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Interest Rates, Convenience Yields, and Inflation Expectations: Drivers of US Dollar Exchange Rates

Kerstin Bernoth[*] Helmut Herwartz[†] Lasse Trienens[‡]

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Abstract

Using a data-driven approach to identify structural vector autoregressive models, we examine key factors influencing the US dollar exchange rate across eight advanced economies from 1980 to 2022. We find that shocks to inflation expectations, which are closely tied to unfunded government transfer payments, have a pronounced effect on the US dollar's value. This underscores the fiscal dimension of exchange rates. External shocks, related to the convenience yield investors forgo to hold US dollar assets, have emerged over time as the most powerful driver of US dollar exchange rate fluctuations. These findings provide new insights into the complex interplay of monetary policy, fiscal dynamics, and global market forces in shaping US dollar exchange rates.

Keywords: exchange rates, convenience yield, inflation expectations, monetary policy, fiscal policy, unfunded government transfer payment, monetary-fiscal policy mix.

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

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1 Introduction

Since 2000, the global economy has undergone significant and far-reaching economic shifts. Financial markets worldwide have faced recurring bouts of heightened risk aversion, driven by a series of crises, including the 2008 Global Financial Crisis (GFC), the Covid-19 pandemic, and the energy shock following Russia's invasion of Ukraine. Simultaneously, government debt levels in the United States and other advanced economies have surged, sparking concerns about the long-term sustainability of fiscal policies and their potential effects on price stability. Since 2021, the Federal Reserve has responded to accelerating inflation by sharply tightening monetary policy, signaling a decisive shift from its previously accommodative approach. Given the central role of the US dollar in global trade, asset issuance and official reserves, understanding how it is affected by these economic developments is both timely and critical.

The macroeconomic literature on the responsiveness of the US dollar exchange rate is extensive and insightful, yet the majority of studies focus primarily on the marginal role of individual determinants. A triad of prominent drivers in the current context includes (i) US short-term interest rates governed by monetary policy (see, e.g., Eichenbaum & Evans, 1995; Faust & Rogers, 2003; Stavrakeva & Tang, 2019), (ii) the value international investors place on safe US dollar-denominated assets (i.e., convenience yields, see, e.g., Krishnamurthy & Vissing-Jorgensen, 2012; Krishnamurthy & Lustig, 2019; Rey, 2015; Miranda-Agrippino & Rey, 2022), and (iii) inflation expectations, which reflect public perceptions of the sustainability of fiscal debt service plans (see, e.g., Jiang, 2021a,b; Bianchi & Melosi, 2022; Schmitt-Grohé & Uribe, 2022). A key finding from the recent monetary policy literature (Müller et al., 2024) suggests that the effects of interest rate surprises on the US dollar exchange rate are likely time-contingent. While the informational content conveyed by central bank decisions (Nakamura & Steinsson, 2018; Gürkaynak et al., 2021) is identified as one possible explanation for this time variation (Müller et al., 2024), this triad of exchange rate determinants provides a complementary perspective. Recognizing that these drivers operate simultaneously raises concerns about empirical approaches focusing on single-factor effects, while also implying that the responsiveness of the US dollar varies over time as a reflection of the relative strength of these interacting causal forces.

This paper investigates those factors influencing the US dollar exchange rate, focusing on the intricate interactions between US monetary policy, global demand for safe-haven US dollar assets, and long-term US inflation expectations. For this purpose, we analyze the causal relationships between short-term US interest rates, the US dollar exchange rate, and long-term US inflation expectations, identifying three structural shocks that drive our model variables.[1] The first is a standard US short-term nominal interest rate shock.[2] The second shock is measured as exogenous innovations to long-term US inflation expectations. Following Bianchi et al. (2023a), Cochrane (2023), and Herwartz & Trienens (2024), we interpret this shock as a fiscally induced inflation shock resulting from uncovered changes in fiscal policy. Finally, like Bernoth & Herwartz (2021) and Cormun & De Leo (2022), we identify an external shock, defined as an exogenous change in the US dollar exchange rate. We demonstrate that this shock is associated with the US dollar convenience yield and, consequently, the global demand for safe US dollar-denominated assets.

To identify the structural shocks, we apply a data-based identification approach of structural vector autoregressive (SVAR) models that takes advantage of the uniqueness of independent components in linear non-Gaussian systems (Comon, 1994).[3] Model implied structural shocks have sound economic properties. The main advantage of this full-system identification approach is that it does not impose explicit restrictions on the behavior of our model variables while allowing for a full and simultaneous interaction between them. We then trace the dynamic responses of changes in US short-term interest rates, the US dollar exchange rate, and long-run inflation expectations.

Using a monthly data set covering the period 1980M1 to 2022M12 and a cross-section of

[1] Given the importance of the interest rate differential between the two countries in determining exchange rates, we also estimated the model with four shocks by adding a foreign nominal interest rate shock. However, it turns out that the US and foreign nominal interest rate shocks are highly correlated. Thus, for identification reasons, we refrain from adding the foreign nominal interest rate shock.

[2] Note that we are explicitly not talking about a monetary policy shock here. As elucidated by Müller et al. (2024) and Gürkaynak et al. (2021), among others, interest rate shocks encompass both a monetary policy shock and a central bank's proprietary insights regarding the real economy, such as the natural interest rate. However, since interest rates are important determinants of exchange rates and we aim for full model identification, we focus on interest rate shocks.

