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The Impacts of Global Risk and US Monetary Policy on US Dollar Exchange Rates and Excess Currency Returns

Kerstin Bernoth* Helmut Herwartz† Lasse Trienens‡

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Abstract

We examine the causal relationship between US monetary policy shocks, exchange rates and currency excess returns for a sample of eight advanced countries over the period 1980M1 to 2022M11. We find that the dynamics of the US dollar exchange rate is the main driver of currency excess returns. The exchange rate is significantly affected by US monetary policy shocks, where the persistence of this shock is important, as well as by an external shock. This external shock is strongly related to global risk aversion and the convenience yield that investors are willing to pay for holding US Dollar assets. A significant part of the response of excess currency returns is also expected, suggesting a violation of the UIP. Focusing only on the post-crisis period, the impact of both the external shock and the inflation targeting shock on exchange rates and currency excess returns disappears in the cross-section.

Keywords: exchange rates, excess currency returns, uncovered interest parity, convenience yield, global financial cycle, global risk, monetary policy.

JEL Classification: E52, C32, E43, F31, G15, F41

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1 Introduction

In view of the sharp rise in inflation observed since 2021, the US Federal Reserve has initiated a turnaround after years of accommodative monetary policy. Interest rates have been raised faster and more sharply than at any time since former Fed Chairman Paul Volcker's rate hikes in the early 1980s. At the same time, risk aversion in global financial markets is high, given the large uncertainty about the further course of the Covid-19 pandemic and the energy crisis triggered by the war in Ukraine. The ongoing tightening of monetary policy will not only affect the US economy, but will inevitably spill over to the rest of the world, with the US Dollar exchange rate and excess dollar returns being important transmission channels given the global US Dollar dominance in trade invoicing, asset issuance, and official reserve holdings. Excess returns after monetary policy shocks are understood as evidence against the uncovered interest parity (UIP) condition.

The impact of monetary policy shocks on exchange rates and excess dollar returns has been the subject of a large number of studies throughout the 2010s, with results being far from conclusive. In standard models for open economies, a monetary policy shock affecting the interest rate differential between countries should change the exchange rate in such a way that the uncovered interest parity (UIP) condition is satisfied: A positive interest rate differential is associated with an (expected) currency depreciation. However, there are also theoretical arguments and some empirical evidence that US monetary tightening has a depreciating effect on the US Dollar exchange rate (Stavrakeva & Tang, 2019; Ilzetzki & Jin, 2021; Inoue & Rossi, 2019).

In addition to the wide range of estimation approaches used, a plausible explanation for the inconclusiveness is that the interaction of monetary policy and the exchange rate is shadowed by omitted influencing factors. Consistent with the literature on the global financial cycle (Rey, 2015; Miranda-Agrippino & Rey, 2020), studies show that global factors affect responses of exchange rates and excess currency returns. This holds especially for the US Dollar with its prominent role as the dominant world currency (Lustig et al., 2011; Kalemli-Özcan, 2019; Kalemli-Özcan & Varela, 2021; Cormun & De Leo, 2022). In a related context, Inoue & Rossi (2019) show that the impact of monetary policy on exchange rates differs depending on how monetary policy affects agents' expectations and their perceived effects on the riskiness/uncertainty in the economy during particular episodes. Thus, the strong linkages between exchange rates, interest rates, and global risk are pronounced, making it difficult to accurately identify their causal relations. Moreover, recent studies also argue that the persistence of monetary policy shocks plays a significant role in how market interest rates, exchange rates, and, correspondingly, excess returns react (De Michelis & Iacoviello, 2016; Schmitt-Grohé & Uribe, 2022).

This paper examines the causal relationships between US monetary policy shocks, exchange rates, and currency excess returns by estimating a structural vector autoregressive (SVAR) model. To trace the dynamic responses of soundly identified shocks, we rely on Local Projection Impulse Response Functions (LPIRFs), as suggested by Jordà (2005). Our first contribution to the existing literature is methodological. Various assumptions are used in the existing literature to solve the identification problem, most turning out to be too restrictive for the research question under review. For instance, recursive approaches, as used by Hnatkovska et al. (2016), must either assume that the policy rate does not directly affect exchange rates or that central banks do not respond to the exchange rate, both of which are highly controversial.[1] Identification with sign restrictions, as applied, for example, by Faust & Rogers (2003); Scholl & Uhlig (2008) and Kim et al. (2017), allows simultaneous linking of financial variables, but has the disadvantage of being based on otherwise stringent assumptions about the qualitative effects of monetary policy shocks (Baumeister & Hamilton, 2019). While narrative arguments for identification - or similarly - high frequency information (see, for instance Romer & Romer, 2004; Jarociński & Karadi, 2020; Müller et al., 2022) typically aim at the reliable detection of partially identified shocks, their scope is limited for full system identification in light of restrictive exogeneity conditions and demanding assumptions with regard to instrument relevance. We apply an identification approach that places no explicit constraints on the behavior of our model variables, while allowing for a full and simultaneous interaction between US monetary policy, global risks, and exchange rates. It builds on contributions to data-based identification of SVARs that take advantage of the uniqueness

[1] See also Gertler & Karadi (2015) and Caldara & Herbst (2019), who caution against the recursive approach in VARs that model both macroeconomic and financial variables.

of independent components in linear non-Gaussian systems that are considered as structural shocks (Comon, 1994).[2] Model implied structural shocks have sound economic properties.

Also from an economic perspective, we provide valuable contributions to the existing literature in several respects. To the best of our knowledge, we are the first to systematically capture different dimensions in which monetary policy affects exchange rates and currency excess returns by including an external shock as well as by both a transitory and a persistent monetary policy shock into the SVAR system. More specifically, like Cormun & De Leo (2022), we separately identify monetary policy shocks and an external shock. The external shock is identified from exogenous innovations in the exchange rate of the US Dollar. We interpret the external shock as being closely related to global risk. This is based on the results of, e.g., Krishnamurthy & Lustig (2019); Stavrakeva & Tang (2019); Georgiadis et al. (2021) and Cormun & De Leo (2022), who show that a global risk factor accounts for a sizeable share of the variation in the dollar exchange rate. This also finds confirmation in Corbo & Di Casola (2022), who interpret an exogenous exchange rate shock as a change in the general risk premium charged by investors for holding assets in foreign currency.

Similar to Williamson (2016); Uribe (2022); Evans & McGough (2018); Garin et al. (2018); Garcia-Schmidt & Woodford (2019); Lukmanova & Rabitsch (2020) and Bilbiie (2022), we explicitly distinguish between temporary and persistent monetary policy shocks. This distinction is currently of particular importance given that it is reasonable to consider that a longer-term trend reversal in the monetary policy stance has set in. As is standard practice in Dynamic Stochastic Equilibrium (DSGE) models, we study the transmission of a transitory monetary policy shock through the response to an innovation in the short-term nominal interest rate and the transmission of a persistent monetary policy shock through the response to an innovation in a low-frequency inflation measure. The underlying idea is that a persistent monetary policy shock raises both the nominal interest rate and inflation in the long run. Therefore, it can also be interpreted as an increase in the inflation target (Uribe, 2022). Similar to Lukmanova & Rabitsch (2020), we use the Survey of Professional Forecasters' (SPF) long-term inflation forecasts as a proxy for the central bank's long-term

[2] Identification by means of independent components as detected in this work is successfully employed in the context of US monetary policy analysis and exchange rate modelling (see, e.g., Bernoth & Herwartz, 2021; Jarociński, 2022; Herwartz et al., 2022b,c) and (Herwartz & Wang, 2023)

inflation target. As such, our paper also ties into a relatively new strand of macroeconomic studies on Neo-Fisher effects. It allows us to verify Schmitt-Grohé & Uribe (2022)'s result that, in contrast to a temporary monetary tightening that leads to an appreciation of the US Dollar and thus excess returns to be earned on the US Dollar, a permanent monetary policy shock depreciates the US Dollar and causes excess returns to be earned on the foreign currency.[3]

Our third contribution is that, similar to Müller et al. (2022), we analyze whether the emergence of excess currency returns can be associated with a failure of UIP. Since UIP is an ex-ante relationship between exchange rates and interest rate differentials, we also estimate the response of *expected* excess currency returns on the dollar to monetary policy shocks and the external shock. While Müller et al. (2022) rely on survey data, we determine the exchange rate expectations needed for the calculation of ex-ante excess currency returns as in-sample fitted values implied by our model regressions.

Fourth, our paper also contributes to the literature manifesting a break in the relationship between US monetary policy, exchange rates, and excess currency returns during the Great Recession following the Great Financial Crisis (GFC). Stavrakeva & Tang (2019) find that the US Dollar surprisingly appreciated during the GFC and in the years after in response to expansionary US monetary policy shocks. Bernoth et al. (2022) find that the appearance of excess currency returns appears to be very much a pre-crisis phenomenon for most currencies. For the period prior to the GFC, they show that cross-country differences in currency excess returns against the US Dollar can be related to conventional measures of risk, most predominantly a global risk factor in line with the result of Lustig et al. (2011). However, for the post-crisis period, they reject that currency excess returns reflect a risk compensation.

And finally, we analyze in more detail the nature of our external shock. Following Krishnamurthy & Vissing-Jorgensen (2012); Engel & Wu (2018); Krishnamurthy & Lustig (2019) and Jiang et al. (2021), foreign investors receive a convenience yield by holding dollar-

[3]Lukmanova & Rabitsch (2020) confirm that the statistical properties of the US inflation targeting process imply that targeting shocks can indeed be viewed as long-lasting changes in monetary policy. However, unlike in Uribe (2022), in this study inflation targeting shocks are not considered permanent in a strict sense, but very persistent. For this reason, we speak of persistent rather than permanent monetary shocks.

denominated safe assets, which reflects their value in liquidity and safety while lowering their return requirements. Thus, shifts in the demand and supply of safe dollar assets are important drivers of the US Dollar exchange rate and affect investment returns. Accordingly, the US Dollar exchange rate reflects, to a significant extent, a global risk factor in the form of a shortage of safe dollar assets. Thus, we investigate whether the external shock we identify is related to global risk aversion and the convenience yield on US Dollar Treasuries, which would be evidence for the existence of a so-called convenience yield channel of monetary policy. Our study also contributes to the growing literature that sheds light on how the role of the United States as the world's safe asset supplier has shaped the dynamics of the dollar exchange rate and excess returns (Gourinchas et al. (2010); Bruno & Shin (2015a); Maggiori (2017); Ilzetzki & Jin (2021); Jiang et al. (2021).

Using a monthly data set covering the period 1980M1 to 2022M12 and a cross-section of eight advanced economies, i.e. Australia, Canada, Germany, Japan, New Zealand, Sweden, Switzerland, and the United Kingdom, we find that the dynamics of the US dollar exchange rate is the main driver of excess currency returns, rather than that of interest rates. The impact of US monetary policy on the US Dollar exchange rate depends on the type and persistence of the monetary policy shock as well as on an external shock. This external shock is strongly related to global risk aversion and the convenience yield that investors are willing to pay to hold US dollar assets. A significant part of the response of excess currency returns to the shocks considered is also expected, suggesting a violation of UIP. Focusing only on the post-crisis period, the impact of both the external shock and the inflation targeting shock on exchange rates and currency excess returns has vanished. This may explain why currency excess returns declined or even disappeared after the GFC and the Great Recession, as shown by Burnside (2019) and Bernoth et al. (2022).

The remainder of this paper proceeds as follows. In the next section, the data and the VAR model are presented and the data-based identification approach is described in detail. Section 3 presents the theoretical features of the structural shocks and the assignment of sound economic labels to the statistically identified shocks. Section 4 presents the estimation results of the macroeconomic response profiles to the distinguished shocks. Section 5 characterizes the identified external shocks in detail. Section 6 concludes. The appendices provide further information on the implementation of the data-based identification (Appendix A), on the data sources (Appendix B), on the diagnostic tests for normality and fundamentality (Appendix C) and on the structural parameter estimates (Appendix D).

2 Empirical Strategy

2.1 Data

We analyze the causal relationship between US monetary policy shocks, exchange rates, and currency excess returns by means of a set of country-specific structural VARs. This section briefly sketches the employed VAR models in reduced and structural form and encounters the sufficient conditions for uniqueness of independent structural shocks.

