of elements as $\mathbf{X}_{t+1}$ with an entry of 1 corresponding to the position of $q_{t+1}^{i,j}$ in $\mathbf{X}_{t+1}$ and zeros elsewhere. Likewise, e_i^i and e_i^j are selection vectors corresponding to the short-term interest rates of countries i and j, respectively.

VAR-based expectations of nominal exchange rates and interest rates can be obtained from those for future $\mathbf{X}_{t+k+1}$, as follows:[16]

$$
\begin{aligned}
\hat{E}_t\mathbf{X}_{t+k+1} &= \left(\mathbf{I}-\mathbf{\Gamma}\right)^{-1}\left(\mathbf{I}-\mathbf{\Gamma}^{k+1}\right)\bar{\mathbf{X}} + \mathbf{\Gamma}^{k+1}\mathbf{X}_t \\
\hat{E}_t\Delta s_{t+k+1}^{i,j} &= \left(e_q + e_{\pi,i} - e_{\pi,j}\right)\left[\left(\mathbf{I}-\mathbf{\Gamma}\right)^{-1}\left(\mathbf{I}-\mathbf{\Gamma}^{k+1}\right)\bar{\mathbf{X}} + \mathbf{\Gamma}^{k+1}\mathbf{X}_t\right] \\
&\quad -e_q\left[\left(\mathbf{I}-\mathbf{\Gamma}\right)^{-1}\left(\mathbf{I}-\mathbf{\Gamma}^{k}\right)\bar{\mathbf{X}} + \mathbf{\Gamma}^{k}\mathbf{X}_t\right] \\
\hat{E}_t\tilde{\imath}_{t+k+1} &= \left(e_i^i - e_i^j\right)\left[\left(\mathbf{I}-\mathbf{\Gamma}\right)^{-1}\left(\mathbf{I}-\mathbf{\Gamma}^{k+1}\right)\bar{\mathbf{X}} + \mathbf{\Gamma}^{k+1}\mathbf{X}_t\right].
\end{aligned}
$$

Note that the change in expectations over future $\mathbf{X}_{t+k+1}$ can be written simply as a linear combination of the time $t+1$ reduced-form residuals:

$$
\hat{E}_{t+1}\mathbf{X}_{t+k+1} - \hat{E}_t\mathbf{X}_{t+k+1} = \mathbf{\Gamma}^k\mathbf{\Xi}_{t+1}.
$$

[15]Note that the futures contracts we use typically are written on interbank interest rates while our VAR produces expectations of 3-month T-bill rates. By basing our comparisons on expected interest rate changes and surprises, we are able to abstract from differences in the rates that do not vary at a quarterly frequency. Nonetheless, the differences in the financial instruments might make it harder to detect a high correlation between our VAR-implied expectations and the ones implied by futures prices even if our VAR accords well with the market's expectations formation process.

[16]The $\hat{E}_t$ operator denotes expectations based on the linear projections performed in the VAR estimation. Although not explicitly delineated, the operator conditions only on the set of regressors included in the estimation of each equation. Due to the restrictions set out above, this means that the relevant information set differs across variables.

As long as all eigenvalues of $\mathbf{\Gamma}$ are strictly inside the unit circle in the complex plane, then the exchange rate change components can be constructed using equations (4) and (7) as follows:

$$\begin{aligned}
\tilde{i}_t^{i,j} &= \left(e_i^i - e_i^j\right)\mathbf{X}_t \\
\sigma_t &= \left(e_q + e_\pi^i - e_\pi^j\right)\left(\bar{\mathbf{X}} + \mathbf{\Gamma}\mathbf{X}_t\right) - \left(e_q + e_i^i - e_i^j\right)\mathbf{X}_t \\
\varphi_{t+1}^C &= \left(e_i^i - e_i^j\right)\mathbf{\Xi}_{t+1} \\
\varphi_{t+1}^F &= \left(e_i^i - e_i^j\right)\mathbf{\Gamma}\left(\mathbf{I} - \mathbf{\Gamma}\right)^{-1}\mathbf{\Xi}_{t+1} \\
\sigma_{t+1}^F &= \left[\left(e_q + e_\pi^i - e_\pi^j\right)\mathbf{\Gamma} - \left(e_q + e_i^i - e_i^j\right)\right]\left(\mathbf{I} - \mathbf{\Gamma}\right)^{-1}\mathbf{\Xi}_{t+1} \\
s_{t+1,\infty}^{\Delta E} &= \left(e_\pi^i - e_\pi^j\right)\left(\mathbf{I} - \mathbf{\Gamma}\right)^{-1}\mathbf{\Xi}_{t+1}.
\end{aligned} \tag{20}$$

One key thing to note above is that since our specification implies constant expectations over long-run levels of the real exchange rate, the change in expectations regarding long-run nominal exchange rates, $s_{t+1,\infty}^{\Delta E}$, simply reflects changes in expectations over long-run relative prices levels. Hence, this component can be computed from VAR-based expectations of the future path of inflation rates.

Note that in the subsequent sections, the results for the NOK-JPY and NZD-JPY pairs for specification *YF* are not presented because the VAR estimates in this case violate the requirement that all eigenvalues of $\mathbf{\Gamma}$ are strictly inside the unit circle in the complex plane. When we conduct subsample analyses in the *YF* case, we also do not report results for pairs including JPY in the most recent subsample since our yields data for Japan end in 2009:Q2 (see the Section A.1 in the Appendix for details on the data).

4 The Link Between Monetary Policy and Exchange Rate Changes

4.1 Further Decomposing Monetary Policy

Before presenting the results from our exercise, we introduce an alternate decomposition of our total monetary policy term, $\varphi_{t+1}^{MP} \equiv \varphi_{t+1}^C + \varphi_{t+1}^F$. This decomposition will allow us to account for the fact that the period $t+1$ monetary policy surprise for a currency i, $\varphi_{t+1}^{C,i}$, is one of the main drivers of the changes in expectations over future policy rates, $\varphi^{F,i}$, due to the strong persistence of short-term interest rates. Hence, we decompose $\varphi_{t+1}^{MP,i} \equiv \varphi_{t+1}^{F,i} + \varphi_{t+1}^{C,i}$ by orthogonalizing $\varphi_{t+1}^{F,i}$ with respect to $\varphi_{t+1}^{C,i}$ in order to obtain a component of the expectations over future policy rates which is uncorrelated with the surprise in the current short-term policy rate. More precisely, we perform a Cholesky decomposition of $\varphi_{t+1}^{F,i}$ and $\varphi_{t+1}^{C,i}$ to obtain two new orthogonalized subcomponents of

$\varphi_{t+1}^{MP,i}$,

$$\begin{aligned}
\varphi_{t+1}^{P,i} &\equiv \varphi_{t+1}^{F,i} - \frac{Cov\left(\varphi_{t+1}^{F,i}, \varphi_{t+1}^{C,i}\right)}{Var\left(\varphi_{t+1}^{C,i}\right)} \varphi_{t+1}^{C,i} \qquad (21) \\
\text{and } \varphi_{t+1}^{T,i} &\equiv \left(1 + \frac{Cov\left(\varphi_{t+1}^{F,i}, \varphi_{t+1}^{C,i}\right)}{Var\left(\varphi_{t+1}^{C,i}\right)}\right) \varphi_{t+1}^{C,i} \\
\text{where we define } \varphi_{t+1}^{P} &\equiv \varphi_{t+1}^{P,i} - \varphi_{t+1}^{P,j} \\
\text{and } \varphi_{t+1}^{T} &\equiv \varphi_{t+1}^{T,i} - \varphi_{t+1}^{T,j}.
\end{aligned}$$

Hence, we have a new decomposition of the total monetary policy term, $\varphi_{t+1}^{MP} = \varphi_{t+1}^{P} + \varphi_{t+1}^{T}$. The term $\varphi_{t+1}^{T,i}$ is proportional to $\varphi_{t+1}^{C,i}$ and can be interpreted as the full effect of the contemporaneous monetary policy on the total monetary policy term, $\varphi_{t+1}^{MP,i}$, inclusive of its effect on future policy rate expectations. On the other hand, $\varphi_{t+1}^{P,i}$ reflect changes in the market's expectations of the future path of relative policy rates that is orthogonal to the surprise in the current policy rate.[17] The superscripts ^{T} and ^{P} are a reference to the oft-mentioned "target" and "path" factors of monetary policy, as introduced by Gürkaynak, Sack, and Swanson (2005), since our decomposition is conceptually similar to theirs.

Given the presence of a binding zero lower bound and the increased use of unconventional monetary policy such as forward guidance and QE for many advanced economies, the link between φ_{t+1}^{P} and the exchange rate change becomes particularly interesting in the recent period since December 2008.[18] The *YF* specification will also be of particular interest in this exercise because it can better capture changes in policy expectations due to unconventional policy measures.

Equipped with the decompositions given by equations (8) and (21), we proceed to investigate the relationship between monetary policy and exchange rates. The empirical exchange rates literature has taken two approaches to studying this relationship. One approach, employed by Froot and Ramadorai (2005), Engel and West (2010), and Evans (2012), is to use decompositions similar to equation (8) to study the fraction of the volatility explained by the different components of the exchange rate change. This type of analysis is in the spirit of the variance decomposition pioneered by Campbell and Shiller (1988a;b) and further developed in Campbell (1991) and Campbell and Ammer (1993). Another approach is to examine the coefficients obtained by regressing changes in the exchange rate on particular components. Standard UIRP regressions are one version of this exercise in which the relevant component is the lagged relative policy rate. The VAR literature

[17] In the AR(1) process example considered in Section 2.1, we would have $\varphi^{P} = 0$ by construction.

[18] QE is assumed to have worked both through a portfolio balance channel and through a signalling channel. It is this latter channel that would be captured in φ^{P}.

has also delivered estimates of variance decompositions and responses with the main difference from our exercise being that these VAR results are with respect to an alternate decomposition of the exchange rate into the effects of various orthogonal exogenous shocks (see Eichenbaum and Evans (1995), for example). Using our decompositions, we show that variance decompositions and estimates from univariate regressions are very much related and the questions addressed by both approaches can be studied jointly.

We begin with the variance decomposition approach. Given the historical distribution of shocks in the data, we examine the fraction of exchange rate volatility that can be explained by each one of the components in equation (8). Since these components are not orthogonal, we cannot simply report relative variances of each component versus exchange rates. One potential approach to accounting for the nonzero covariances between components is to take a stand on the ordering of the components and assign the covariances to one of the two underlying components. Instead, we take an agnostic approach, splitting each covariance evenly between the two relevant components (see Gourinchas and Rey (2007) for an example of this approach in a different context). This method is equivalent to reporting the covariances between each term in the decomposition and the exchange rate change, each scaled by the variance of the exchange rate change:

$$1 = \frac{Cov\left(\tilde{\imath}_t, \Delta s_{t+1}\right)}{Var\left(\Delta s_{t+1}\right)} + \frac{Cov\left(\sigma_t - \sigma^F_{t+1}, \Delta s_{t+1}\right)}{Var\left(\Delta s_{t+1}\right)} + \frac{Cov\left(-\varphi^{MP}_{t+1}, \Delta s_{t+1}\right)}{Var\left(\Delta s_{t+1}\right)} + \frac{Cov\left(s^{\Delta E}_{t,\infty}, \Delta s_{t+1}\right)}{Var\left(\Delta s_{t+1}\right)}. \tag{22}$$

Since this decomposition sums to 1, we can interpret these scaled covariances as the shares of the variance of the exchange rate change that can be explained by each component from the decomposition. Note that these scaled covariances are also equivalent to coefficients from univariate regressions of each component on the exchange rate change. Given that this is not an orthogonal decomposition, the scaled covariance terms may be negative and greater than one in absolute value.

Note now that each of these scaled covariance measures can be related to both the R^2 and regression coefficient from univariate regressions of the exchange rate change on the relevant component as follows:

$$\frac{Cov\left(x, \Delta s_{t+1}\right)}{Var\left(\Delta s_{t+1}\right)} = \frac{Var\left(x\right)}{Var\left(\Delta s_{t+1}\right)} \underbrace{\frac{Cov\left(x, \Delta s_{t+1}\right)}{Var\left(x\right)}}_{\beta_x} = \frac{R^2_x}{\beta_x}$$

$$\text{for } x \in \left\{\tilde{\imath}^{i,j}_t, \sigma_t - \sigma^F_{t+1}, -\varphi^{MP}_{t+1}, s^{\Delta E}_{t,\infty}\right\},$$

where R^2_x and β_x pertain to the regression,

$$\Delta s_{t+1} = \alpha_x + \beta_x x + \epsilon^x_{t+1}.$$

The main measure of interest will be the fraction of the variance of the exchange rate change that can be attributed to monetary policy—both to contemporaneous monetary policy surprises as well as changes in expectations of future monetary policy. Using (21), $\frac{Cov\left(\varphi_{t+1}^{MP},\Delta s_{t+1}\right)}{Var(\Delta s_{t+1})}$ can be further decomposed into the parts due to: (i) Π^C, the relative monetary policy surprise and its effect on changes in expectations over future policy rates (through persistence), and (ii) Π^P, the change in future policy rate expectations that is orthogonal to the contemporaneous monetary policy surprise.

$$\begin{aligned}
\frac{Cov\left(\varphi_{t+1}^{MP},\Delta s_{t+1}\right)}{Var\left(\Delta s_{t+1}\right)} &= \underbrace{\frac{Cov\left(\varphi_{t+1}^{T},\Delta s_{t+1}\right)}{Var\left(\Delta s_{t+1}\right)}}_{\Pi^C} + \underbrace{\frac{Cov\left(\varphi_{t+1}^{P},\Delta s_{t+1}\right)}{Var\left(\Delta s_{t+1}\right)}}_{\Pi^P} \\
&= \underbrace{\beta_{\varphi^C}\frac{Var\left(\varphi_{t+1}^{C}\right)+\frac{\beta_{\varphi^{C,i}}}{\beta_{\varphi^C}}Cov\left(\varphi_{t+1}^{F,i},\varphi_{t+1}^{C,i}\right)-\frac{\beta_{\varphi^{C,j}}}{\beta_{\varphi^C}}Cov\left(\varphi_{t+1}^{F,j},\varphi_{t+1}^{C,j}\right)}{Var\left(\Delta s_{t+1}\right)}}_{\Pi^C} \\
&\quad \underbrace{+\beta_{\varphi^P}\frac{Var\left(\varphi_{t+1}^{P}\right)}{Var\left(\Delta s_{t+1}\right)}}_{\Pi^P}, \qquad (23) \\
\text{where } \beta_{\varphi^x} &\equiv \frac{Cov\left(\varphi_{t+1}^{x},\Delta s_{t+1}\right)}{Var\left(\varphi_{t+1}^{x}\right)} \text{ for } x\in\{C,P,\{C,i\},\{C,j\}\}.
\end{aligned}$$

Note that β_{φ^x} is the OLS regression coefficient obtained from regressing the exchange rate change on φ_{t+1}^{x}.

If the countries are roughly symmetric such that $Cov\left(\varphi_{t+1}^{F,j},\varphi_{t+1}^{C,j}\right)\approx Cov\left(\varphi_{t+1}^{F,i},\varphi_{t+1}^{C,i}\right)$ and $\beta_{\varphi^{C,i}}\approx-\beta_{\varphi^{C,j}}\approx\beta_{\varphi^C}$, then equation (23) simplifies further to

$$\frac{Cov\left(\varphi_{t+1}^{MP},\Delta s_{t+1}^{i,j}\right)}{Var\left(\Delta s_{t+1}^{i,j}\right)}\approx\beta_{\varphi^C}\frac{Var\left(\varphi_{t+1}^{C}\right)+2Cov\left(\varphi_{t+1}^{F,i},\varphi_{t+1}^{C,i}\right)}{Var\left(\Delta s_{t+1}\right)}+\beta_{\varphi^P}\frac{Var\left(\varphi_{t+1}^{P}\right)}{Var\left(\Delta s_{t+1}^{i,j}\right)}.$$

Here, we see that by holding the regression coefficients $\left\{\beta_{\varphi^C},\beta_{\varphi^P}\right\}$ fixed, equation (23) implies that monetary policy explains a larger fraction of the volatility of the exchange rate change when: (i) relative monetary policy surprises, φ_{t+1}^{C}, are volatile, (ii) monetary policy is more persistent (that is, a high $Cov\left(\varphi_{t+1}^{F,i},\varphi_{t+1}^{C,i}\right)$), or (iii) relative path components, φ_{t+1}^{P}, are volatile. Furthermore, a larger fraction of the variance of the exchange rate change will be attributed to monetary policy when the regression coefficients $\left\{\beta_{\varphi^C},\beta_{\varphi^P}\right\}$ are high.

The VAR literature's impulse responses can be interpreted as one variant of β_{φ^C} with the main difference being that the explanatory variable is a monetary policy shock that is identified differently and treated as exogenous. Another difference in our approach is that we have decomposed our exchange rate changes into components with economic interpretations which allows us to study the contribution of each of these terms to the estimated coefficient. That is, we can see whether the direct effect of monetary policy surprises through the total monetary policy term is amplified or dampened by their comovement with the remaining components:[19]

$$\begin{aligned} \beta_{\varphi^x} &= \frac{Cov\left(\varphi^x_{t+1}, \Delta s^{i,j}_{t+1}\right)}{Var\left(\varphi^x_{t+1}\right)} \\ &\approx \frac{Cov\left(\varphi^x_{t+1}, -\varphi^{MP}_{t+1}\right)}{Var\left(\varphi^x_{t+1}\right)} + \frac{Cov\left(\varphi^x_{t+1}, s^{\Delta E}_{t+1,\infty}\right)}{Var\left(\varphi^x_{t+1}\right)} + \frac{Cov\left(\varphi^x_{t+1}, \sigma_t - \sigma^F_{t+1}\right)}{Var\left(\varphi^x_{t+1}\right)} \end{aligned} \tag{24}$$

for $x \in \{C, P\}$.

If conventional monetary policy surprises, as captured by φ^C_{t+1}, affect the exchange rate change only through its persistence and is uncorrelated with the expected excess returns components and changes in long-run nominal exchange rate expectations, then $\beta_{\varphi^C} \approx \frac{Cov\left(\varphi^C_{t+1}, -\varphi^{MP}_{t+1}\right)}{Var\left(\varphi^C_{t+1}\right)}$. However, comovement between φ^C_{t+1} and the remaining terms in the decomposition can play an important role in amplifying or dampening the link between monetary policy and exchange rate dynamics. Below, we will examine these comovements with respect to both our measures of conventional and unconventional monetary policy.

Both the variance decomposition of the exchange rate change, given by equation (22), and the regression coefficients β_{φ^C} and β_{φ^P} convey useful but somewhat different information. The former helps us to understand how much monetary policy has contributed to exchange rate fluctuations historically. In contrast, β_{φ^C} and β_{φ^P} can be interpreted as the change in the exchange rate that is associated with a 1 percentage point change in the relevant monetary policy term, given historical correlations between the monetary policy component in question and the remaining terms in the decomposition as shown in equation (24). The monetary policy components, φ^C and φ^P, can themselves capture a variety of structural shocks related to the real economy, the health of the financial sector, and even the nature of the central bank's policy function. In this paper, we do not attempt to identify the structural shocks that drive the relationships presented below.

4.2 Decomposition Results

In this subsection, we present the result from the decomposition given by equations (22), (23), and (24). Throughout this section, we highlight the results for the USD and GBP bases, while the results for the JPY and DEM/EUR bases are reported in Section A.3 of the Appendix. We start with the decomposition given in equation (22) where the variance of the exchange rate change

[19]The reason why the relationship holds approximately and not with an equality is because we use a constrained VAR specification.

is divided into the parts attributable to monetary policy, $\frac{Cov\left(-\varphi_{t+1}^{MP},\Delta s_{t+1}^{i,j}\right)}{Var\left(\Delta s_{t+1}^{i,j}\right)}$; uncovered interest rate parity, $\frac{Cov\left(\tilde{i}_t^{i,j},\Delta s_{t+1}^{i,j}\right)}{Var\left(\Delta s_{t+1}^{i,j}\right)}$; long-run exchange rate expectations, $\frac{Cov\left(s_{t,\infty}^{\Delta E},\Delta s_{t+1}^{i,j}\right)}{Var\left(\Delta s_{t+1}^{i,j}\right)}$; and expected excess returns, $\frac{Cov\left(\sigma_t-\sigma_{t+1}^{F},\Delta s_{t+1}^{i,j}\right)}{Var\left(\Delta s_{t+1}^{i,j}\right)}$.