Our empirical analysis employs monthly data spanning the period 1980M1 to 2022M11.[4] Throughout, we consider the United States as the domestic country, while a set of eight foreign countries, i.e., the United Kingdom, Japan, Canada, New Zealand, Australia, Sweden, Switzerland and Germany, give rise to a cross section of alternative empirical model implementations. The country selection obtains from the following considerations. First, we want to focus on advanced economies. Various studies, in fact, show that exchange rate behavior and the occurrence of UIP deviations differ significantly between emerging and advanced economies (Kalemli-Özcan, 2019; Kalemli-Özcan & Varela, 2021). Hence, mixing these two types of economies could lead to inconclusive results. Second, we would like to look at a time period as long as possible to have sufficient sample information to examine the hypothesis that a potential change of structural relations can be traced back to changes in the importance of the US Dollar as an international reserve currency. Third, we intend to compare our results with those of Schmitt-Grohé & Uribe (2022), who focus their analysis on the United Kingdom, Japan, and Canada. Therefore, our dataset includes these three economies as well, but we also provide evidence on the robustness of the results by using an extended set of economies (including Australia, Germany, New Zealand, Sweden, and Switzerland).

[4] We also split the full sample information into pre-crisis and post-crisis subsamples to shed light on the eventually modified transmission of structural shocks after the GFC and the Great Recession (see Section 4.6).

2.2 Excess currency returns

Deviations from the uncovered interest rate parity (UIP) are described by the concept of *excess currency returns (ECR)*. ECRs measure the potential gains of carry trade strategies that consist, for instance, of borrowing in a low-yield currency, investing in a high-yield currency, and conversion of the portfolio at maturity. Let ex-post ECR be defined as:

$$ECR_{t+12} = (r_t^* - r_t) - (s_{t+12} - s_t), \tag{1}$$

where r_t^* denotes the foreign one year treasury bill rate. Thus, ECRs are zero if a positive (negative) interest rate differential between foreign countries and the US is offset by a US Dollar appreciation (depreciation) of a similar magnitude. Thus, if $ECR_{t+12} > 0$, investors holding foreign assets over the next 12 months will earn an excess return, while if $ECR_{t+12} < 0$, investors holding US Dollar assets will benefit from an excess return.

However, when ECRs are present, one cannot straightforwardly conclude that UIP does not hold. Ex-post ECRs can also occur because markets systematically misjudge exchange rate developments, but this should not be interpreted as a failure of UIP. UIP only fails when investors expect excess returns, as Froot & Frankel (1989); Gourinchas & Tornell (2004); Kalemli-Özcan & Varela (2021) and Candian & De Leo (2023) argue. Accordingly, we also examine *Expected Excess Currency Returns* (EECR), which are defined as follows:

$$EECR_{t+12} = (r_t^* - r_t) - (E_t[s_{t+12}] - s_t), \tag{2}$$

where $E_t[s_{t+12}]$ denotes the, at time t, expected exchange rate for 12 months ahead. Thus, expected exchange rates are determined as in-sample fitted values of the stylized regression:

$$s_{t+12} = \alpha + r_t^* \beta_1 + r_t \beta_2 + s_t \beta_3 \hat{\pi}_t \beta_4 + \varepsilon_t \tag{3}$$

2.3 A cross section of structural VARs

Conditional on presample values $y_0, y_1, ..., y_{1-p}$, we consider country specific VARs of dimension $K = 3$ and lag order p that read in their reduced and structural form, respectively, as

$$\begin{aligned}\Delta y_t &= \nu + A_1 \Delta y_{t-1} + \ldots + A_p \Delta y_{t-p} + u_t, &(4)\\ &= \nu + A_1 \Delta y_{t-1} + \ldots + A_p \Delta y_{t-p} + D\epsilon_t,\ t = 1, 2, ..., T, &(5)\end{aligned}$$

where Δ is the first difference operator[5] (i.e. $\Delta y_t = y_t - y_{t-1}$) and $y_t = (r_t, s_t, \pi_t^e)'$ is the vector of endogenous variables.[6] In particular, r_t denotes the US one year treasury bill rate. We choose a maturity of one year because, as Gertler & Karadi (2015) and Rüth (2020) also argue, a monetary policy interest rate indicator with such a slightly longer maturity has a wider distance to the zero lower bound and is also an effective strategy to capture the role of forward guidance during the Great Recession following the GFC.[7] Moreover, s_t is the log nominal FX rate in foreign currency per US Dollar, while $\hat{\pi}_t$ is a measure of US inflation target, which serves as an indicator of persistent shifts in monetary policy (Uribe, 2022). Like Lukmanova & Rabitsch (2020), we use as a proxy for the inflation target the mean of inflation expectations for the next ten years from the SPF.[8] Note that we refrain from adding ECR or EECR as a fourth variable to our VAR to avoid a multicollinearity problem, since they are composed of the variables included in our three-dimensional VAR. In order to analyze their cause and effect relationships with the variables and shocks under consideration, we rely on the concept of local projections, as described in the following.

By assumption, the model in (4) is causal, i.e., $\det(A(z)) \neq 0\ \forall |z| \leq 1$, where $A(L) = I_K - A_1 L - A_2 L^2 - \ldots - A_p L^p$ and L is the lag operator such that, e.g., $L\Delta y_t = \Delta y_{t-1}$. The corresponding Wold representation reads as $Y_t = \Phi(L)u_t$, $\text{Cov}[u_t] = \Sigma_u$. Finally, u_t is a serially uncorrelated vector process with mean zero and covariance Σ_u, ν is a vector of intercepts, $A_1, A_2, \ldots, A_p$ are $K \times K$ parameter matrices. By means of OLS or ML estimation

[5] As discussed further below, the variables in y_t are not cointegrated according to conventional diagnostics, so we estimate the model in first differences.

[6] We have also considered $K = 4$ dimensional models including foreign treasury yields. With regard to the three shocks of interest in this work, the informational content of the four dimensional system is similar to the one of trivariate models. For instance, regarding largest available samples for the UK, Japan, and Canada the correlation between model specific (i.e., $K = 3$ vs. $K = 4$) US temporary shocks is 0.894, 0.943, and 0.973, respectively. For the remaining two shocks, the respective six correlation statistics are between 0.961 and 0.988.

[7] Bernoth et al. (2022) show that the appearance of excess currency returns is maturity dependent and only occurs for maturities longer than one month. However, they analyze excess currency returns unconditional on monetary policy shocks.

[8] Lukmanova & Rabitsch (2020) show that differences across specifications with alternative inflation target measures are minor and that estimation results are robust across various measures of low-frequency inflation, including 10-year ahead inflation expectations of the SPF.

the reduced form parameters and the residuals u_t can be estimated consistently.

2.4 Identification

2.4.1 Uniqueness of non-Gaussian independent components

An important contribution of our work to the existing literature is its innovative identification of structural shocks that account for potential bidirectional causalities among the variables in u_t (and, hence, y_t) in a largely agnostic manner. By assumption, the structural parameter matrix D in (5) is nonsingular.[9] Hence,

$$\epsilon_t = D^{-1}u_t \text{ and } \mathrm{Cov}[u_t] = DD' =: \Omega. \tag{6}$$

It is well known that, in a Gaussian framework ($u_t \sim N(0, \Omega)$), the identification of the parameter matrix D requires external information (e.g. the assumption of a recursive causal structure; Sims, 1980), since rotations of Gaussian random vectors are observationally equivalent. An important result in Comon (1994) states that the linear transmission scheme on the left hand side of (6) allows for a unique recovery of D from (estimates of) u_t, if (i) the components of ϵ_t are mutually independent, and (ii) at most one of the elements ϵ_{it} exhibits a Gaussian distribution. It is worth noting that, for the present case of analyzing financial market variables and outcomes, the deviations from Gaussianity (e.g. fat tails) are well established in the respective literature. In this context, Jarociński (2022) explores the non-Gaussian properties of monetary policy shocks and uses independent components analysis by Comon (1994) to identify their underlying structure. The author notes that the identified shocks provide an intuitive interpretation and plausible effects. Furthermore, despite not imposing external information, the shocks are remarkably similar to those identified in the existing literature using Gaussian methods. Hence, independent components detection

[9] We also follow the convention to investigate effects of positive structural shocks and assume that the diagonal elements of D greater than zero.

appears as a promising solution to achieve identification in a data-based manner. [10]

The data-based approach to identification that we pursue in this study consists of determining the matrix D in a way that joint dependence among the implied shocks $\epsilon_t = D^{-1}u_t$ is minimal in terms of a flexible non-parametric dependence measure, namely the so-called Cramér-von-Mises (CvM) distance of Genest et al. (2007).[11] Since the statistical identification does not build upon behavioral economic relationships, the discussion of theoretical impact effects in Section 3 and the 'shock labelling' analysis in Section 4.1 provide important complementary insights into the properties of the statistically identified independent shocks. While the use of economic a-priori information fixes the structural shocks by construction, shocks identified by means of a statistical criterion (such as mutual independence) do not necessarily feature sound economic properties. Herwartz & Lütkepohl (2014) discuss the problem of so-called 'shock-labeling' in detail. In fact, using data-based identification in SVARs requires the assignment of sound economic labels to the detected shocks as an additional modelling step. To support the economic labelling of the statistically identified shocks (i.e. independent components), we provide an extensive literature review in Section 3 below on the theoretical and empirical transmission channels that shape the contemporaneous relationships among short-term US yields, exchange rates, and long-term inflation expectations as our measure for the (perceived) inflation target. This helps us plausibly identify the expected impact of exogenous shocks hitting our SVARs via the three endogenous variables under consideration.

[10]By means of Monte-Carlo experiments Herwartz et al. (2022a) compare several alternative data-based approaches to identification in SVARs. An important finding of this study is that nonparametric variants of independent component analysis, such as those employed in this study, perform accurate and largely robust under a wide variety of data-generating models, including scenarios of heteroskedastic shocks that are likely to affect our model variables due to the coverage of the GFC. While informative (co)variance changes have also been suggested for SVAR identification in a number of papers (e.g., Rigobon, 2003; Lanne & Lütkepohl, 2008), we consider the robust performance of independent component analysis in a cross-section of VAR models as an important merit of the identification of shocks in the form of independent components.

[11]For more details on the adopted ICA-based approach to identification and a formal representation of this estimator see Appendix A. For computation, we employ modified functions of the R package *svars* of Lange et al. (2017).

2.5 Local projections and rolling regressions

To evaluate the dynamic responses of macro variables to the identified shocks $\epsilon_{k,t}$, - we employ local projection impulse responses along the lines of Jordà (2005). Specifically, we determine local projection impulse response functions (LPIRFs) as (cumulated) parameter estimates $\hat{\beta}_h$ from the following iteration of regressions

$$z_{t+h} = \alpha + \beta_h \epsilon_{k,t} + \sum_{i=1}^{p} \gamma_i z_{t-i} + u_{t+h}^{(h)}, h = 0, 1, ..., H-1, \tag{7}$$

where $z_t \in \{\Delta r_t, \Delta r_t^*, \Delta s_t, (E)ECR_t\}$ and the parameters α and γ are estimated separately for each horizon.[12] Throughout, we set $p = 4$, which aligns with our interest in modelling monthly data and the order of the reduced form VAR models.

To further characterize the identified external shock and to test the hypothesis that it is related to international demand for safe US dollar assets and global risk aversion, we follow Lilley et al. (2022) and evaluate rolling regressions of the form:

$$y_s = \alpha + \beta_0 \epsilon_{2,s} + u_s, \; s = \tau_1, \tau_1 + 1, \ldots, \tau_2, \tag{8}$$

where $y_s \in \{\Delta TB_t, \Delta log(VXO_t)\}$, $\epsilon_{2,s}$ is the external shock and τ_1 (τ_2) is the lower (upper) bound of rolling samples of size 60. Iterations cover time periods from $\tau_1 = 1, \ldots, T - 60 - 1$ to $\tau_2 = 60, \ldots, T - 1$.[13] Moreover, $\Delta log(VXO_t)$ is a proxy for global risk appetite calculated as the monthly change in the log implied volatility of the S&P100 stock index, and ΔTB_t denotes the change in the average US Treasury basis against the G10 economies. The treasury basis is determined as the average of the differences between the yield on an actual one-year US Treasury and the yield on an equivalent synthetic US Treasury constructed from G10 foreign bonds.[14] A negative US Treasury basis indicates that an actual US Treasury is

[12] Note that the response of the ex-post spot exchange rate 12 months ahead does not capture the effects originated in the horizons before. This fault is also present in the directly estimated responses of the ex-post exchange rate differential and ex-post ECRs. To address this issue, we reconstruct these responses with the estimated response of the spot exchange rate and the interest rate differential. For ex-ante ECRs we do not face this issue as the ex-post spot rate is substituted by its expectations.