From Figure 2 and Figure A1 in the Appendix, it is apparent that most of the exchange rate change variation can be explained by the expected excess return components. However, for some currency pairs, the monetary policy component plays an important role. This is the case for most GBP base currency pairs where, for GBP-USD, GBP-DEM/EUR and GBP-CHF, $\frac{Cov\left(-\varphi_{t+1}^{MP},\Delta s_{t+1}^{i,j}\right)}{Var\left(\Delta s_{t+1}^{i,j}\right)}$ equals $0.38\ (0.31)$, $0.23\ (0.61)$ and $0.21\ (0.29)$ for the *YF*(*TR*) specifications, respectively.

In order to further understand the importance of monetary policy for explaining volatility in exchange rate fluctuations, Figure 3 reports the decomposition given by equation (23). Additionally, we present this particular decomposition for two subsamples: a pre-ZLB sample of 1990:Q1-2008:Q4 and a ZLB sample of 2009:Q1 until the end of our available data.[20] This subsample exercise illustrates how the contribution of φ^C (which we interpret as conventional policy surprises) and φ^P (interpreted as an unconventional policy component) has changed since the United States hit the ZLB. For our subsample exercises, we report results from only the *YF* specification since this is the case that best captures forward guidance and the signaling channel of QE, both of which are thought to be important drivers of φ^P.[21]

When considering the GBP or the USD as base currencies, it is apparent that the overall contribution of monetary policy to explaining the variance of the exchange rate change, given by $\Pi^C + \Pi^P$, decreased over the ZLB period. Moreover, φ^C is the main driver of the contribution of monetary policy prior to the ZLB period, but the importance of φ^P has increased since then for most currency pairs. However, the same pattern does not hold when one considers the DEM/EUR as a base currency. Given the binding ZLB in the recent period, the lack of uncertainty over one-period-ahead policy rates, and the use of unconventional monetary policy in the United States and United Kingdom, the results for the GBP and USD bases are not too surprising. The lack of higher importance of φ^P in the case of the DEM/EUR base can potentially be attributed to the delayed use of QE and forward guidance by the ECB.

Equation (23) indicates that the scaled covariance between the total monetary policy term, φ^{MP}, and the nominal exchange rate change over the recent ZLB period for the GBP and USD bases could have dropped for a number of reasons. Here, we focus on two possibilities: (i) lower volatil-

[20]In our subsample exercises, we do not re-estimate the VAR due to the small sample sizes within these date ranges at the quarterly frequency. For details on the exact date ranges for which we have data, see Section A.1 in the Appendix.

[21]To be consistent with equation (23), these graphs use a φ^P that is comptued by regressing $\varphi^{F,i}$ on $\varphi^{C,i}$ and $\varphi^{F,j}$ on $\varphi^{C,j}$ separately for each subsample. However, these results are robust to using a φ^P that is computed with regressions over the full sample.

ity of the two monetary policy components, or (ii) smaller changes in exchange rates associated with monetary policy surprises (that is, smaller β_{φ^C} and β_{φ^P}).

In order to assess these two channels, we first examine the variances of both components before and after the global financial crisis. Figures 4 and 5 show that the variance of conventional monetary policy surprises, φ_{t+1}^C, dropped sharply during the crisis while variances of the unconventional policy term, $Var\left(\varphi_{t+1}^P\right)$, remained relatively unchanged.

In Figures 6 and 7, we plot β_{φ^C} and β_{φ^P} for the two subsamples. While there is heterogeneity across currency pairs, one can see that the regression coefficients for both components actually become more negative over the ZLB period for most currency pairs. In the *YF* specification, β_{φ^C} becomes more negative for 18 out of 21 currency pairs while β_{φ^P} is more negative for 13 out of 21 currency pairs. Thus, we conclude that the contribution of monetary policy terms to the variation of exchange rate changes dropped in the ZLB period, and that this finding can be attributed to a drop in the variation of conventional monetary policy surprises, φ^C.

To better understand why β_{φ^C} and β_{φ^P} are actually higher in the ZLB period, we report the decomposition of this coefficients given in equation (24).

Figures 8 and 9 reveal a few patterns. First of all, the long-run nominal exchange rate expectations component tends to comove positively with both monetary policy components, φ^C and φ^P, thus creating a dampening effect on β_{φ^C} and β_{φ^P}. This comovement could potentially capture policy responses to inflationary shocks, which generate positive correlation between relative policy surprises and inflation. Because the VAR imposes purchasing power parity in the long run, this correlation would dampen the estimated effect of φ^Con contemporaneous exchange rate changes. This pattern appears in both the pre- and post-ZLB subsamples so it is not the main contributor to the larger negative β_{φ^C} and β_{φ^P} in the recent period. Instead, the change over the two subsamples seems to be more attributable to a change in the dynamics of the expected excess return component, $\sigma_t - \sigma_{t+1}^F$. For many currency pairs, this component dampens the response of the exchange rate change to relative monetary policy surprises prior to the ZLB period while the opposite effect occurs during the ZLB period.

This asymmetry implies that prior to 2009, if the eurozone had a higher monetary policy surprise relative to the United States, the direct monetary policy effect would push the euro to appreciate against the dollar. However, this higher monetary policy surprise in the eurozone is also associated with a lower expected long-run value of the euro vis-à-vis the dollar and lower expected excess returns in the future from holding the dollar-denominated bond, lower $\sigma_{t+1}^{F,£,\$}$. These latter two factors move in the direction of depreciating the euro, leading to a small and economically insignificant β_{φ^C} prior to the Federal Reserve beginning their zero interest rate policy. In contrast, during the ZLB period, a higher monetary policy surprise in the eurozone has now become associated with a higher expected excess returns in the future from holding the dollar-denominated bond. Thus, a higher monetary policy surprise in the eurozone is now associated with a very large, negative and statistically significant appreciation of the euro. In concurrent work, Stavrakeva and Tang (2015) delve deeper into the determinants of expected excess currency returns and study how this change in the link between monetary policy and exchange rates is related to changes in the health of the financial sector.

5 Conclusion

In this paper, we examine the link between monetary policy and exchange rate changes at a quarterly frequency, focusing both on measures of conventional as well as unconventional monetary policies. We confirm the folk wisdom prevalent in financial markets that a country's currency tends to appreciate when there are higher contemporaneous and expected future policy rates in that country relative to others. Historically, the variation in contemporaneous policy rate surprises and changes in expectations of future policy can explain a sizable fraction of the exchange rate change volatility for a number of currency pairs. However, variation in expected excess returns, which captures currency risk premia in addition to wedges arising from financial frictions, accounts for the majority of exchange rate change fluctuations.

We also find that the contribution of conventional monetary policy to exchange rate volatility has decreased since the United States hit the ZLB. In the recent ZLB period, a marginal increase in the relative conventional monetary policy surprise is associated with a bigger exchange rate appreciation vis-à-vis the pre-ZLB sample period of 1990:Q1-2008:Q4. These two facts can be reconciled by noting the lower volatility of monetary policy surprises over the recent period. We found similar results for unconventional monetary policy.

Lastly, we investigate how the comovements between relative monetary policy surprises and other components of exchange rate changes have evolved between the pre- and post-ZLB periods. In the pre-ZLB period, we find that a higher monetary policy surprise in country i had been associated with higher expected excess returns from holding the currency of country j. However, in the more recent period, this relationship has reversed. Thus, the dynamics of expected excess returns were previously dampening the effect of monetary policy prior to the ZLB period, but now play an amplifying role instead.

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Table 1: Change in Relative Short-Term Yields, 1990:Q1-2015:Q1

\$ base currency	**AUD**	**CAD**	**CHF**	**DEM/EUR**	**JPY**	**NOK**	**NZD**	**SEK**	**GBP**
$\Delta\tilde{i}_{t+1}^{i,\$}$	-8.617^{***}	1.624	-10.244	-4.224	-7.701	-6.319^{**}	-4.818^{*}	-8.136^{***}	-15.795^{***}
	[2.822]	[1.866]	[6.336]	[4.862]	[6.851]	[3.103]	[2.784]	[2.667]	[4.869]
Constant	-0.323	-0.044	-0.663	-0.234	-0.206	0.029	-0.123	0.007	-0.300
	[0.581]	[0.365]	[0.563]	[0.551]	[0.648]	[0.600]	[0.537]	[0.543]	[0.434]
# Observations	99	99	90	85	88	77	103	99	99
Adjusted R^2	0.041	-0.006	0.050	-0.000	0.008	0.040	0.016	0.220	0.199
RMSE	5.658	3.582	5.385	4.993	6.059	5.233	5.436	5.389	4.376

£ base currency	**AUD**	**CAD**	**CHF**	**DEM/EUR**	**JPY**	**NOK**	**NZD**	**SEK**	**USD**
$\Delta\tilde{i}_{t+1}^{i,£}$	-17.946^{***}	-5.131^{*}	-8.191	-11.797	-32.252^{***}	-5.854^{**}	-10.097^{***}	-5.245^{***}	-15.795^{***}
	[2.356]	[2.813]	[8.052]	[7.280]	[8.399]	[2.516]	[2.376]	[1.067]	[4.869]
Constant	-0.160	0.018	-0.550	-0.125	0.108	0.107	-0.141	0.168	0.300
	[0.510]	[0.465]	[0.535]	[0.466]	[0.647]	[0.466]	[0.515]	[0.415]	[0.434]
# Observations	99	99	90	85	88	77	99	99	99
Adjusted R^2	0.228	0.030	0.041	0.055	0.309	0.044	0.106	0.119	0.199
RMSE	5.072	4.643	5.200	4.138	6.040	4.065	5.143	4.127	4.376

¥ base currency	**AUD**	**CAD**	**CHF**	**DEM/EUR**	**NOK**	**NZD**	**SEK**	**GBP**	**USD**
$\Delta\tilde{i}_{t+1}^{i,¥}$	-19.095^{**}	-3.303	-11.955^{*}	-14.642^{*}	-14.263^{*}	-9.284	-11.880^{***}	-32.252^{***}	-7.701
	[8.930]	[7.490]	[6.844]	[8.793]	[8.451]	[6.513]	[2.977]	[8.399]	[6.851]
Constant	-0.086	0.005	-0.313	-0.297	-0.380	-0.213	0.003	-0.108	0.206
	[0.832]	[0.762]	[0.619]	[0.654]	[0.812]	[0.789]	[0.778]	[0.647]	[0.648]
# Observations	88	88	88	85	77	88	88	88	88
Adjusted R^2	0.079	-0.007	0.036	0.037	0.087	0.039	0.141	0.309	0.008
RMSE	7.786	7.148	5.824	6.070	7.341	7.343	7.260	6.040	6.059

DEM/€ base currency	**AUD**	**CAD**	**CHF**	**JPY**	**NOK**	**NZD**	**SEK**	**GBP**	**USD**
$\Delta\tilde{i}_{t+1}^{i,DEM/e}$	-10.974^{**}	6.772	0.595	-14.642^{*}	0.001	-8.223^{**}	0.255	-11.797	-4.224
	[5.191]	[5.251]	[3.291]	[8.793]	[3.124]	[3.921]	[4.723]	[7.280]	[4.862]
Constant	-0.061	-0.073	-0.425^{*}	0.297	-0.072	-0.147	0.045	0.125	0.234
	[0.633]	[0.563]	[0.244]	[0.654]	[0.404]	[0.571]	[0.370]	[0.466]	[0.551]
# Observations	85	85	85	85	77	85	85	85	85
Adjusted R^2	0.033	0.025	-0.011	0.037	-0.013	0.052	-0.012	0.055	-0.000
RMSE	5.845	5.178	2.270	6.070	3.547	5.227	3.403	4.138	4.993

Frequency: Quarterly
Dependent variable: $\Delta s_{t+1}^{i,j}$ Quarterly log exchange rate change (increase implies base currency appreciation)
Independent variable: Change in relative 3-month yields (not annualized).
Heteroskedasticity-robust standard errors are in brackets.
$^{***}/^{**}/^{*}$ Statistically significant at 1, 5, and 10 percent, respectively.

Table 2: Change in Relative Long-Term Yields, 1990:Q1-2014:Q3

\$ base currency	**AUD**	**CAD**	**CHF**	**DEM/EUR**	**JPY**	**NOK**	**NZD**	**SEK**	**GBP**
$\Delta n\tilde{y}_{t+1}^{20,i,\$}$	-0.623^{*}	0.171	-1.134^{***}	-1.328^{***}	-0.865^{***}	-0.242	-0.090	-0.413^{**}	-0.836^{***}
	[0.327]	[0.170]	[0.227]	[0.219]	[0.205]	[0.219]	[0.191]	[0.192]	[0.239]
Constant	-0.289	-0.042	-0.509	-0.246	-0.333	-0.302	-0.307	-0.134	-0.159
	[0.574]	[0.365]	[0.509]	[0.480]	[0.566]	[0.678]	[0.557]	[0.568]	[0.448]
# Observations	99	99	99	99	99	66	98	87	99
Adjusted R^2	0.041	-0.003	0.226	0.241	0.130	0.008	-0.009	0.026	0.135
RMSE	5.659	3.578	5.060	4.767	5.638	5.553	5.501	5.240	4.547

£ base currency	**AUD**	**CAD**	**CHF**	**DEM/EUR**	**JPY**	**NOK**	**NZD**	**SEK**	**USD**
$\Delta n\tilde{y}_{t+1}^{20,i,£}$	-1.114^{***}	-1.026^{***}	-0.582^{*}	-0.268	-1.043^{**}	-0.127	-0.315	-0.424	-0.836^{***}
	[0.188]	[0.273]	[0.350]	[0.254]	[0.412]	[0.164]	[0.261]	[0.308]	[0.239]
Constant	-0.190	0.024	-0.399	-0.088	-0.143	-0.299	-0.236	0.021	0.159
	[0.515]	[0.434]	[0.483]	[0.431]	[0.661]	[0.483]	[0.549]	[0.456]	[0.448]
# Observations	99	99	99	99	99	66	98	87	99
Adjusted R^2	0.211	0.153	0.047	0.000	0.121	-0.005	0.006	0.022	0.135
RMSE	5.126	4.338	4.968	4.283	6.718	3.966	5.431	4.265	4.547

¥ base currency	**AUD**	**CAD**	**CHF**	**DEM/EUR**	**NOK**	**NZD**	**SEK**	**GBP**	**USD**
$\Delta n\tilde{y}_{t+1}^{20,i,¥}$	-1.103^{***}	-0.899^{**}	-0.651^{**}	-0.961^{***}	-0.400	-0.806^{***}	-0.572	-1.043^{**}	-0.865^{***}
	[0.293]	[0.354]	[0.311]	[0.327]	[0.371]	[0.292]	[0.505]	[0.412]	[0.205]
Constant	-0.056	0.184	-0.184	0.113	0.027	0.113	0.053	0.143	0.333
	[0.730]	[0.664]	[0.583]	[0.588]	[0.917]	[0.715]	[0.756]	[0.661]	[0.566]
# Observations	99	99	99	99	66	98	87	99	99
Adjusted R^2	0.159	0.089	0.051	0.109	0.014	0.073	0.032	0.121	0.130
RMSE	7.317	6.677	5.787	5.918	7.597	7.095	7.303	6.718	5.638

DEM/€ base currency	**AUD**	**CAD**	**CHF**	**JPY**	**NOK**	**NZD**	**SEK**	**GBP**	**USD**
$\Delta n\tilde{y}_{t+1}^{20,i,DEM/e}$	-1.069^{***}	-0.688^{**}	-0.236	-0.961^{***}	-0.086	-0.668^{**}	0.681^{**}	-0.268	-1.328^{***}
	[0.264]	[0.322]	[0.235]	[0.327]	[0.125]	[0.323]	[0.272]	[0.254]	[0.219]
Constant	-0.152	0.071	-0.365	-0.113	0.048	-0.124	0.151	0.088	0.246
	[0.603]	[0.560]	[0.229]	[0.588]	[0.451]	[0.587]	[0.363]	[0.431]	[0.480]
# Observations	99	99	99	99	66	98	87	99	99
Adjusted R^2	0.151	0.047	0.006	0.109	-0.010	0.042	0.099	0.000	0.241
RMSE	5.908	5.556	2.313	5.918	3.634	5.776	3.378	4.283	4.767

Frequency: Quarterly; $n = 20$
Dependent variable: $\Delta s_{t+1}^{i,j}$ Quarterly log exchange rate change (increase implies base currency appreciation).
Independent variable: Change in relative 5-year quarterly yields (not annualized).
Heteroskedasticity-robust standard errors are in brackets.
$^{***}/^{**}/^{*}$ Statistically significant at 1, 5, and 10 percent, respectively.

Table 3: Correlation between VAR-Implied and Market-Based Expected Monetary Policy Changes

	AUD	CAD	CHF	DEM/EUR	NOK	NZD	SED	GBP	USD
Taylor Rule	0.44	0.21	0.51	0.69	-0.25	0.58	0.10	0.62	0.62
Yield Factor	0.72	0.39	0.62	0.74	0.21	0.74	0.66	0.78	0.75
# Observations	104	102	97	87	79	102	102	109	114

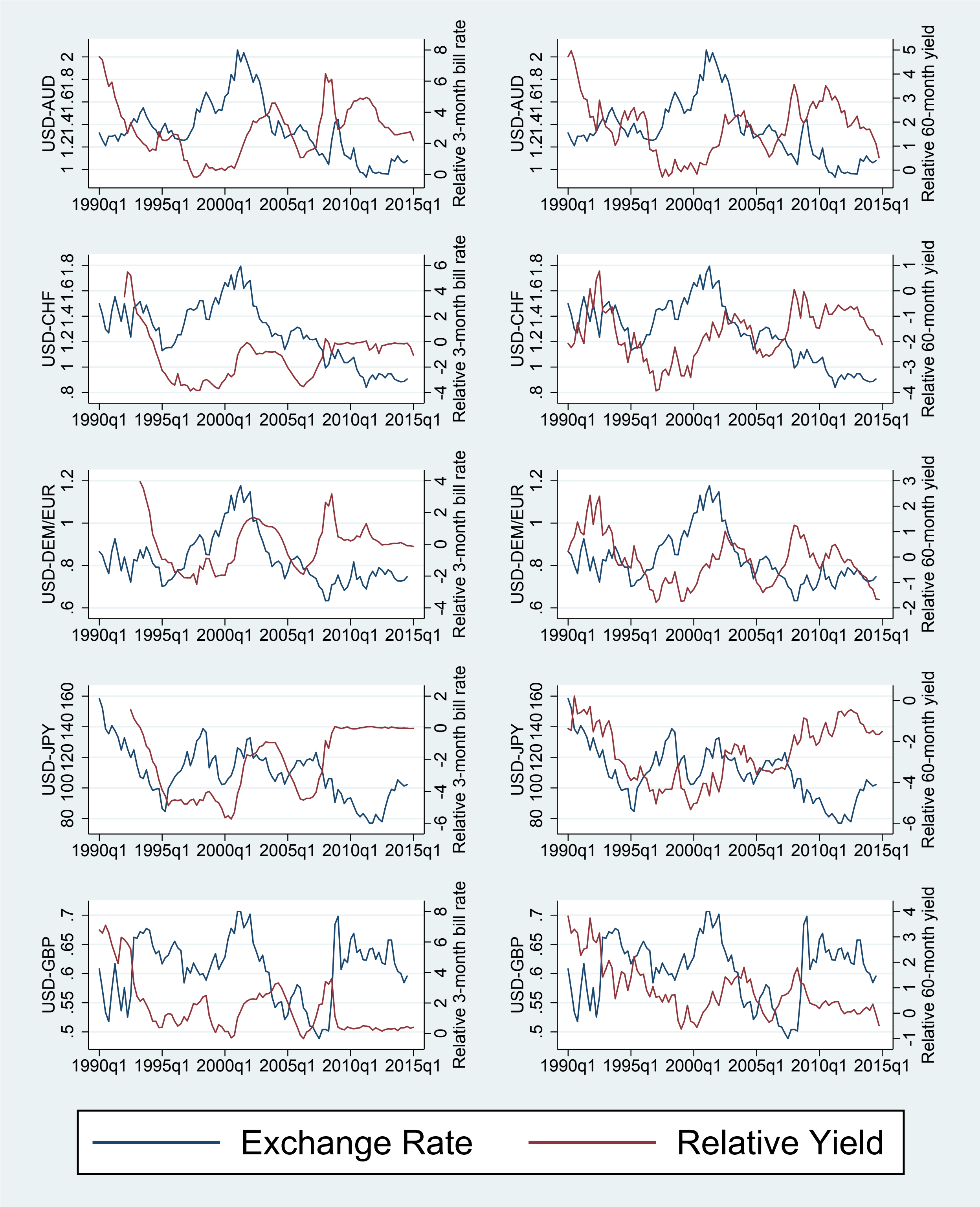
Frequency: Quarterly

Table 4: Correlation between VAR-Implied and Market-Based Expected Monetary Policy Surprises