[13] The overall patterns and trends identified in rolling regressions with time windows comprising 120 observations are more persistent but of a similar shape. Respective figures are available upon request.

[14] The yield on a synthetic US Treasury is calculated as the yield on a foreign bond converted into dollars at maturity and hedged against currency risks.

more expensive than a foreign equivalent counterpart, which represents a deviation from the Covered Interest Rate Parity (CIP) in government bond markets. Thus, the US Treasury basis serves as a measure for the convenience yield on US Treasury securities reflecting the value that investors place on liquidity and safety, thus lowering their yield requirements. Hence, the convenience yield rises when the preference for US Dollar assets increases. For more information, please refer to Krishnamurthy & Lustig (2019).

3 Shock labelling

The first shock considered is the temporary innovation to the short-term US nominal interest rate (*US Temp. shock*). As Rey (2015), Georgiadis (2016), and Ehrmann & Fratzscher (2009) show, conventional US monetary policy has significant spillover effects on foreign economies, which are particularly strong for countries with relatively liquid and open financial markets. However, the reaction of the US Dollar exchange rate to an exogenous increase in US interest rates is not clear-cut, as there are theoretical arguments for effects in both directions. According to the UIP condition, higher US interest rates should lead to an immediate appreciation of the US Dollar. Krishnamurthy & Lustig (2019) add the consideration that when the Fed tightens monetary policy, bond markets assume that a reduction in the supply of safe dollar assets is imminent. As a result, the marginal willingness of global investors to pay for the safety and liquidity of dollar-denominated assets increases, leading to an appreciation of the dollar.

However, there are also arguments for a depreciating effect of a US monetary tightening on the US Dollar exchange rate. One is that higher interest rates increase the debt service burden on companies and governments, which reduces overall investment and growth prospects, while also increasing pressure on the banking system. As pointed out by Gürkaynak et al. (2021), another argument is that an increase of US policy rates may signal higher than expected inflation, which invokes a depreciation of the US Dollar.

The ambiguity of the impact of a monetary policy shock on exchange rates is also reflected in the empirical literature. Consistent with seminal papers by Eichenbaum & Evans (1995) and Faust & Rogers (2003), Müller et al. (2022), Rüth (2020), and Schmitt-Grohé &

Uribe (2022) find an immediate positive relationship between the US Dollar exchange rate (appreciation) and US interest rates. However, Stavrakeva & Tang (2019) find that this relationship flipped during the Great Recession. Their explanation is that, in times of crisis, the described signaling effect of monetary policy dominates. An unexpected tightening of US monetary policy signals economic strength, leading to a decline in risk aversion and higher expected inflation in the US, thus resulting in a depreciation of the US Dollar. Last, but not least, Inoue & Rossi (2019) find evidence that the exchange rate responses differ with the effects of monetary policy on agents' expectations of risk premia in the short, medium, and long run during specific episodes. Thus, the response of the US Dollar exchange rate to a monetary policy shock might be state-dependent and, consequently, the theoretical sign is left open.

According to a DSGE model of Lukmanova & Rabitsch (2020), the on-impact response of the inflation target to a positive nominal US interest rate shock should be either zero under full information or negative under imperfect information. Their empirical estimations suggest that the (perceived) inflation target does not respond to temporary monetary policy shocks, in line with predictions of the DSGE model under full information. Allowing for potential state dependence, the respective theoretical (net) effect, as stated explicitly in Table 1, is either negative or zero.

The second shock under consideration is an unexpected US Dollar exchange rate change (*FX shock*). The response of US short-term interest rates to a surprise increase in the exchange rate (US Dollar appreciation) can be expected to be negative. According to results of Krishnamurthy & Lustig (2019), Stavrakeva & Tang (2019), Georgiadis et al. (2021), and Cormun & De Leo (2022), a considerable part of the fluctuations in the US Dollar exchange rate is due to a global risk factor. This is also confirmed by Corbo & Di Casola (2022), who interprets an exogenous exchange rate shock as a change in the overall risk premium charged by investors for holding assets in a foreign currency. According to Krishnamurthy & Lustig (2019), the US Dollar appreciates when the marginal willingness of foreign investors to pay for US Dollar-denominated safe assets increases, which is the case, for example, when global risk appetite declines. As a result, US short term interest rates decline. Thus, the theoretical impact effect of an exogenous FX shock on the short-term US interest rate is negative.

To our knowledge, so far there is little theoretical and empirical evidence in the literature on how an exogenous exchange rate shock to the US Dollar affects the (perceived) inflation target. Therefore, we leave the theoretical sign open.

The third shock considered is the inflation target shock, an indicator of persistent shifts in monetary policy (*US target shock*). Theoretical models do not provide a clear indication of the direction in which short-term US interest rates and the US Dollar exchange rate should respond to a persistent monetary policy shock. When agents are fully informed that a monetary policy shock is persistent in nature via an upward revision of the inflation target, the associated immediate increase in inflation expectations leads to a decline in real interest rates, which has an expansionary effect on output, as Lukmanova & Rabitsch (2020), Uribe (2022) and Schmitt-Grohé & Uribe (2022) argue. In response, short-term nominal interest rates in the US will gradually and persistently rise in response to the increase in the inflation target, which is also confirmed by a New Keynesian model in Garin et al. (2018) and is referred to as the Neo-Fisher effect. With imperfect information, however, market participants have limited information about the central bank's objective and must learn about the nature of the monetary shock over time to distinguish permanent shifts in the inflation target from temporary disruptions in the monetary policy rule. In this case, agents may initially misinterpret an inflation target increase with an expansionary interest rate shock and inflation would not adjust immediately leading to a negative gap between actual and targeted inflation. Under such conditions, short-term nominal interest rates would initially react negatively, then become positive after some time, when the learning process is complete (Lukmanova & Rabitsch, 2020).

Mumtaz & Theodoridis (2018), Lukmanova & Rabitsch (2020) and Schmitt-Grohé & Uribe (2022) study the effects of a persistent expansionary monetary policy shock by means of an SVAR model. They find, on impact, a negative effect on short-term nominal US interest rates that turns positive after a few quarters, which is in line with the Neo-Fisher effect under imperfect information. We take this information and remain agnostic about the way a permanent monetary policy shock affects our model variables.

In response to a positive US inflation target shock and the corresponding increase in long-term inflation expectations, the US Dollar should immediately depreciate (negative reaction), although nominal interest rates increase in response, as described above. Empirical estimates by Schmitt-Grohé & Uribe (2022) are consistent with this prediction, which they call the open economy Neo-Fisher effect. However, since the impact of a persistent monetary policy shock on the US Dollar exchange rate is one of our research questions, we do not take an a-priori stand on the expected sign.

Table 1 summarizes the theoretical sign pattern behind our shocks.

Table 1: Theoretical sign patterns of structural shocks

Variable	US Temp Shock	FX-Shock	US Target Shock
r_t	+	−	?
s_t	?	+	?
π_t^*	?	?	+

4 Estimation Results

4.1 Structural estimates and an empirical 'test of concept'

The uniqueness of the identified components $\epsilon_{k,t}$, as determined in this study, only holds under informative deviations from the joint Gaussian model. In addition, structural impulse response estimates, as determined in this study, are only reliable if the shocks can be considered fundamental. Diagnostic results documented in Appendix C confirm that both statistical preconditions (i.e. non-Gaussianity and fundamentalness) are fulfilled for the considered set of empirical (S)VARs. As a further underpinning of the informational content of the shocks retrieved from the model in (4) we notice that the variables in y_t are not cointegrated according to conventional diagnostics. For instance, testing for cointegration among US interest rates and inflation expectations with full sample information yields an ADF-statistic of -1.502 which lacks significance at conventional levels.

The discussion of theoretical impact effect directions (see Table 1) reveals that we consider, in particular, the marginal response of US interest rates to an external shock as important for a sound economic labelling of the statistically identified shocks. Noting that our analysis covers cross sectional results for a total of eight structural models, it is interesting to unravel how far the unrestricted data-based estimates allow for a cross sectionally (almost) uniform interpretation. In this regard, the empirical estimates of the structural parameter d_{12} could be considered of special importance, as the theoretical discussion postulates negative impact effects of the external shock on US short-term interest rates.

Instead of providing a set of metric estimation results, Table 2 displays the absolute frequencies of estimated effect directions on impact (left hand side panel) and for the sum of structural IRFs from $h = 0$ up to horizon $h = 3$ (i.e. effects within one quarter, right hand side panel). As it turns out, several unrestricted structural parameter estimates imply effect directions that are (almost) common for the entire cross section. US interest rates largely respond in a negative way to unexpected increases in the FX rates. More specifically, the documented directional estimates are theory-conforming for six and seven (out of eight) economies when focusing on impact and within-quarter effects, respectively. Interestingly, interest rate responses to the persistent US monetary policy shock are uniformly positive, consistent with the neo-Fisher effect hypothesis in the case of fully informed investors.

Although the sensitivity of excess exchange rate returns will be discussed in detail below, it worth highlighting that the empirical analysis points to effect directions of structural shocks that apply to (almost) all foreign economies under scrutiny. For instance, at the cross sectional level exchange rates respond positively (negatively) to temporary (persistent) monetary policy shocks and inflation expectations tend to fall in response to an unexpected depreciation of the US dollar.

Given both the diagnostic evidence pointing to the fundamentalness of the structural shocks and the cross sectionally comparable results for several structural parameters, we can - in summary - conclude that the agnostic data-based approach to identification yield structural shocks featuring sound economic labels. Accordingly, it is tempting to address how these shocks affect short-term US and foreign treasury yields, US Dollar exchange rates or excess currency returns.

4.2 Responses to temporary US monetary policy shocks

Figure 1 shows the cumulative local projection impulse response functions (LPIRFs) of a temporary contractionary US monetary policy shock on US and foreign 12-month government bond yields, the nominal exchange rate, and ex-post ECRs for the longest periods of

Table 2: Empirical impact directions of structural shocks

	On Impact ($h = 0$)			Within one quarter ($h = 0, 1, 2, 3$)		
	US Temp	FX-S.	US Target	US Temp	FX-S.	US Target
r_t	$+(8) - (0)$	$+(2) - (6)$	$+(8) - (0)$	$+(8) - (0)$	$+(1) - (7)$	$+(8) - (0)$
s_t	$+(6) - (2)$	$+(8) - (0)$	$+(1) - (7)$	$+(7) - (1)$	$+(8) - (0)$	$+(0) - (8)$
π_t^e	$+(0) - (8)$	$+(2) - (6)$	$+(8) - (0)$	$+(8) - (0)$	$+(0) - (8)$	$+(8) - (0)$

Notes: The table shows the absolute number directional estimates obtained in the sample of eight economies (AUS, CAN, CHE, DEU, GBR, JPN, NZL, SWE). Identified SVAR models are estimated with the sample period 1980-2022. Explicit parameter estimates are documented in Appendix D.

available data. The upper four panels show the IRFs for the UK as reference country. Due to space considerations, the lower four panels summarize estimation results by providing LPIRF estimates for all considered economies jointly. While this collection of estimates lacks a complementation with model specific confidence bands, 'overall' significance of the displayed dynamics can be evaluated in terms of mean group criteria (Pesaran & Smith, 1995).[15]

For both observed sample periods, we confirm the results of Rey (2015), Georgiadis (2016), and Ehrmann & Fratzscher (2009) that conventional US monetary policy has significant spillover effects on foreign economies. As shown in panels (a) and (b), we find that foreign short-term nominal interest rates increase significantly in response to a positive *US temp shock*. The collection of cross sectional evidence in panel (f) confirms the positive responsiveness of foreign treasury yields to a conventional tightening of US monetary policy at the mean group level. A plausible explanation for the positive interest rate spill-over is provided by Bruno & Shin (2015b), who argue that an increase in the federal funds rate leads to a decline in bank leverage and cross-border credit flows. Thus, a rise in short-term US yields leads to a decline in available funds abroad, which exerts upward pressures on funding conditions and foreign interest rates. However, the foreign interest rate is rising at a somewhat more modest pace than US short-term nominal yields. As a result, the interest differential between the foreign country and the US ($r^* - r$) tends to decline (see panel ?).

Panel (c) shows the response of the GBP/US Dollar exchange rate to a temporary contractionary US monetary policy shock. The GBP/US Dollar exchange rate is not affected by a US interest rate hike. However, with regard to the entire cross section (see panel (g)), we observe that the displayed IRFs turn positive, on average, after about one year. Hence, mean

[15]Detailed country specific results are available from the authors upon request.

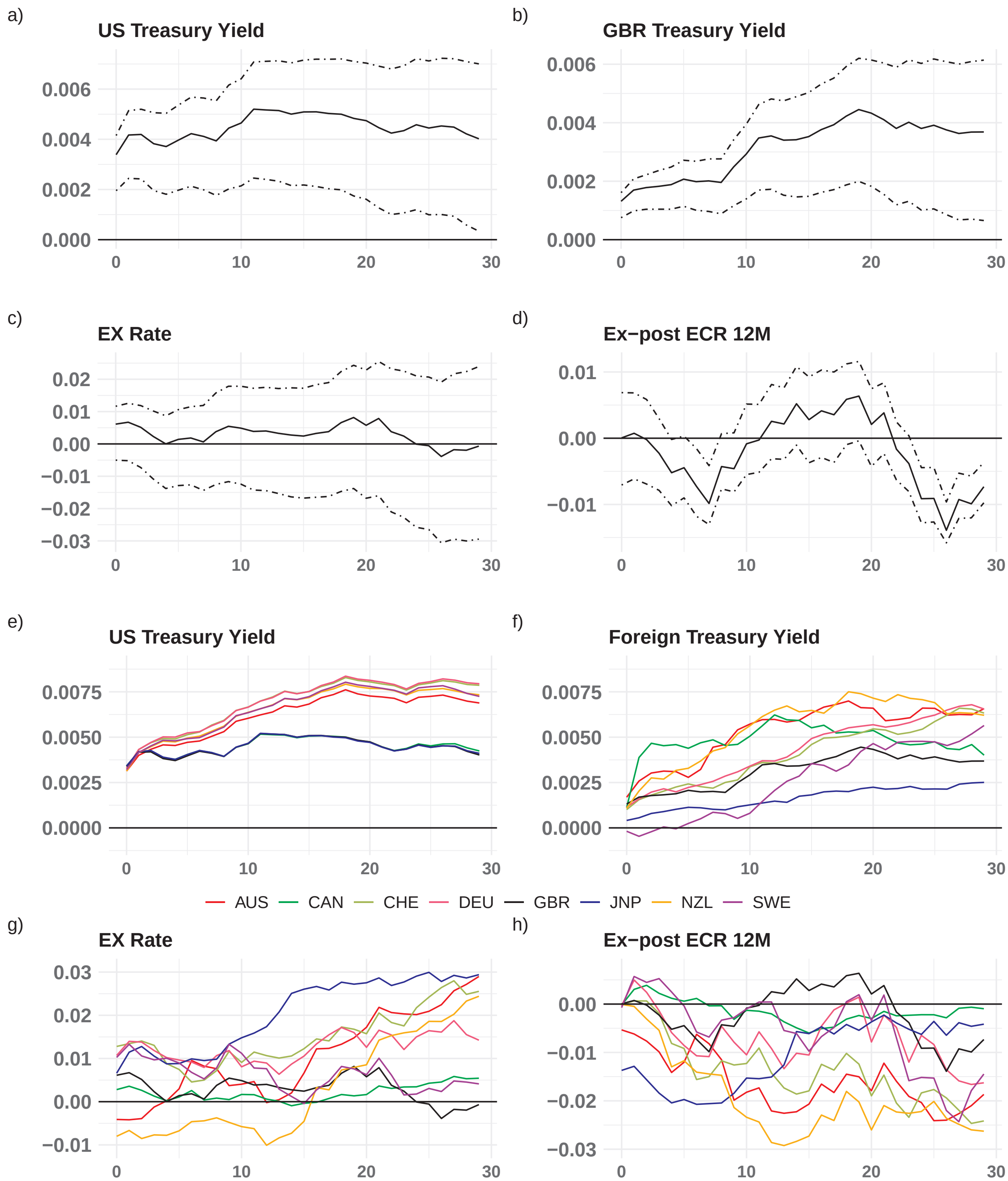

Figure 1: LPIRFs to a US temp shock, full sample

Notes: Cumulative LPIRFs of US and Foreign Treasury yields (12M), nominal EX rates and ex-post excess currency returns to US temp shocks. Sample periods are 1980M6-2021M11 (CAN, GBR and JPN) and 1987M6-2021M11 (AUS, CHE starting 88M5, DEU starting 88M3, NZL, SWE). Upper panels show results for the UK as reference economy joint with 95% bootstrap confidence intervals (panels (a) to (d)). We employ a Moving Block Bootstrap as suggested by Brüggemann et al. (2016) with block length of 25 ($\approx 5.03 * T^{1/4}$). Point estimates for the entire cross section of economies are in lower panels (e) to (h). The time axis refers to a period of up to 30 months.

group evidence indicates a lagged appreciation of the US Dollar rather than an immediate appreciation as it is predicted in the standard Dornbusch model.[16] Looking more specifically at the response of the exchange rate differential ($s_{t+12} - s_t$), the cross-sectional analysis confirms this pattern. The US Dollar exchange rate differentials turn positive (appreciation) after a few months and tends to increase for some time thereafter. In contrast to Müller et al. (2022) and Gürkaynak et al. (2021), we do not observe that the US Dollar exchange rates tend to depreciate thereafter again - at least not within the observed 30 month window.

The combination of a tendentially declining interest rate differential ($r^* - r$) and an appreciation of the US Dollar in response to a temporary increase in the nominal US interest rate indicates a rather negative reaction of excess currency returns. For the UK, the LPIRF of the ex-post ECR is mostly negative, but insignificant throughout (panel (d)). However, when looking at the entire cross-section, we indeed find a significant negative impact on excess US Dollar returns, indicating that investors holding US Dollar assets benefit from a temporary contractionary US monetary policy shock (see panel (h)), which is in line with the findings of Eichenbaum & Evans (1995), Schmitt-Grohé & Uribe (2022), and Müller et al. (2022).[17] When looking at the individual components of ECR, we find that the response of interest differentials is by far too small to compensate for the US Dollar appreciation following a US temporary monetary policy shock. Thus, we confirm the result of Kalemli-Özcan (2019) that, in advanced economies, excess currency returns are more likely associated with exchange rate fluctuations.

4.3 Responses to external shocks

Figure 2 shows the cumulative LPIRFs of an appreciation shock to the US Dollar exchange rate on US and foreign 12-month nominal interest rates, the nominal exchange rate, and the ex-post ECR for the period 1980M6 to 2022M11. Similar to Cormun & De Leo (2022), we interpret this *FX shock* as an external shock.

[16] To provide an example for mean-group inference, the average response of US Dollar exchange rates at the 12 month horizon is 0.62%. Given that the variance of empirical responses is about 5E[-05], we consider $\sqrt{5E[-05]/8} = 0.24\%$ to approximate the estimation uncertainty characterizing the average response. Hence, an implied mean-group t-ratio of $0.62/0.24 = 2.58$ indicates significance at conventional nominal levels.

[17] Note that Müller et al. (2022) show that the identification method of a monetary policy shock also plays a role in whether the ECR response is significant or not.

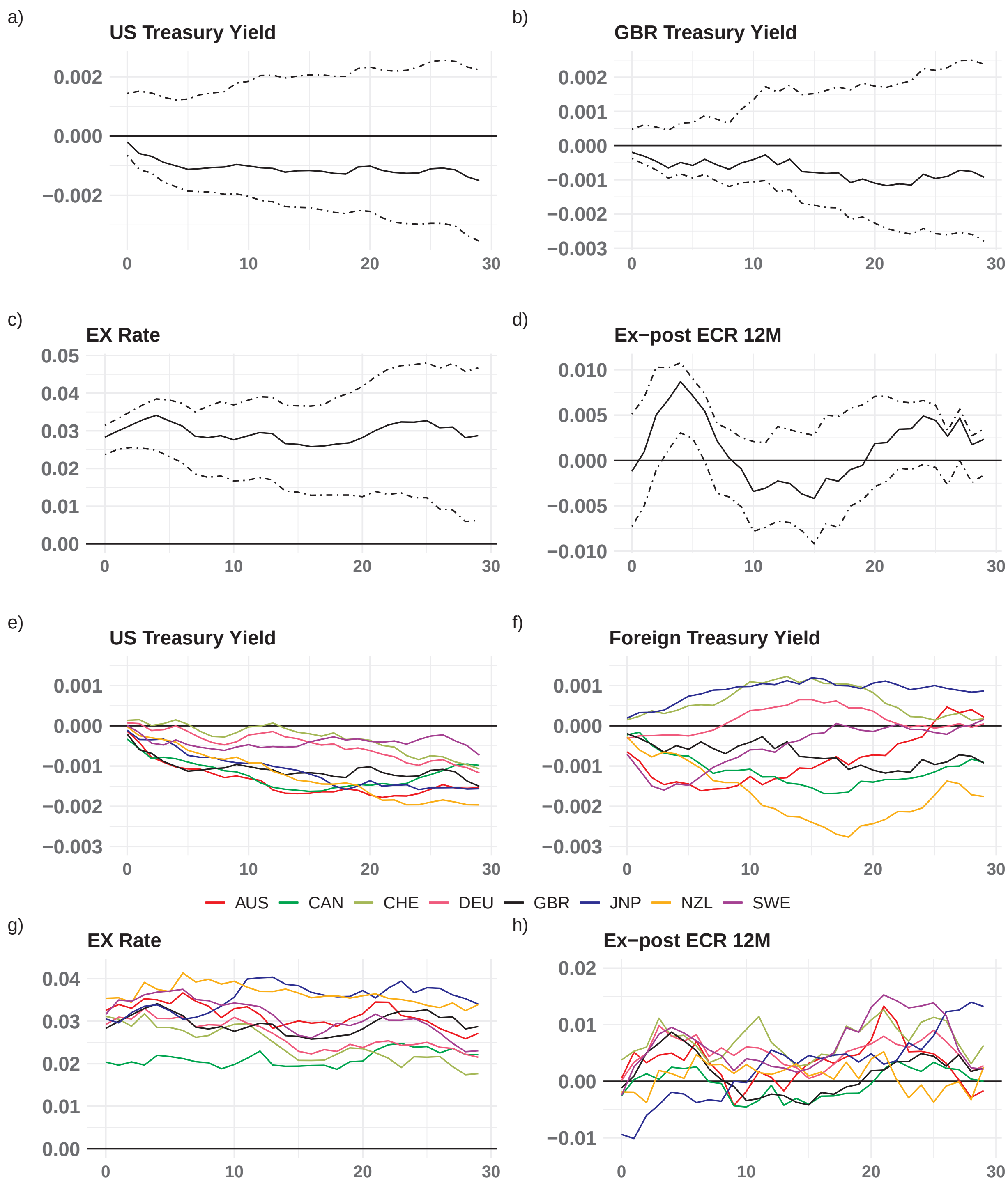

Figure 2: LPIRFs to a FX rate shock, full sample (For notes see Figure 1)

Consistent with our prediction, the short-term nominal interest rate in the US falls in response to an exchange rate shock. Although this response is not significant in the case of the UK (panel (a)), the cross-sectional analysis (panel (h)) suggests a significant decline after only a few months. In contrast, foreign short-term nominal interest rates do not seem to react significantly to an exchange rate shock, as suggested by both the LPIRF for the UK (panel (b)) and the cross-sectional results (panel (f)). In this way, the interest rate differential between foreign countries and the US rises significantly on average in response to an unexpected appreciation of the US Dollar. These results support the empirical findings of Cormun & De Leo (2022). They also align with theoretical arguments that the US Dollar exchange rate shock reflects a decline in global risk appetite. As such it increases the willingness of international investors to hold US Dollar-denominated assets, which tends to increase interest rate spreads against the US (Krishnamurthy & Lustig, 2019).