	AUD	CAD	CHF	DEM/EUR	NOK	NZD	SED	GBP	USD
Taylor Rule	0.62	0.56	0.37	0.81	0.05	0.74	0.60	0.58	0.69
Yield Factor	0.83	0.65	0.49	0.84	-0.11	0.78	0.79	0.69	0.77
# Observations	104	103	97	88	80	103	103	109	114

Frequency: Quarterly

Figure 1: Exchange Rates versus Short and Long-term Yields Relative to the U.S. Dollar, 1990:Q1-2015:Q1

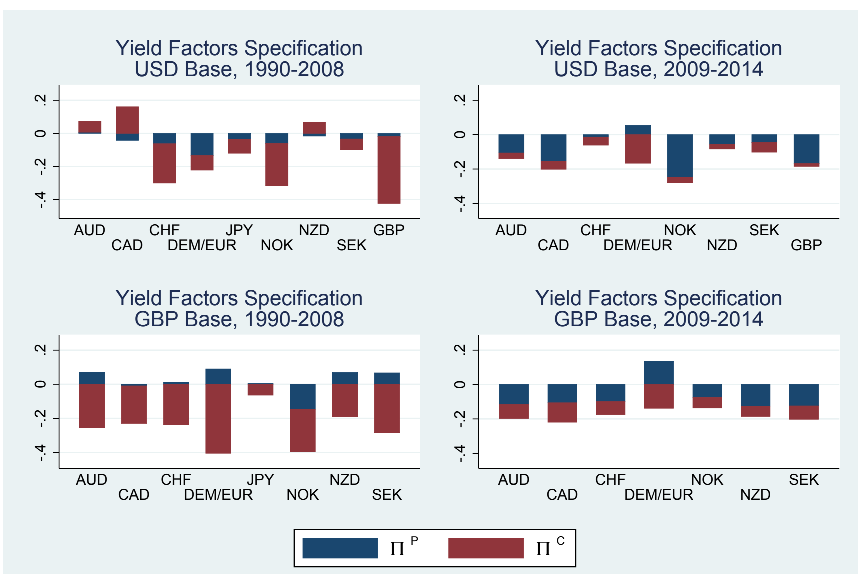


Sources: Global Financial Data, Wright (2011), Gürkaynak, Sack, and Wright (2007), Reserve Bank of Australia, Swiss National Bank, Bundesbank, Bank of England, and authors' calculations.

Figure 2: Fraction of $Var\left(\Delta s_{t+1}^{i,j}\right)$ Attributed to Different Components

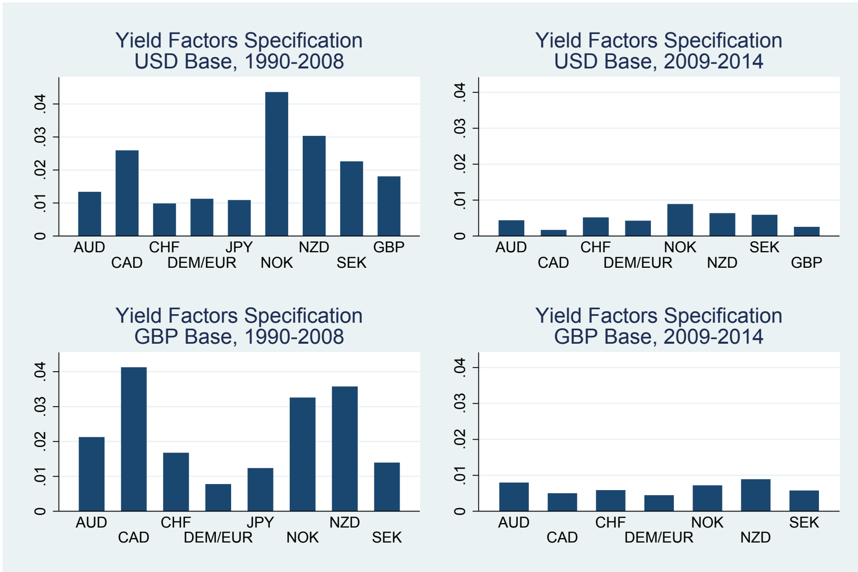
Source: Authors' calculations.

Figure 3: Decomposing $\frac{Cov\left(\varphi_{t+1}^{MP},\Delta s_{t+1}^{i,j}\right)}{Var\left(\Delta s_{t+1}^{i,j}\right)}$



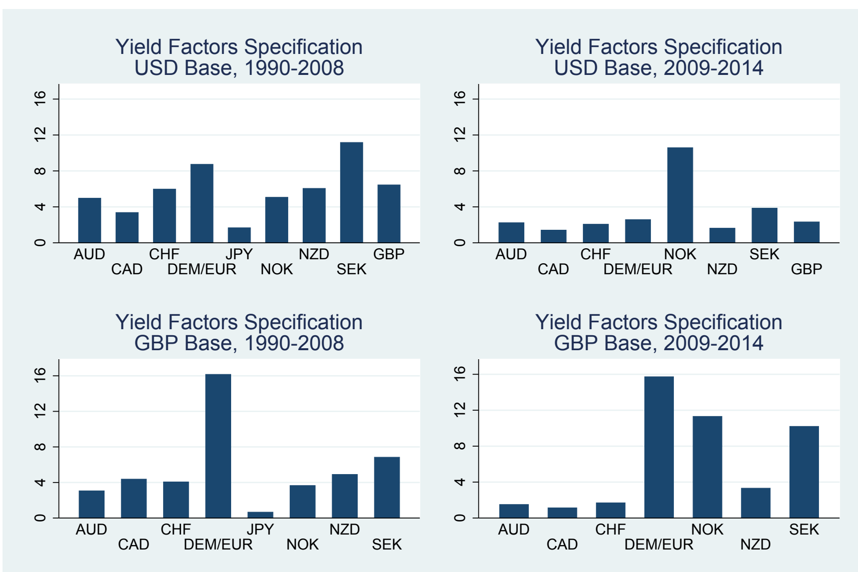
Source: Authors' calculations.

Figure 4: Variance of Conventional Relative Monetary Policy Surprises, φ^C_{t+1}



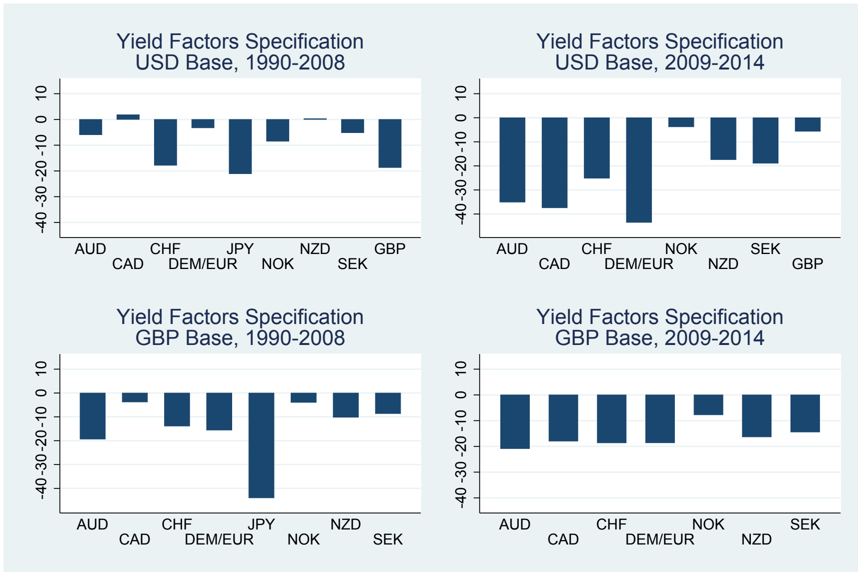
Source: Authors' calculations.

Figure 5: Variance of Unconventional Monetary Policy Terms, φ^P_{t+1}



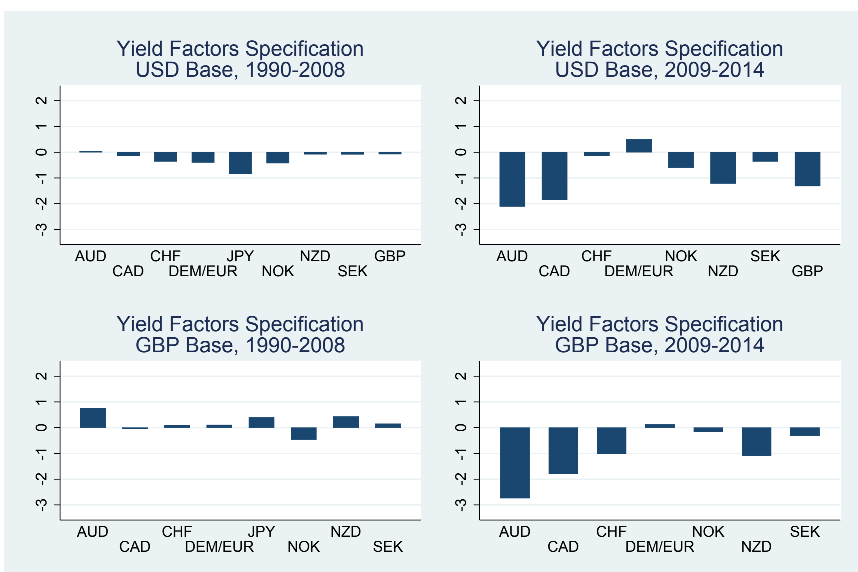
Source: Authors' calculations.