Exchange rates remain at the elevated level after a foreign exchange shock and do not show any noticeable appreciation or depreciation in the following months, as shown in panels (c) and (g). However, looking more specifically at the LPIRF of the exchange rate differential, $s_{t+12} - s_t$, our cross-sectional analysis shows an insignificant effect of an exogenous appreciation of the US Dollar for the first 18 months or so, which then turns significantly negative, indicating a lagged US Dollar depreciation.

Rising foreign versus US interest rate differentials, coupled with a lagged US Dollar depreciation, suggest that foreign excess returns will respond positively to an external shock, implying that investors in US Dollar assets will lose out. This is confirmed when looking at the cross-section of LPIRFs in panel (h). Although insignificant for selected countries like the UK or Canada, the cross sectional mean response of the ECR to an external shock is positive and increases over time, supporting the result of Cormun & De Leo (2022). Again, exchange rate dynamics are the main drivers rather than interest rate differentials. This result fits nicely with our hypothesis that the external shocks can be interpreted as a measure of global risk. When global risk increases, the US Dollar appreciates as the willingness of foreign investors to pay for US Dollar denominated assets rises. This induces an expected depreciation and thus lowers the returns of foreign investors on their holdings of US Dollar denominated assets as also described by Krishnamurthy & Lustig (2019).

4.4 Responses to inflation target shocks

Figure 3 shows the cumulative LPIRFs of an inflation target shock, our indicator of a persistent US monetary policy shift, on US and foreign 12-month nominal interest rates, the nominal exchange rate, and the ex-post ECR for the period 1980M6 to 2022M11. Focusing on the UK, Canada, and Japan, we find that short-term US Treasury bond yields gradually rise for about 12 months in response to a positive inflation target shock (see panels (a) and (e)). For these three countries, this positive co-movement between the inflation target and short-term interest rates fits Neo-Fisher dynamics and confirms the estimation results of Schmitt-Grohé & Uribe (2022), who focus on these three countries only. However, as shown in panel (e), this result does not seem to be robust and is rather country-specific.[18] The LPIRFs resulting from the estimates for the remaining five countries considered - Australia, Germany, New Zealand, Sweden, and Switzerland - show that US short-term interest rates do not react to an US target shock in the first 12 months and then fall continuously. This conditional negative co-movement between interest rates and the inflation target opposes the theoretical Neo-Fisher effect and empirical results of Uribe (2022), Valle e Azevedo et al. (2022), and De Michelis & Iacoviello (2016).

Unlike for US short-term interest rates, however, we find a clear answer as to how a positive US inflation target shock affects foreign nominal interest rates. Looking at both the case for the UK in panel (b) and the cross-sectional results in panel (f), we conclude that foreign nominal short-term interest rates rise significantly. This is consistent with the model results of Gürkaynak et al. (2021) and once more confirms the positive spillover effects of US monetary policy. Given the heterogeneous responses of US nominal interest rates, the overall effect of the US inflation target shock on the interest rate differential between foreign countries and the US is ambiguous and insignificant on average in the first 12 months. Only for subsequent periods does the cross-sectional analysis suggest a significant positive response

[18] Note that correlations between the structural innovations of the US target shocks identified in the country-specific SVARs are 0.99 and vary only in the third digit. The heterogeneous responses observed in panel (e) of figure 3 can therefore be attributed to the different starting points of the samples periods analyzed for each of the countries investigated. The responses obtained by analyzing the same truncated sample period for the United Kingdom, Japan, and Canada also show that neo-Fisherian effect after US target shocks disappears in the mid-1980s. This is also the case for the identified US temporary shocks, where the cross-country correlation is around 0.98. The specific samples used in the analysis are listed in the notes to Figure 1.

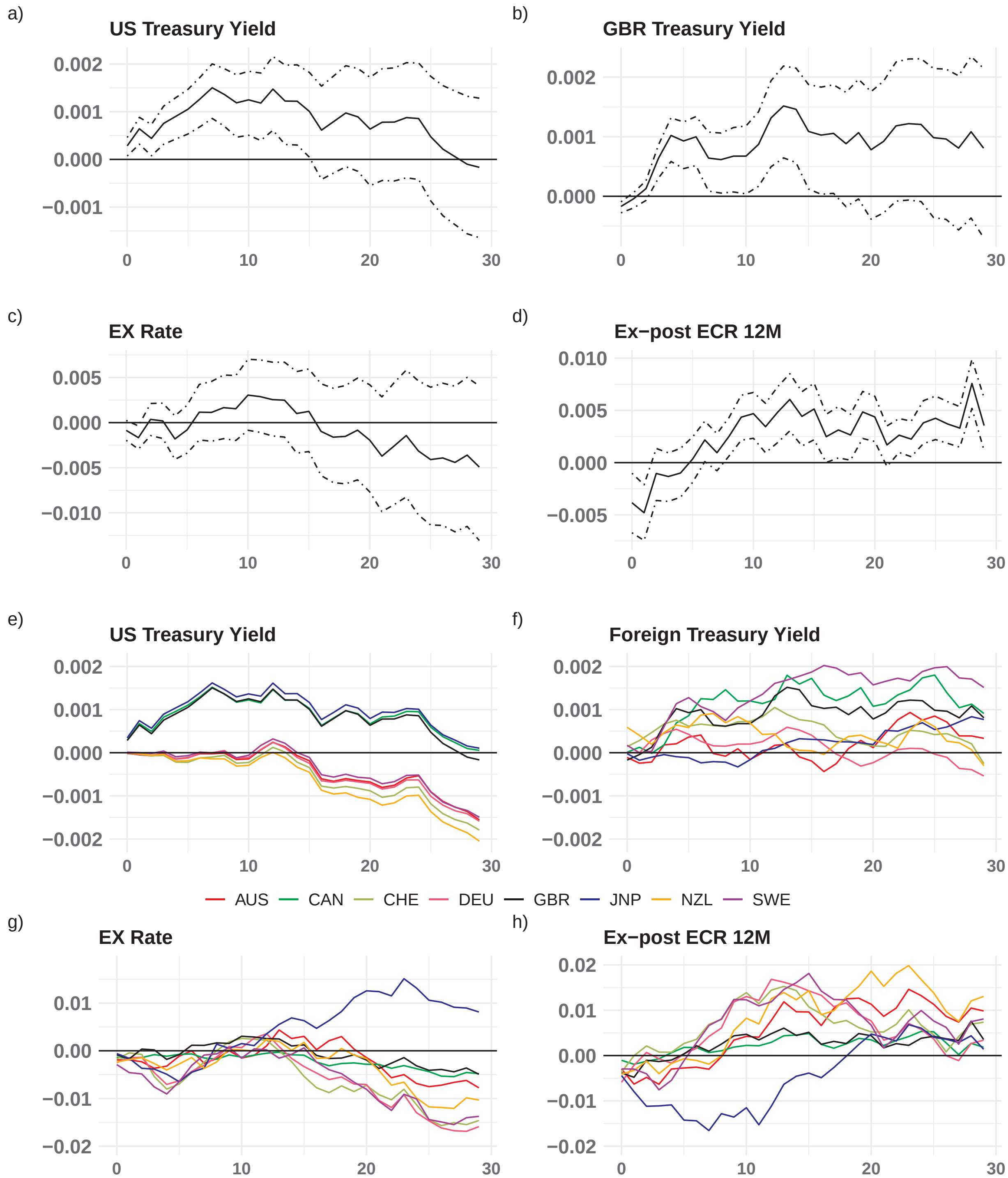

Figure 3: LPIRFs to a US target shock, full sample (For notes see Figure 1)

to the interest rate differential.

It should be recalled that, according to the Neo-Fisher hypothesis, the US Dollar should immediately depreciate in response to a positive US inflation target shock (negative exchange rate reaction). However, the LPIRF of the sterling exchange rate shows a rather insignificant reaction (panel (c)). However, looking at the entire cross-section, we indeed find a significant depreciation effect for the US Dollar exchange rate, with the only exception being the Japanese yen rate. Accordingly, the cross-sectional evidence in panel (i) confirms that the response of the exchange rate differential ($s_{t+12} - s_t$) is negative and indicates a depreciation of the US Dollar in response to a positive inflation target shock.

The lagged positive reaction of the foreign-US interest rate differential, combined with the continued depreciation of the US Dollar, suggests that the impact of an inflation target shock in the US on ex-post excess currency returns will be positive, as expected. This is confirmed by the LPIRFs for the British pound (panel(d)) and also the cross-sectional analysis (panel (h)). In the first two months after the inflation target shock, the impulse response is initially significantly negative, but then rises steadily and becomes positively significant after a few months. Thus, investors holding foreign assets earn a positive return in response to a persistent shift in US monetary policy that raises the inflation target and long-run inflation expectations. Again, also in the case of the inflation target shock, excess currency returns are mainly driven by exchange rate movements, which are not sufficiently compensated by the interest rate differential.

In summary, we find that the dynamics of the US dollar exchange rate are the main driver of excess currency returns. Both external factors and US monetary policy have an impact on the exchange rate and, hence, on excess currency returns. In response to a temporary tightening of US monetary policy, the US dollar appreciates, generating significant excess currency returns in favor of investors holding US dollar assets. In response to a sustained increase in the inflation target, the US dollar depreciates, generating significant excess returns in favor of investors holding foreign-denominated assets. An external shock in the form of an unexpected appreciation of the US dollar, which, as we show in Section 5, can be associate with a decline in global risk appetite, leads to a depreciation of the US dollar in the following months, which is associated with negative returns earned by investors holding

US dollar assets.

4.5 Ex-post versus ex-ante excess currency returns

So far we have focused on ex-post ECRs. However, as outlined before, in order to investigate whether the UIP holds as a function of the shocks under study, the focus should not be on the actual exchange rate, but on the exchange rate *expected* by the markets (Froot & Frankel, 1989; Gourinchas & Tornell, 2004; Kalemli-Özcan & Varela, 2021; Candian & De Leo, 2023). Similar to Müller et al. (2022), we therefore analyze in this section how expected or ex-ante ECRs respond to temporary and persistent monetary policy as well as to the external shock. While these authors rely on survey data on expected exchange rates to calculate expected excess returns, we determine the exchange rate expectations needed for the calculation as in-sample adjusted values implied by our model regression, as described in equation (2).

Figure 4 shows the results. In contrast to Müller et al. (2022), who find that the response of expected excess returns to a temporary US monetary policy shock does not differ markedly from zero, our cross-sectional analysis shows that it is significantly negative. Thus, we find that UIP indeed does not hold, since excess returns in the foreign exchange market are expected and cannot be attributed entirely to expectation errors. However, while ex-post currency excess returns fall very persistently in the wake of the monetary policy shock, the drop in expected excess returns is only temporary, and, on average, becomes insignificant after about one and a half years.

Similar to the ex-post results, the response of *expected* excess currency returns to an external shock turns out to be significantly positive and very persistent, both at the country level and at the cross-sectional level. Thus, in this case the occurrence of excess returns cannot be attributed to expectation errors. Investors holding foreign assets also expect to earn an excess return, indicating a violation of UIP.

The response to the US inflation target shock only seems to be different when we focus on expected rather than observed exchange rate data. Our cross-sectional evidence suggests that the expected excess returns to the US Dollar conditional on a US inflation target shock are now insignificant. In this case, rather than attributing the excess returns to a violation of the UIP, we have to conclude that investors systematically do not expect the subsequent

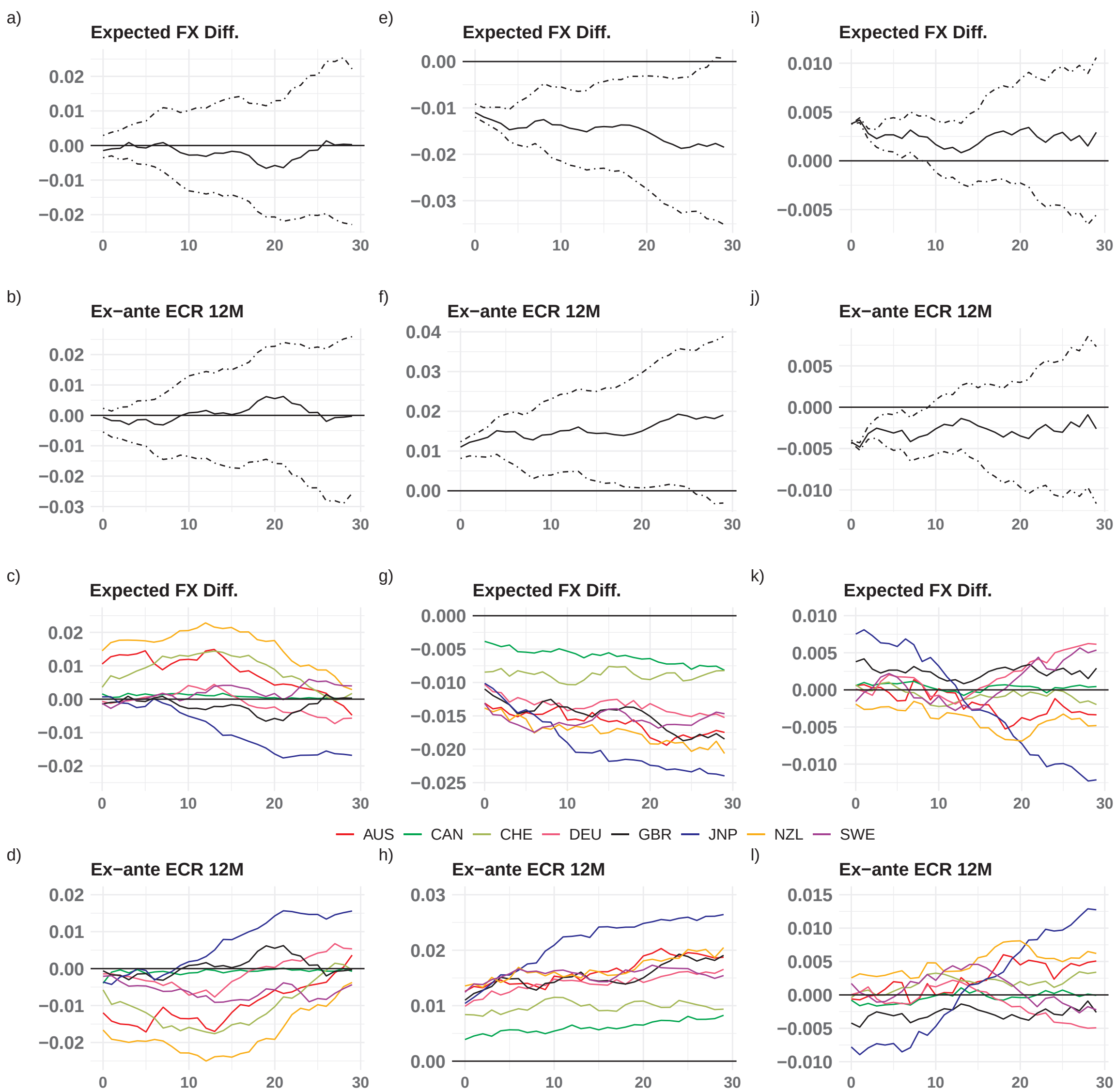

Figure 4: LPIRFs for expected FX differentials and ex-ante ECRs, full sample

Notes: The figure displays local projections for the entire sample and all economies. Endogenous variables comprise the expected exchange rate differential and ex-ante ECRs, both with a maturity of 12-month. Three blocks from left to right display results for the US temp, the external, and the US target shock.

depreciation of the US Dollar.

In summary, we cannot rule out the possibility that a significant part of the response of the excess currency returns to the shocks considered is due to a violation of the UIP and not to expectation bias.

4.6 Pre- and post crisis period

There is some evidence in the literature that the occurrence of excess currency returns depends on the sample period. For instance, Bernoth et al. (2022) find a structural break in 2007 with the onset of the GFC. Focusing on the pre-crisis period, they estimate significant excess currency returns in favor of US Dollar denominated assets for a set of currencies of advanced economies. Moreover, they find that excess currency returns are indeed matched by covariances with (global) risk factors. However, when they focus on the post-crisis period, excess returns have disappeared and no longer reflect risk compensation. In a similar vein, Falconio (2016) examines the relationship between US monetary policy and excess returns from currency carry trades. He shows that, prior to the onset of the financial crisis in 2008, expansionary policy shifts lead to a decrease in international risk aversion, which in turn leads to higher carry trade returns. However, when they focus on the post-crisis period from 2008 to 2015, they find that international risk aversion is no longer influenced by US monetary policy and, hence, neither are excess currency returns.

Both of these above mentioned papers look at reduced form correlations, which are generally not indicative of the strength of structural relationships, like those explored in this paper. In this section, we test whether we also observe a different response of excess currency returns conditional on US monetary policy shocks and the external shock before and after the GFC. For this, we split our data set into two subsamples, a pre-crisis period 1980M6 to 2007M4 and a post-crisis period 2008M4-2021M12. Figure 5 shows the results.[19]

[19] As we demonstrate in Section 5, the nature of the external shock is subject to a transition process starting from mid-2007. In order to exclusively capture the new external variation, we exclude the period of transition and leave a gap of one year between the two sub-samples.

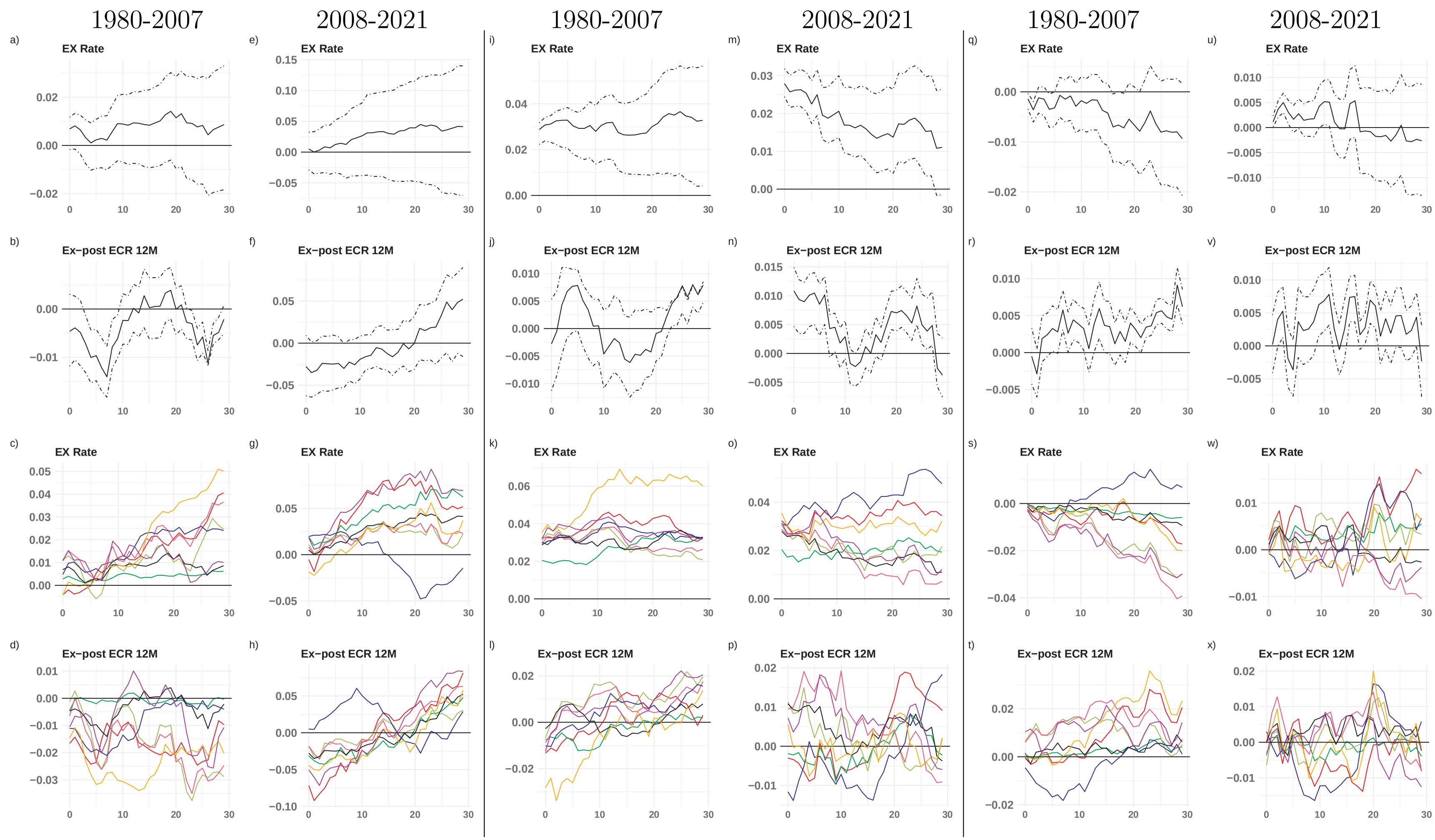

Figure 5: LPIRFs to a US temp, external, and US target shock, sub-samples

Notes: The figure displays local projections for two sub samples and all three shocks. Sub-samples comprise the periods 1980M6-2007M4 (CAN, GBR and JPN) and 1987M6-2007M4 (AUS, CHE starting 88M5, DEU starting 88M3, NZL, SWE), and 2008M4-2021M11 (AUS, CAN, CHE, DEU, GBR, JPN, NZL and SWE). The endogenous variables considered are the expected exchange rate and ex-post ECRs with 12-month maturity. Three blocks from left to right display results for the US temp, the external shock, and the U.S. target shock.

For all three shocks considered, the LPIRFs for the pre-crisis period are very similar to those for the full sample. However, if we focus only on the period after 2008M4, we find markedly distinct responses, mainly of the exchange rate and, accordingly, of the excess currency returns, which confirms a structural break after the GFC.

In response to a temporary monetary policy shock, the US Dollar also appreciates on impact, as it does in the full sample. What is new, however, is that the US Dollar starts to depreciate after about 20 months, such that the cross-sectional exchange rate return $s_{t+12} - s_t$ becomes significantly negative after about five years. As excess currency returns are mainly driven by exchange rate dynamics, we find a similar pattern for ex-post ECRs in the post-crisis period. They fall after a temporary US monetary policy shock and then rise steadily, turning positive after about 20 months.

Interestingly, the cross-sectional analysis for the post-crisis period suggests that the impacts of both the external shock and the inflation target shock on currency excess returns lack significance. Looking at the underlying components of these returns, we see that exchange rates in individual countries react very heterogeneously and not clearly in one direction to these two shocks. Therefore, their response is now insignificant from the perspective of the cross-sectional analysis. This corresponds to the findings of Bernoth et al. (2022) that, when focusing on the post-crisis period, currency excess returns no longer reflect just a compensation for (global) risk but could explain why currency excess returns declined or even vanished after the GFC and the Great Recession, as Burnside (2019) and Bernoth et al. (2022) show. Moreover, the Neo-Fisher effect, which states that monetary policy influences exchange rates and excess currency returns via shocks to the inflation target, seems to apply only to the pre-crisis period.

5 The nature of the external factor

In this section, we analyze the nature of the external/foreign exchange rate shock in more detail. According to Krishnamurthy & Vissing-Jorgensen (2012); Engel & Wu (2018); Krishnamurthy & Lustig (2019), and Jiang et al. (2021), important drivers of fluctuations in US Dollar exchange rates are shifts in the demand and supply of safe dollar assets. The

US Dollar exchange rate clears the global market of these safe assets. The supply of safe dollar assets is largely determined by monetary policy, while the demand for US Dollar safe assets is significantly impacted by global risk, as highlighted in the literature on the existence of a global financial cycle (Rey, 2015; Miranda-Agrippino & Rey, 2022). During periods of relatively low global risk appetite, global cross-border capital flows contract and demand for safe US Dollar assets increases, causing the US Dollar to appreciate.