Figure 6: Coefficients from Regressing Exchange Rate Changes on Conventional Relative Monetary Policy Surprises, β_{φ^C}



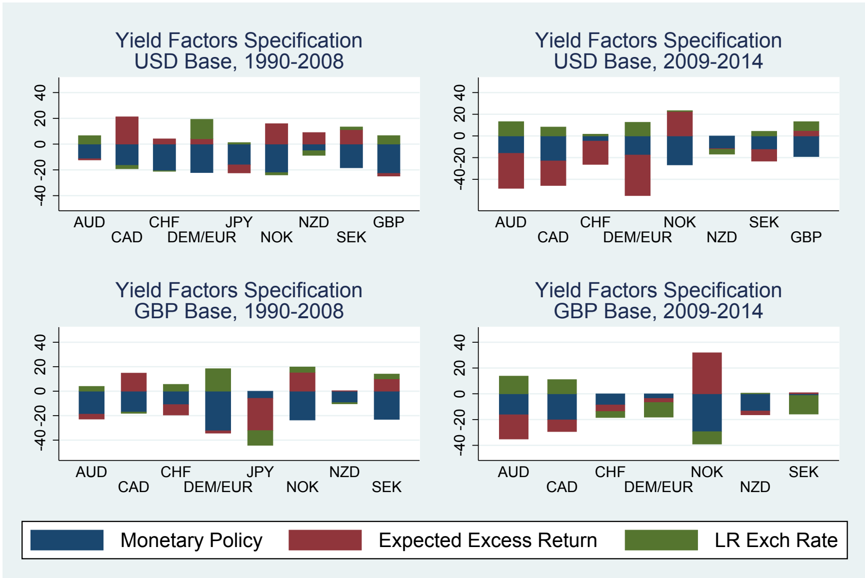
Source: Authors' calculations.

Figure 7: Coefficients from Regressing Exchange Rate Changes on Unconventional Monetary Policy Terms, β_{φ^P}



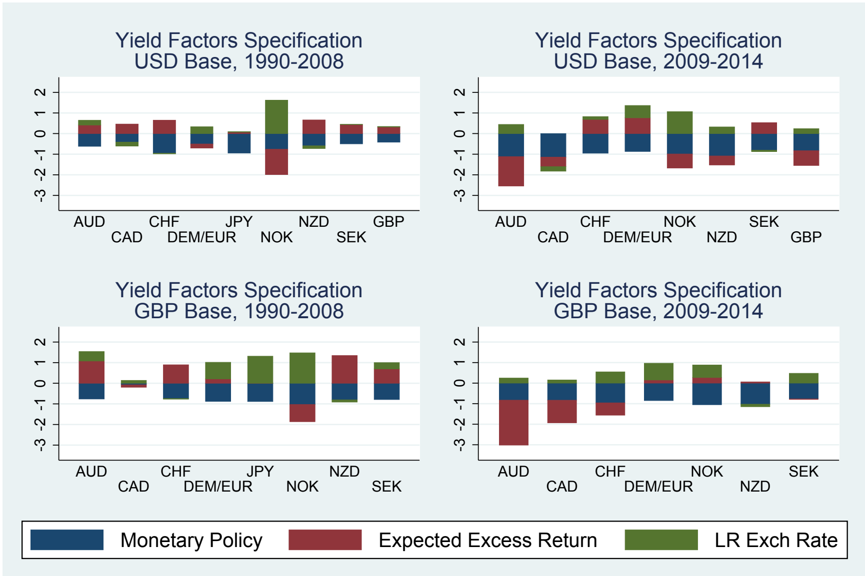
Source: Authors' calculations.

Figure 8: Decomposition of Regression Coefficients β_{φ^C} Into Parts Associated with Different Components of Exchange Rates



Source: Authors' calculations.

Figure 9: Decomposition of Regression Coefficients β_{φ^P} Into Parts Associated with Different Components of Exchange Rates



Source: Authors' calculations.

A Appendix

A.1 Data Description

- *Exchange Rates*: End-of-quarter exchange rates are obtained using daily data from Global Financial Data.
- *Short-term rates*: End-of-quarter 3-month bill rates were 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 3-month bill rates are replaced with 3-month EONIA OIS swap rates starting in 1999:Q1.
  - Japan: Bloomberg
- *Zero-Coupon Yields:* End-of-quarter zero-coupon yields were obtained from the following sources:
  - Canada, Germany, Sweden, Switzerland, and United Kingdom: Central banks
  - 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: Wright (2011)
  - United States: Gürkaynak, Sack, and Wright (2007)
- *GDP Deflator, Output Gap, and Current Account-to-GDP ratio:* All macro data is from the OECD Main Economic Indicators and Economic Outlook databases. The GDP gap is computed using the OECD's annual estimates of potential GDP which were log-linearly interpolated to the quarterly frequency. German data are replaced with euro-area data starting in 1999:Q1.
- *Market-Based Interest Rate Surprises and Expected Changes:* These are computed using prices of futures on 3-month interest rates on the last trading day of each quarter. These expectations refer to the 3-month rates on each contract's last trading day which typically falls within the 2nd to last week of each quarter. When computing the surprises and expected changes in these interest rates, the actual rate used is the underlying rate of each futures contract. The futures data are all obtained from Bloomberg and are based on the following underlying rates:

- Australia: Australian 90-day bank accepted bills
- Canada: Canadian 3-month bankers' acceptance
- Switzerland: 3-month Euroswiss
- Germany/EU: ICE 3-month Euribor
- Norway: 3-month NIBOR
- New Zealand: New Zealand 90-day bank accepted bills
- Sweden: 3-month Swedish T-bill (1992:Q4-2007:Q4); 3-month STIBOR (2008:Q1-present)
- United Kingdom: 3-month Sterling LIBOR
- United States: 3-month Eurodollar

Table A1: Data Sample Ranges

	Exchange rates, short-term yields, and macroeconomic variables	Zero-coupon long-term yields
Australia	1990:Q1 – 2014:Q3	1990:Q1 – 2014:Q3
Canada	1990:Q1 – 2014:Q3	1990:Q1 – 2014:Q3
Germany	1993:Q2 – 2014:Q3	1993:Q2 – 2014:Q3
Japan	1992:Q3 – 2014:Q3	1992:Q3 – 2009:Q2
New Zealand	1990:Q1 – 2015:Q1	1990:Q1 – 2014:Q3
Norway	1995:Q2 – 2014:Q3	1998:Q1 – 2014:Q3
Sweden	1990:Q1 – 2014:Q3	1992:Q4 – 2014:Q3
Switzerland	1992:Q1 – 2014:Q3	1992:Q1 – 2014:Q3
United Kingdom	1990:Q1 – 2014:Q3	1990:Q1 – 2014:Q3
United States	1990:Q1 – 2015:Q1	1990:Q1 – 2015:Q1

A.2 Additional Tables

Table A2: Monetary Policy Persistence, 1990:Q1-2015:Q2

	AUD	CAD	CHF	DEM/EUR	JPY	NOK	NZD	SEK	GBP	USD
Δi_t^i	0.884***	0.941***	0.924***	0.950***	0.891***	0.960***	0.931***	0.936***	0.955***	0.969***
	[0.022]	[0.026]	[0.044]	[0.020]	[0.031]	[0.029]	[0.028]	[0.075]	[0.015]	[0.015]
Constant	0.132***	0.029	0.006	0.012	0.003	0.023	0.083**	0.040	0.020	0.005
	[0.031]	[0.021]	[0.014]	[0.011]	[0.003]	[0.019]	[0.038]	[0.059]	[0.016]	[0.010]
# Observations	103	103	94	89	92	81	103	103	101	103
Adjusted R^2	0.947	0.946	0.938	0.946	0.959	0.905	0.922	0.898	0.975	0.961
RMSE	0.138	0.166	0.116	0.099	0.039	0.165	0.190	0.312	0.131	0.112

Frequency: Quarterly
Dependent variable: i_{t+1}^i 3-month quarterly yields
Independent variable: Lagged 3-month quarterly yields
Heteroskedasticity-robust standard errors are in brackets.
$^{***}/^{**}/^{*}$ Statistically significant at 1, 5, and 10 percent, respectively.

Table A3: Uncovered Interest Rate Parity, 1990:Q1-2015:Q1

\$ base currency	**AUD**	**CAD**	**CHF**	**DEM/EUR**	**JPY**	**NOK**	**NZD**	**SEK**	**GBP**
$\tilde{i}_t^{i,\$}$	-0.398	0.276	-1.446	-2.044	-1.422	-0.213	0.909	1.584	1.156
	[1.399]	[0.537]	[1.899]	[1.520]	[1.231]	[1.656]	[1.606]	[1.597]	[1.494]
Constant	0.125	-0.121	-0.974	-0.202	-0.976	0.069	-0.865	-0.520	-0.564
	[1.054]	[0.410]	[0.929]	[0.542]	[0.902]	[0.604]	[1.487]	[0.691]	[0.587]
# Observations	99	99	90	85	88	77	103	99	99
Adjusted R^2	-0.009	-0.009	0.005	0.010	0.003	-0.013	-0.006	0.029	0.002
RMSE	5.805	3.588	5.511	4.967	6.075	5.376	5.497	6.014	4.884

£ base currency	**AUD**	**CAD**	**CHF**	**DEM/EUR**	**JPY**	**NOK**	**NZD**	**SEK**	**USD**
$\tilde{i}_t^{i,£}$	-0.640	0.630	-0.962	-1.369	0.571	-0.806	0.134	0.468	1.156
	[1.626]	[1.385]	[1.448]	[1.438]	[1.399]	[1.070]	[1.826]	[1.267]	[1.494]
Constant	0.053	0.125	-1.202	-0.456	0.250	0.088	-0.288	0.172	0.564
	[0.797]	[0.506]	[0.950]	[0.679]	[1.344]	[0.476]	[1.051]	[0.469]	[0.587]
# Observations	99	99	90	85	88	77	99	99	99
Adjusted R^2	-0.008	-0.009	-0.006	-0.001	-0.010	-0.005	-0.010	-0.007	0.002
RMSE	5.794	4.734	5.325	4.258	7.304	4.170	5.468	4.412	4.884