Against this background, we examine whether the external shock is indeed related to the attractiveness of US Dollar assets to international investors and to global risk aversion. As described in Section 2.5, we proxy global risk appetite with VXO, the implied volatility of the S&P100 stock index. Similar to Krishnamurthy & Lustig (2019), we measure the attractiveness of holding US dollar assets with the so-called convenience yield on US Treasury securities, proxied by the US Treasury basis against the G10 countries, and conduct rolling regressions on the identified external shock as described by equation (8). We examine the immediate effects of external shocks in the period of their occurrence. Figure 6 displays the results.

Panel (a) of Figure 6 shows coefficient estimates from rolling regressions of the US Treasury basis on the external shock for the period 2000M3 to 2022M11.[20] Until mid-2003, the cross-sectional average of these coefficient estimates suggests an insignificant relationship between our exchange rate shock and the convenience yield. However, from 2004 onwards, the cross-sectional average of the slope coefficient becomes significantly negative. Around 2007, there is an abrupt and sustained further decline in the estimated slope coefficients. From mid-2011, the average slope coefficient increases again and becomes insignificant around 2018. These results suggest that the convenience yield explains a significant part of our identified external factor. In particular, during periods of high uncertainty and global risk, the role of the US dollar as the primary global safe-haven asset is a strong driver of US dollar exchange rate developments that are unexplained by US monetary policy shocks. This provides evidence for the existence of a so-called convenience yield channel of monetary policy.

[20]We can only plot the rolling window coefficients from 2000M3 onwards, as data for the US Treasury basis is only available from 1995M2 and we always use 60 observations in the estimation process. For the VXO, data availability starts in the 1980s but ends in 2021M8. However, for reasons of comparability, we only present results for the same sample as the US Treasury basis.

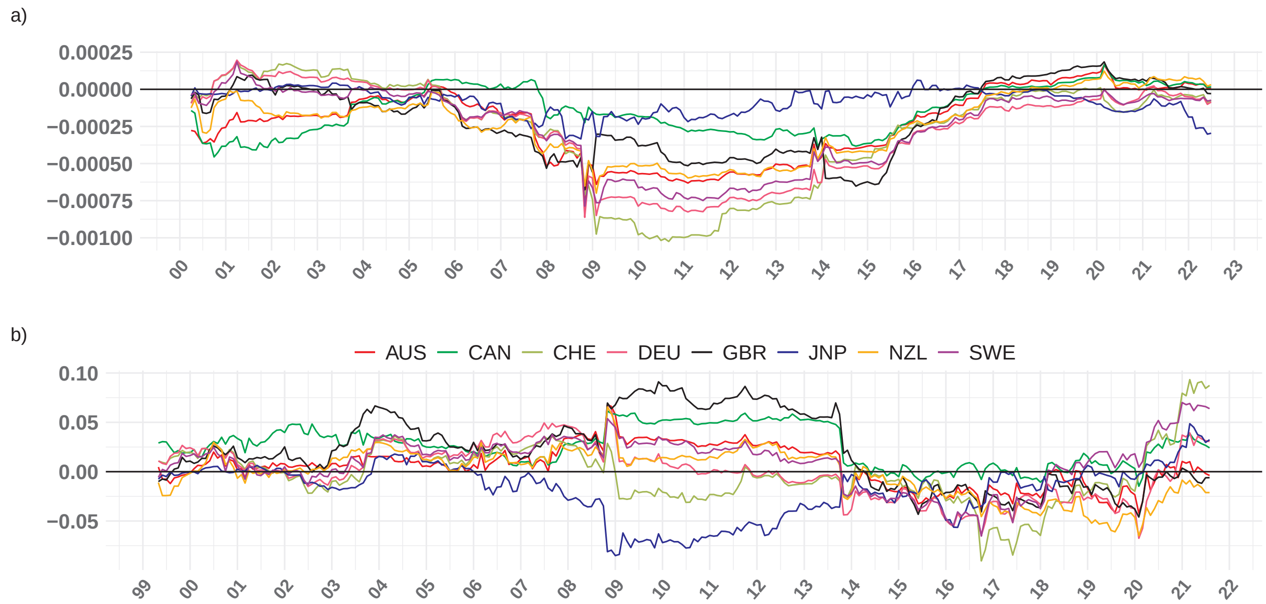

Figure 6: Historical immediate effects of past external shocks

Notes: The figure shows the estimated slope coefficients for the rolling regression analysis using time windows of 60 months (see section 2.5 and equation 8). Panels (a) and (b) show rolling regression results for the immediate effects of external shocks on the US Treasury basis and the VXO, respectively. The sample period for panel (a) is from 2000M3 to 2022M6, and that for panel (b) is from 2000M3 to 2021M8 due to limited data availability of the VXO index.

US monetary policy spills over to other economies through changes in the demand for and supply of safe US dollar assets, even in the absence of changes in policy rates.

Panel (b) of Figure 6 displays the coefficient of the VXO from the rolling regression results. Until 2014, the collection of cross-sectional evidence suggests a significant positive relationship between our exchange rate shock and the VXO, indicating that the US Dollar experiences external appreciation in periods when global risk appetite is low (high VXO). This positive link became, on average, stronger during the GFC between 2008 and 2014. However, we also see some variation across countries. For Japan, Switzerland, and, in part, Germany, the relationship between VXO and the external shock turns negative during the GFC. Thus, during the GFC and the years after, an unexpected appreciation of the US dollar against these three countries is associated with a decline in global risk appetite. A plausible explanation is that these three currencies are also seen as safe havens by international investors. Therefore, in times of crisis, the US dollar does not appreciate unexpectedly against these three currencies, as can be observed for the other currencies. Interestingly, between 2014 and 2021, the period of the Great Recession, the mean group estimate of the rolling window coefficient on VXO is no longer positive, but significantly negative, suggesting that an unexpected appreciation of the US dollar is associated with periods of low global risk appetite. It is not until 2021 that this relationship reverses again. We leave it to further research to determine why this sign reversal occurred.

In sum, the external shock we identify is significantly related to a conventional proxy for global risk aversion, albeit with changing signs. Thus, we find that the dynamics of the US dollar exchange rate that are unrelated to monetary policy can be explained, to a significant extent, by global risk factors. Further, contrary to the finding of Lilley et al. (2022), this holds not only after the onset of the GFC in 2007, but throughout the investigated sample period.

6 Conclusion

How does US monetary tightening affect the US dollar exchange rate and currency excess returns? In this paper, we show that the persistence of monetary policy shocks as well as the level of global risk aversion and the corresponding demand for safe-haven US dollar assets play an important role in answering this question.

Using a monthly data set covering the period 1980M1 to 2022M12 and a cross-section of eight advanced economies, i.e. Australia, Canada, Germany, Japan, New Zealand, Sweden, Switzerland and the United Kingdom, we find that excess currency returns are mainly driven by exchange rate movements, which are not sufficiently compensated by the interest rate differential. The impact of US monetary policy on the US Dollar exchange rate, and hence on excess currency returns, depends on the type and persistence of the monetary policy shock. In response to a temporary tightening of US monetary policy, the US Dollar appreciates, generating significant excess currency returns in favor of investors holding US Dollar assets. In response to an increase in the inflation target, which represents a persistent monetary policy shock, the US Dollar depreciates, generating significant excess returns in favor of investors holding foreign-denominated assets.

We show that US Dollar exchange rate excess returns are also affected by an external shock. An external shock in the form of an unexpected appreciation of the US dollar leads to a depreciation of the US dollar in the following months, which is associated with negative returns for investors holding US dollar assets. The identified external shock is strongly related to global risk aversion and the convenience yield that investors are willing to pay for holding US dollar assets. Thus, lower global risk appetite and greater demand for safe US dollar assets are associated with a US Dollar appreciation that cannot be explained by US monetary policy shocks.

Focusing only on the post-crisis period, the impact of both the external shock and the inflation targeting shock on exchange rates and currency excess returns has vanished. This may explain why currency excess returns declined or even disappeared after the GFC and the Great Recession, as shown by Bernoth et al. (2022). Moreover, the Neo-Fisher effect, which suggests that monetary policy influences exchange rates and currency excess returns via shocks to the inflation target, also seems to apply only to the pre-crisis period. Why this is the case should be the subject of future research.

Finally, we analyze whether the emergence of excess currency returns can be associated with a failure of UIP. Since UIP is an ex-ante relationship between exchange rates and interest rate differentials, we re-run our estimations using *expected* rather than ex-post excess currency returns. We find that a significant part of the response of excess currency returns to the shocks considered is also expected, suggesting a violation of UIP.

Our findings add to the understanding of the properties of exchange rates, interest rate fluctuations, and excess currency returns. Shocks to the demand for safe dollar assets as well as global risk aversion explain part of the variation in the dollar exchange rate and, hence, excess dollar returns. This information could help solve the exchange rate disconnect puzzle and shed light on the optimal design of monetary policy.

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Appendices

A - Estimation of structural model parameters

Building upon the result of Comon (1994), a variety of approaches to ICA-based point estimation of the structural parameter matrix D in (5) have been suggested (e.g., Moneta et al., 2013; Matteson & Tsay, 2017; Lanne et al., 2017; Gouriéroux et al., 2017).[21] In this study, we estimate D by means of an approach that can be considered as a modification of the estimator in Matteson & Tsay (2017), which has been successfully employed, for instance, by Bernoth & Herwartz (2021). Avoiding an explicit distributional assumption, the estimator of D is obtained by selecting the particular structural matrix that obtains implied shocks with weakest dependence in terms of the Cramér-von-Mises (CvM) distance (Genest et al., 2007),

$$\mathcal{B} = \int_{(0,1)^K} \left[\sqrt{T} \left(C(\tilde{\epsilon}) - \prod_{k=1}^{K} U(\tilde{\epsilon}_k) \right) \right]^2 d\tilde{\epsilon}, \tag{9}$$

where C and U denote the empirical copula of orthogonalized model disturbances and the implied copula under independence, respectively. Since the CvM-distance is constructed from (joint) ranks, it is scale free. Genest et al. (2007) consider it an 'ideal' choice for nonparametric dependence diagnosis unless sufficient support for a local dependence alternative is available. As we are not aware of such an alternative in the analysis of heterogeneous economies, we determine an estimator for D by solving the minimization problem

$$\widehat{D} = \widetilde{D}_{\hat{\theta}}, \text{with } \hat{\theta} = \text{argmin}_{\theta}\{\mathcal{B}|\tilde{\epsilon}_t = \widetilde{D}_{\theta}^{-1} u_t\}. \tag{10}$$

To implement (10), we use rotation matrices that structure the space of potential decompositions of the reduced form residual covariance estimates $\widehat{\Sigma}_u = GR_\theta R_\theta' G' = \widetilde{D}_\theta \widetilde{D}_\theta'$, where G is a lower triangular Cholesky factor of $\widehat{\Sigma}_u$ and $R_\theta R_\theta'$ is the identity matrix. Hence, $\widehat{D} = GR_{\hat{\theta}}$. Random vectors $\tilde{\epsilon}$ are determined from orthogonalized reduced form model disturbances ($\tilde{\epsilon}_t = \widetilde{D}_\theta^{-1} \hat{u}_t$), and

[21] Kilian & Lütkepohl (2017) review alternative ICA approaches and embed these variants of data-based identification into the SVAR literature. Assuming independence of shocks is more strict than the typical orthogonality assumption. However, this restriction is also implicit in the stylized construction of impulse response functions tracing the effects of isolated unit shocks (by setting $E[\epsilon_{jt}|\epsilon_{it} = 1] = 0$, $i \neq j$).

the rotation matrices are specified as the product of three Givens rotation matrices, i.e.

$$R_\theta = \begin{pmatrix} \cos\theta_1 & -\sin\theta_1 & 0 \\ \sin\theta_1 & \cos\theta_1 & 0 \\ 0 & 0 & 1 \end{pmatrix} \begin{pmatrix} \cos\theta_2 & 0 & -\sin\theta_2 \\ 0 & 1 & 0 \\ \sin\theta_2 & 0 & \cos\theta_2 \end{pmatrix} \begin{pmatrix} 1 & 0 & 0 \\ 0 & \cos\theta_3 & -\sin\theta_3 \\ 0 & \sin\theta_3 & \cos\theta_3 \end{pmatrix}.$$

The minimization outlined in (10) can be achieved by means of nonlinear optimization.[22]

It is worth noting that the point estimate $\widehat{D}$ that solves (10) is unique up to the signs and ordering of its columns, since changing the column ordering or multiplying single columns with minus unity does not change $\widehat{D}\widehat{D}'$. To establish uniqueness of column signs and ordering (and hence comparability of economy-specific estimates $\widehat{D}_i$), we opt for the particular ordering that yields a maximum sum of (absolute) diagonal elements. Following, for instance, Lütkepohl & Netšunajev (2017) this ordering establishes that a particular shocks exerts its strongest effect on the variable to which it is primarily associated. If - given this column ordering - a particular diagonal element is negative, we multiply the respective column with minus unity. Thereby, sign uniqueness establishes that the analysis focuses on the effects of positive shocks.