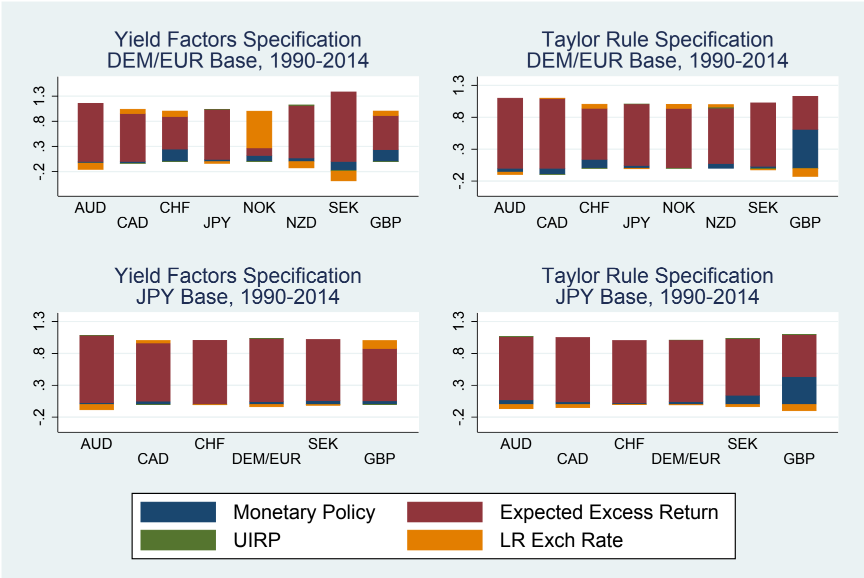
¥ base currency	**AUD**	**CAD**	**CHF**	**DEM/EUR**	**NOK**	**NZD**	**SEK**	**GBP**	**USD**
$\tilde{i}_t^{i,¥}$	1.494	-0.427	1.522	0.699	1.755	1.296	2.952	0.571	-1.422
	[2.817]	[1.941]	[2.961]	[2.217]	[1.752]	[1.431]	[2.268]	[1.399]	[1.231]
Constant	-1.907	0.325	-0.537	-0.529	-1.845	-2.021	-1.881	-0.250	0.976
	[3.196]	[1.520]	[0.891]	[1.468]	[1.593]	[1.957]	[1.658]	[1.344]	[0.902]
# Observations	88	88	88	85	77	88	88	88	88
Adjusted R^2	-0.008	-0.011	-0.007	-0.011	0.000	-0.003	0.035	-0.010	0.003
RMSE	8.145	7.162	5.951	6.217	7.683	7.501	7.694	7.304	6.075

DEM/€ base currency	**AUD**	**CAD**	**CHF**	**JPY**	**NOK**	**NZD**	**SEK**	**GBP**	**USD**
$\tilde{i}_t^{i,DEM/e}$	-1.153	-2.759	-0.746	0.699	-0.386	2.919^{*}	-1.408	-1.369	-2.044
	[2.380]	[2.498]	[2.075]	[2.217]	[1.299]	[1.486]	[1.531]	[1.438]	[1.520]
Constant	0.539	0.288	-0.685	0.529	0.068	-2.632^{**}	0.283	0.456	0.202
	[1.796]	[0.670]	[0.762]	[1.468]	[0.597]	[1.276]	[0.377]	[0.679]	[0.542]
# Observations	85	85	85	85	77	85	85	85	85
Adjusted R^2	-0.008	0.007	-0.009	-0.011	-0.012	0.034	0.005	-0.001	0.010
RMSE	5.969	5.225	2.268	6.217	3.545	5.277	3.374	4.258	4.967

Frequency: Quarterly
Dependent variable: $\Delta s_{t+1}^{i,j}$ Quarterly log exchange rate change (increase implies base currency appreciation).
Independent variable: Lagged relative 3-month quarterly yields
Heteroskedasticity-robust standard errors are in brackets.
$^{***}/^{**}/^{*}$ Statistically significant at 1, 5, and 10 percent, respectively.

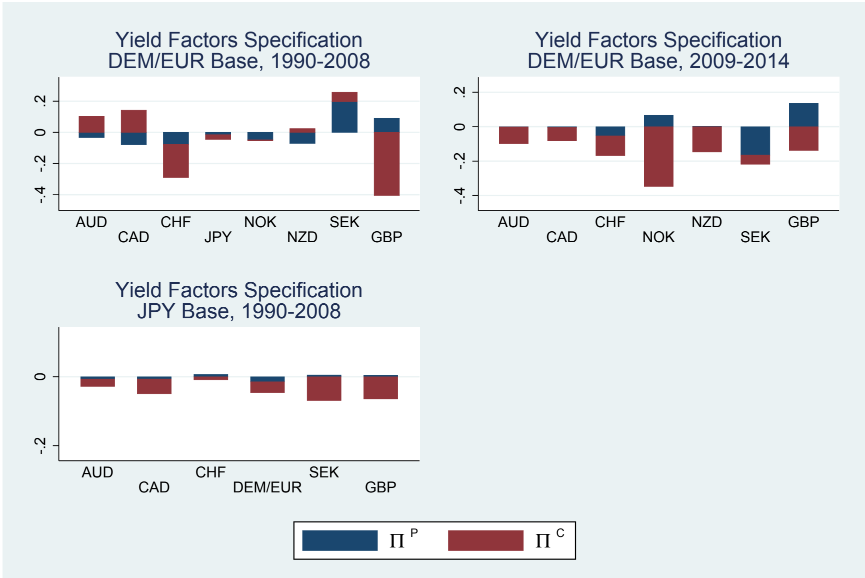
A.3 Additional Figures

Figure A1: Fraction of $Var\left(\Delta s_{t+1}^{i,j}\right)$ Attributed to Different Components



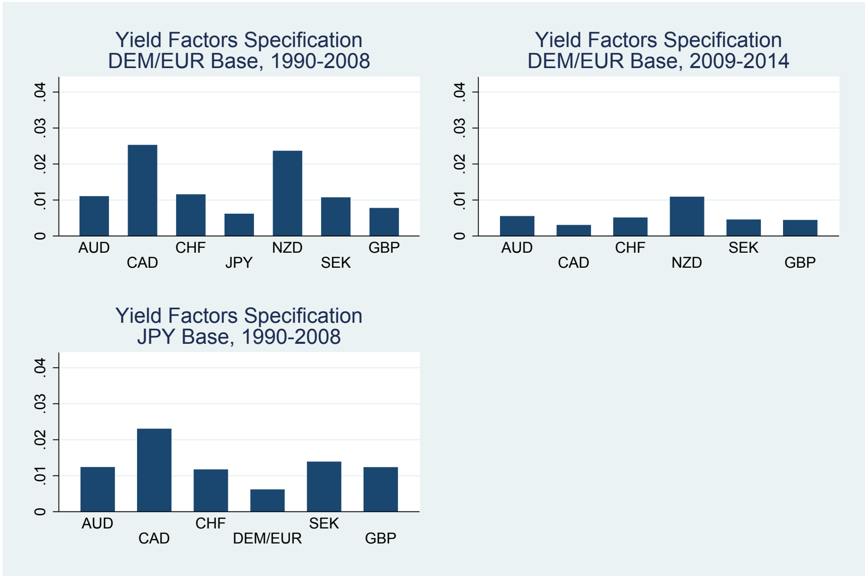
Source: Authors' calculations.

Figure A2: Decomposing $\frac{Cov\left(\varphi^{MP}_{t+1},\Delta s^{i,j}_{t+1}\right)}{Var\left(\Delta s^{i,j}_{t+1}\right)}$



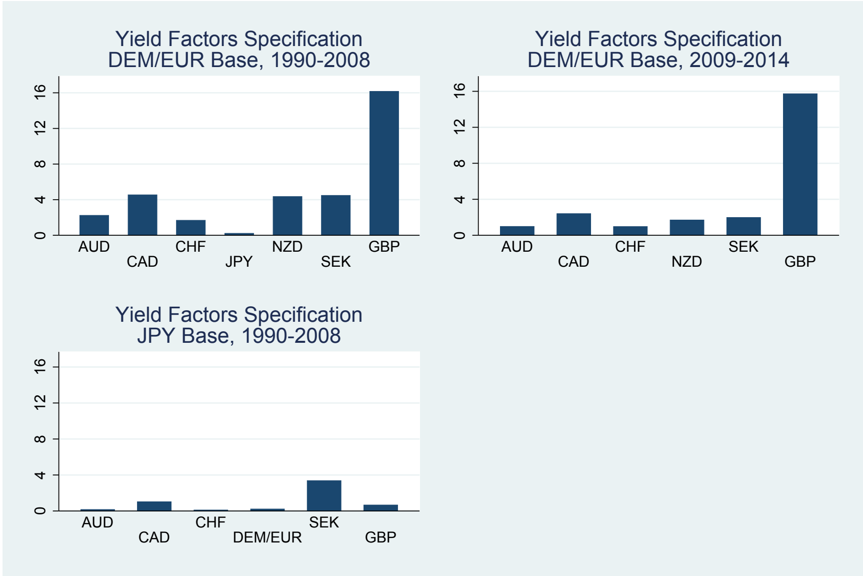
Source: Authors' calculations.

Figure A3: Variance of Conventional Relative Monetary Policy Surprises, φ^{C}_{t+1}



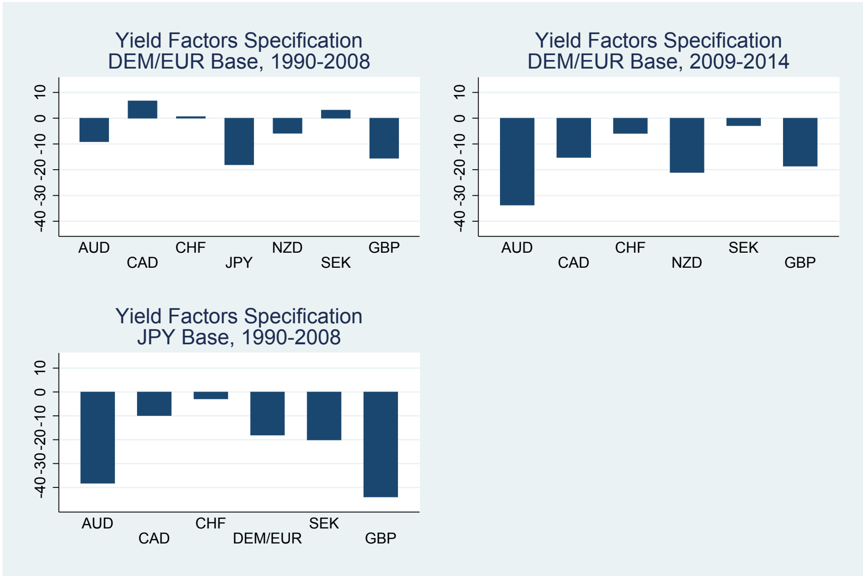
Source: Authors' calculations.