B - Data sources and used samples

- **Interest rates,** r_t, r_t^*: Treasury yields with 12-month maturity (End Month). Source: For Australia, Germany, New Zealand, Japan, Sweden, Switzerland, the United Kingdom, and the United States: *Macrobond*. For Canada: *Refinitiv* from CANSIM - Statistics Canada.

- **US treasury basis with the G10 economies**: Treasury yields with 12-month maturity for the United States, United Kingdom, Japan, Canada, Sweden, Switzerland, Australia, New Zealand are taken from *Macrobond*. For Germany, we obtain data from *Refinitiv*. For observations prior to 1997, we construct implicit bond rates with 12-month maturity from German treasury yields with 10-years maturity.

- **US Inflation expectations,** π^e: Median of the estimate of the CPI inflation rate over the next 10 years in percentage points. Source: Survey of Professional Forecasters, Federal Reserve Bank of Philadelphia downloaded from *Macrobond*.

[22] Procedures are implemented in the R package ‘svars’ (`https://cran.r-project.org/package=svars`) as provided by Lange et al. (2017). To guard against the potential of a local optimum we try 100 alternative initializations with randomized seeds and extract a global optimum accordingly.

- **Nominal exchange rates,** s_t: Source: *Macrobond.* Exchange rates are listed in foreign currency per US Dollar.

- **Forward points**: Source: *Macrobond.* We use the spot rates from Macrobond to transform the forward points into forward rates.

- **VXO**: CBOE S&P 100 Volatility Index. Source: *FRED Economic Data, St. Louis Fed.*

The time series and sample periods used for the calculations and estimates are as follows:

- To estimate the three-dimensional VARs, we use data from 1980M1 to 2022M11 for all eight economies. The time series used are 12-month US Treasury yields, bilateral spot exchange rates, and 10-year US inflation expectations.

- To calculate ex-post and ex-ante ECRs, we use data on foreign treasury yields. However, the availability of foreign treasury yields is limited. Therefore, we use the following sample periods for the different economies: Australia (1987M10-2022M11), Canada (1980M1-2022M11), Japan (1980M1-2022M11), New Zealand (1987M7-2022M11), Sweden (1987M07-2022M11), Switzerland (1988M1-2022M11), and the United Kingdom (1980M1-2022M11).

- For the estimation process of the local projections, the samples start from the same periods as in the ECR calculations. Due to the loss of data in the ECR calculations, we must restrict these samples to 2021M11. We restrict the end of the first sub-sample for all economies to 2007M4. The second sub-sample is homogeneous for all economies from 2008M4 to 2021M11.

- To calculate the average US Treasury basis against the G10 economies, we use time series from 1995M2 to 2022M6 of 12-month US and foreign Treasury yields, spot exchange rates, and 12-month forward rates.

- To estimate the rolling regressions, we use data on the US Treasury basis and the VXO from 1995M2 to 2022M6.

C - Tests of normality and fundamentalness

The structural analysis pursued in this work relies on the identifying assumption of non-Gaussianity of structural shocks and the existence of the Wold representation for the vector valued VAR process y_t. Diagnostic results displayed in Table 3 indicate highly significant deviations from the Gaussian distribution for all identified shocks in all considered economies. In addition, the shocks deviate from moment conditions that are typical for the joint normal. We detect both significant skewness and excess kurtosis. Table 4 documents test outcomes for the null hypothesis that the data are in line with the existence of a Wold representation. With 5% significance, we cannot reject the null hypothesis of fundamentalness for all economies except New Zealand and all implementations of the test statistic. Undocumented results show that (i) extending the lag-order to $p = 12$ results in p-values of at least 35% and (ii) Johansen trace tests of the null hypothesis of a zero cointegration rank are throughout insignificant. From these diagnostics we conclude that a Wold representation exists for the considered economies.

Table 3: Univariate and multivariate normality tests for the structural shocks.

		Univariate			**Multivariate**		
Country		US Temp	External	US Target	Multi JB	Skewness	Kurtosis
AUS	stat.	1442.268	145.461	10235.963	11823.692	655.055	11168.636
	p-value	0	0	0	0	0	0
CAN	stat.	1447.641	656.965	9860.256	11964.86	627.306	11337.56
	p-value	0	0	0	0	0	0
CHE	stat.	1446.906	29.643	10475.759	11952.308	621.369	11330.939
	p-value	0	0	0	0	0	0
DEU	stat.	1421.707	19.952	10221.973	11663.631	611.411	11052.221
	p-value	0	0	0	0	0	0
GBR	stat.	1345.264	74.195	9880.962	11300.42	598.874	10701.55
	p-value	0	0	0	0	0	0
JPN	stat.	1309.574	73.714	10504.484	11887.77	645.159	11242.61
	p-value	0	0	0	0	0	0
NZL	stat.	1317.945	632.87	10084.081	12034.895	684.759	11350.136
	p-value	0	0	0	0	0	0
SWE	stat.	1329.727	151.513	10322.531	11803.77	643.722	11160.048
	p-value	0	0	0	0	0	0

Notes: Univariate Jarque-Bera tests for the single structural shocks (US temp shock, external shock and US target shock) are documented in the left hand side. Tests for joint normality, symmetry, and no excess kurtosis of all structural shocks are shown in the right hand side panel. Diagnostics refer to structural innovations identified in three dimensional VARs of lag order 4.

Table 4: Testing fundamentalness of VAR residuals

p-max	1	2	3	4	5	6	7	8
AUS	0.256	0.255	0.250	0.245	0.241	0.237	0.232	0.229
CAN	0.054	0.061	0.068	0.075	0.080	0.087	0.093	0.099
CHE	0.141	0.151	0.160	0.165	0.167	0.170	0.171	0.171
DEU	0.077	0.083	0.092	0.100	0.106	0.114	0.124	0.130
GBR	0.061	0.059	0.057	0.056	0.055	0.054	0.053	0.052
JPN	0.249	0.284	0.320	0.338	0.351	0.363	0.369	0.372
NZL	0.008	0.011	0.014	0.016	0.017	0.018	0.020	0.020
SWE	0.250	0.271	0.294	0.305	0.313	0.321	0.325	0.326

Notes: The test conducted by Hamidi Sahneh (2016) examines the null hypothesis of fundamentalness, which implies the non-predictability of the VAR residuals. The table presents the p-values for alternative maximum lags used to predict future innovation. The diagnostics are based on VARs of lag order 4 using the Parzen Kernel. Results for alternative kernels and residuals from VAR(12) models are available upon request. We express our gratitude to Mehdi Hamidi Sahneh for providing the relevant codes for this test.

D - Estimated structural parametermatrices

With the values in parentheses $(a; b)$ denoting the bootstrap means (a) and t-ratios (b) the estimated structural impact multipliers $\widehat{D}$ read for full sample information as follows:[23]

$$\widehat{D}_{AUS} = \begin{bmatrix} \underset{(0.325;5.204)}{0.379} & \underset{(0.005;0.046)}{0.002} & \underset{(0.018;1.758)}{0.045} \\ \underset{(-0.243;-0.088)}{-0.051} & \underset{(3.177;13.153)}{3.278} & \underset{(-0.072;-0.02)}{-0.002} \\ \underset{(-0.043;-1.4)}{-0.119} & \underset{(0.009;-0.479)}{-0.024} & \underset{(2.335;8.224)}{2.424} \end{bmatrix}, \widehat{D}_{CAN} = \begin{bmatrix} \underset{(0.329;5.187)}{0.381} & \underset{(-0.011;-0.847)}{-0.034} & \underset{(0.018;1.556)}{0.04} \\ \underset{(0.106;0.922)}{0.291} & \underset{(1.996;9.619)}{2.031} & \underset{(-0.136;-1.358)}{-0.112} \\ \underset{(-0.039;-1.035)}{-0.081} & \underset{(0.02;0.354)}{0.019} & \underset{(2.329;8.273)}{2.421} \end{bmatrix}$$

$$\widehat{D}_{CHE} = \begin{bmatrix} \underset{(0.33;5.244)}{0.384} & \underset{(0.003;-0.024)}{-0.001} & \underset{(0.017;1.397)}{0.036} \\ \underset{(0.662;1.764)}{0.851} & \underset{(3.045;17.767)}{3.091} & \underset{(-0.185;-0.814)}{-0.13} \\ \underset{(-0.019;-0.6)}{-0.047} & \underset{(-0.021;-0.052)}{-0.002} & \underset{(2.34;8.27)}{2.429} \end{bmatrix}, \widehat{D}_{DEU} = \begin{bmatrix} \underset{(0.33;5.213)}{0.384} & \underset{(0.01;-0.214)}{-0.009} & \underset{(0.02;1.386)}{0.036} \\ \underset{(0.431;1.618)}{0.695} & \underset{(2.884;17.521)}{2.909} & \underset{(-0.105;-0.278)}{-0.044} \\ \underset{(-0.03;-0.575)}{-0.043} & \underset{(-0.005;0.025)}{0.001} & \underset{(2.342;8.286)}{2.431} \end{bmatrix}$$

[23]For inferential purposes we use a Moving Block Bootstrap as suggested by Brüggemann et al. (2016). According to their recommendation the block length is set to 25 ($\approx 5.03\ T^{1/4}$). To improve the scaling of documented estimation results structural parameter estimates and bootstrap means are multiplied by 100.

$$\widehat{D}_{GBR} = \begin{bmatrix} \underset{(0.327;5.304)}{0.38} & \underset{(0.008;-0.518)}{-0.024} & \underset{(0.019;1.671)}{0.041} \\ \underset{(0.255;1.327)}{0.632} & \underset{(2.773;15.067)}{2.82} & \underset{(-0.114;-0.596)}{-0.084} \\ \underset{(-0.042;-1.115)}{-0.084} & \underset{(-0.052;-1.088)}{-0.052} & \underset{(2.338;8.379)}{2.43} \end{bmatrix}, \widehat{D}_{JPN} = \begin{bmatrix} \underset{(0.328;5.281)}{0.381} & \underset{(0.028;-0.199)}{-0.008} & \underset{(0.023;1.854)}{0.047} \\ \underset{(0.239;1.607)}{0.677} & \underset{(3.022;16.842)}{3.064} & \underset{(-0.056;0.152)}{0.018} \\ \underset{(-0.055;-1.422)}{-0.108} & \underset{(-0.035;-0.513)}{-0.022} & \underset{(2.347;8.184)}{2.439} \end{bmatrix},$$

$$\widehat{D}_{NZL} = \begin{bmatrix} \underset{(0.322;4.883)}{0.38} & \underset{(0.018;0.453)}{0.031} & \underset{(0.015;1.741)}{0.044} \\ \underset{(-0.392;-0.325)}{-0.292} & \underset{(3.347;10.91)}{3.522} & \underset{(-0.076;-0.017)}{-0.002} \\ \underset{(-0.016;-1.159)}{-0.101} & \underset{(0.022;-0.178)}{-0.01} & \underset{(2.328;8.371)}{2.42} \end{bmatrix}, \widehat{D}_{SWE} = \begin{bmatrix} \underset{(0.329;5.376)}{0.381} & \underset{(0.018;-0.516)}{-0.018} & \underset{(0.022;1.673)}{0.042} \\ \underset{(0.122;1.479)}{0.543} & \underset{(3.113;17.186)}{3.131} & \underset{(-0.069;-0.33)}{-0.062} \\ \underset{(-0.037;-1.106)}{-0.077} & \underset{(-0.035;-0.111)}{-0.005} & \underset{(2.346;8.289)}{2.438} \end{bmatrix}$$