Figure A4: Variance of Unconventional Monetary Policy Terms, φ^P_{t+1}



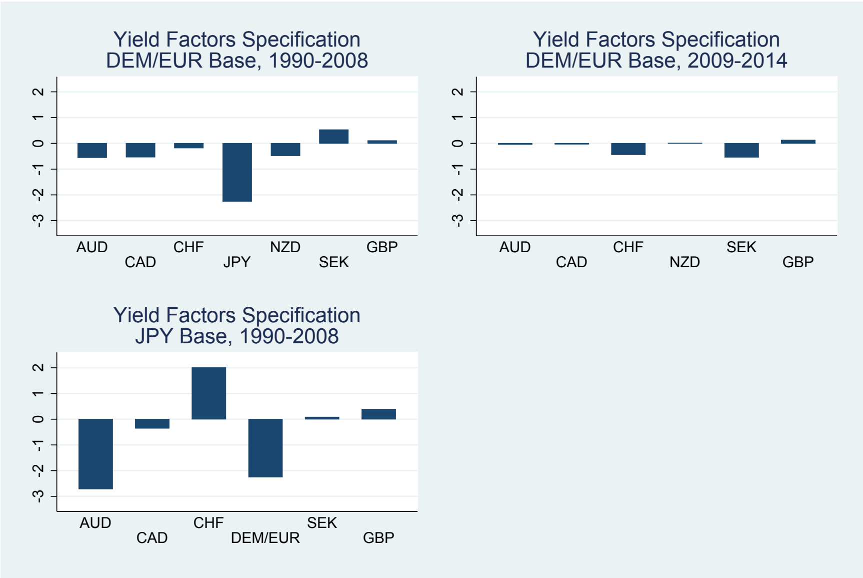
Source: Authors' calculations.

Figure A5: Coefficients from Regressing Exchange Rate Changes on Unconventional Monetary Policy Terms, β_{φ^C}



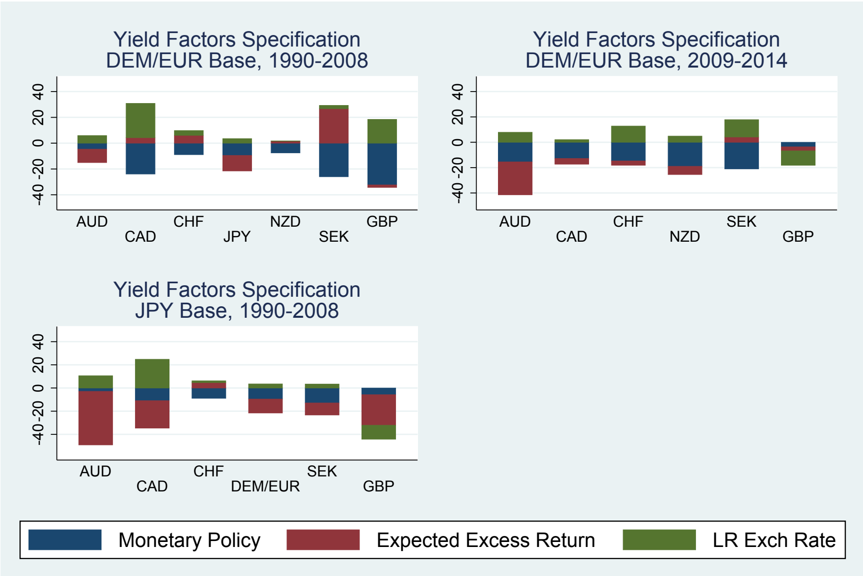
Source: Authors' calculations.

Figure A6: Coefficients from Regressing Exchange Rate Changes on Unconventional Monetary Policy Terms, β_{φ^P}



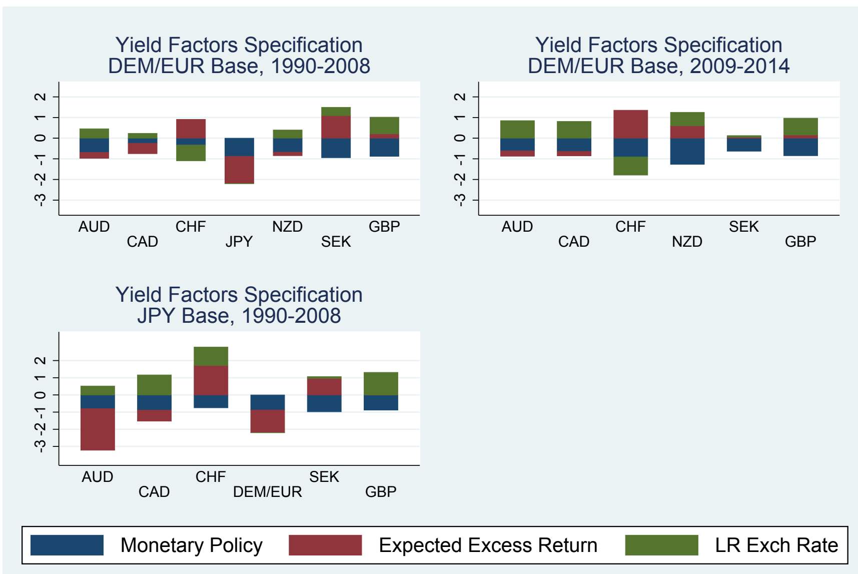
Source: Authors' calculations.

Figure A7: Decomposition of Regression Coefficients β_{φ^C} Into Parts Associated with Different Components of Exchange Rates



Source: Authors' calculations.

Figure A8: Decomposition of Regression Coefficients β_{φ^P} Into Parts Associated with Different Components of Exchange Rates



Source: Authors' calculations.

A.4 Selected Derivations

A.4.1 AR(1) Example

$$E_{t+1}\tilde{i}^{i,j}_{t+k+1} - E_t\tilde{i}^{i,j}_{t+k} = \rho_i^k \Delta i^i_{t+1} - \rho_j^k \Delta i^j_{t+1} \text{ for } k \geq 0.$$

Thus, we have

$$\begin{aligned}
\Delta s^{i,j}_{t+1} &= -\left(E_{t+1}\sum_{k=0}^{\infty}\tilde{i}^{i,j}_{t+1+k} - E_t\sum_{k=0}^{\infty}\tilde{i}^{i,j}_{t+k}\right) - \sigma^F_{t+1} + s^{\Delta E}_{t+1,\infty} + \sigma^{i,j}_t \\
&= -\frac{1}{1-\rho_i}\Delta i^i_{t+1} + \frac{1}{1-\rho_j}\Delta i^j_{t+1} - \sigma^F_{t+1} + s^{\Delta E}_{t+1,\infty} + \sigma^{i,j}_t \\
&= i^i_t - i^j_t - \left(\varepsilon^i_{t+1} - \varepsilon^j_{t+1}\right) - \left(\frac{\rho_i}{1-\rho_i}\varepsilon^i_{t+1} - \frac{\rho_j}{1-\rho_j}\varepsilon^j_{t+1}\right) - \sigma^F_{t+1} + s^{\Delta E}_{t+1,\infty} + \sigma^{i,j}_t.
\end{aligned}$$

A.4.2 Long-term Yields

The no-arbitrage condition for holding an n-period bond for k periods gives the following:

$$P^{n,i}_t = E_t\left[M^i_{t,t+k}P^{n-k,i}_{t+k}\right] \text{ where } M^i_{t,t+k} \equiv \Pi^k_{l=1}M^i_{t+l}.$$

Assuming log-normality and using the definition $y^{n,i}_t \equiv -\frac{1}{n}\ln\left(P^{n,i}_t\right)$ gives

$$\begin{aligned}
p^{n,i}_t &= -ny^{n,i}_t = E_t m^i_{t,t+n} + \frac{1}{2}Var_t\left(m^i_{t,t+n}\right) \\
&= E_t\sum_{k=1}^{n}\left(E_{t+k-1}m^i_{t+k} + \frac{1}{2}Var_{t+k-1}\left(m^i_{t+k}\right)\right) - E_t\sum_{k=1}^{n}\frac{1}{2}Var_{t+k-1}\left(m^i_{t+k}\right) + \frac{1}{2}Var_t\left(m^i_{t,t+n}\right) \\
&= -E_t\sum_{k=0}^{n-1}i^i_{t+k} - E_t\sum_{k=1}^{n}\frac{1}{2}Var_{t+k-1}\left(m^i_{t+k}\right) + \frac{1}{2}Var_t\left(m^i_{t,t+n}\right) \\
y^{n,i}_t &= \frac{1}{n}E_t\sum_{k=1}^{n}i^i_{t+k-1} + \underbrace{\frac{1}{2n}\left[E_t\sum_{k=1}^{n}Var_{t+k-1}\left(m^i_{t+k}\right) - Var_t\left(m^i_{t,t+n}\right)\right]}_{tp^{n,i}_t},
\end{aligned}$$

where we used the short-term bond pricing equation where $n = 1$,

$$i^i_{t+1} = -E_t m^i_{t+1} - \frac{1}{2}Var_t\left(m^i_{t+1}\right).$$

We can then use this expression for the n-period yield to obtain equation (14):

$$
\begin{aligned}
\Delta s_{t+1}^{i,j} &= -n\left(\frac{1}{n}E_{t+1}\sum_{k=0}^{n-1}\tilde{i}_{t+1+k}^{i,j} - \frac{1}{n}E_t\sum_{k=0}^{n-1}\tilde{i}_{t+k}^{i,j}\right) - \\
&\quad \left(E_{t+1}\sum_{k=n}^{\infty}\tilde{i}_{t+1+k}^{i,j} - E_t\sum_{k=n}^{\infty}\tilde{i}_{t+k}^{i,j}\right) - \sigma_{t+1}^{F} + s_{t+1,\infty}^{\Delta E} + \sigma_t^{i,j} \\
&= -n\left(\Delta\tilde{y}_{t+1}^{n,i,j} - \Delta\widetilde{tp}_{t+1}^{n,i,j}\right) - \left(E_{t+1}\sum_{k=n}^{\infty}\tilde{i}_{t+1+k}^{i,j} - E_t\sum_{k=n}^{\infty}\tilde{i}_{t+k}^{i,j}\right) - \sigma_{t+1}^{F} + s_{t+1,\infty}^{\Delta E} + \sigma_t^{i,j}.
\end{aligned}
$$