[3] Identification by means of independent components, as detected in this work, has been successfully employed in the context of US monetary policy analysis and exchange rate modelling (see, e.g., Bernoth & Herwartz, 2021; Jarociński, 2024; Herwartz et al., 2022b,c) and (Herwartz & Wang, 2023).

eight advanced economies, i.e. Australia, Canada, Germany, Japan, New Zealand, Sweden, Switzerland, and the United Kingdom, we find that external factors, US monetary and long-run US inflation expectations influence the US dollar exchange rate. In response to a positive interest rate shock, the US dollar tends to appreciate. An exogenous surge in inflation expectations, which we demonstrate to be closely associated with unfunded government transfers, results in a depreciation of the US dollar. This underscores the fiscal dimension of exchange rates. An external shock in the form of an unexpected appreciation of the US dollar, which can be linked to the value international investors place on liquid and safe dollar assets, leads to a persistent appreciation of the US dollar. The historical decomposition shows that all three shocks considered make an important contribution to explaining US dollar exchange rate changes, with external shocks being of somewhat greater importance on average.

As a robustness test, we estimate our model for three different sub-samples: the Volcker period, the pre-GFC period, and the post-GFC period. We find evidence of time dependence, primarily in the response of US short-term interest rates and the US dollar exchange rate to shocks to inflation expectations. During the Volcker era and in the post-crisis period, short-term US interest rates rose in response to a positive shock to long-term inflation expectations and the US dollar appreciated. We interpret this as the Fed having pursued an active monetary policy stance. Between the late 1980s and the onset of the GFC, there are indications that the US Federal Reserve adopted a more passive monetary stance, as evidenced by a decrease in US short-term interest rates, though this decrease was small. In the post-crisis period, this pattern has reversed, suggesting a return to an active monetary policy stance.

Our work is linked to a number of important areas of research. First, this paper contributes to the large body of research on the impact of monetary policy on exchange rates. Previous literature has used various assumptions to identify exogenous monetary policy shocks, which turn out to be too restrictive for the research question under investigation. 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.[4] Identification with

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

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., 2024) 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. The main advantage of the data-based identification approach used in this paper is that it does not impose explicit restrictions on the behavior of our model variables while allowing for a full and simultaneous interaction between them.

Second, this study contributes to the literature on the global financial cycle (Rey, 2015; Miranda-Agrippino & Rey, 2020) and research highlighting that the US dollar exchange rate is significantly influenced by global demand for safe US dollar assets and the role of the United States as the provider of the dominant global currency with safe-haven status (Bruno & Shin, 2015; Maggiori, 2017; Krishnamurthy & Lustig, 2019; Kalemli-Özcan, 2019; Gourinchas et al., 2010; Georgiadis et al., 2021; Ilzetzki & Jin, 2021; Jiang et al., 2021; Cormun & De Leo, 2022).

Third, by explicitly distinguishing between short-term interest rate shocks and shocks to long-run inflation expectations, it allows us to confirm an important result of Schmitt-Grohé & Uribe (2022), who demonstrate that, in contrast to a temporary monetary tightening, which leads to an appreciation of the US dollar, a persistent monetary shock in the form of an increase in long-run inflation expectations depreciates the US dollar.

Fourth, by arguing that our inflation expectations shock is related to concerns about large unfunded public spending and fiscal sustainability, our paper also contributes to the literature on the impact of fiscal policy on exchange rates. In the fiscal theory of the price level (FTPL) from an international perspective, Jiang (2021a) and Jiang (2021b) emphasize the pivotal role of the United States. While a deterioration in fiscal conditions in the US leads to a depreciation of the US dollar, fiscal conditions in other advanced economies have less significance for exchange rate developments.

Finally, we also contribute to the literature on monetary and fiscal interactions; the debate on monetary versus fiscal dominance; and how these regimes have changed over time by analyzing the response of US interest rates and exchange rate to inflation expectations shocks (Sargent & Wallace, 1981; Leeper, 1991; Sims, 1994; Woodford, 1995; Cochrane, 2001; Herwartz & Trienens, 2024).

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 identified shocks. Section 5 explores whether interest rates and exchange rates react differently over time. 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 fundamentalness (Appendix C), on the structural parameter estimates (Appendix D), and on regression results for US inflation expectations shocks and unfunded fiscal shocks identified by Bianchi et al. (2023a) (Appendix E).

2 Empirical model

2.1 Data

We analyze the causal relationship between short-term nominal interest rates, exchange rates, and long-term inflation expectations 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 2022M6.[5] 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, Swe-

[5] We also split the full sample information into three subsamples: the Volcker-era, a pre-crisis era, and post-crisis era to shed light on the eventually modified transmission of structural shocks after the GFC and the Great Recession (see Section 5).

den, 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 differs 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 Canada, Japan, and the United Kingdom. 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).

2.2 A cross-section of structural VARs

Conditional on presample values $y_0, y_1, ..., y_{1-p}$, we consider a set of eight country specific VARs of dimension $K = 3$. Omitting a country indexation for notational clarity, the models read in their reduced and structural form, respectively, as

$$y_t = \nu + A_1 y_{t-1} + \ldots + A_p y_{t-p} + u_t, \tag{1}$$

$$= \nu + A_1 y_{t-1} + \ldots + A_p y_{t-p} + D\epsilon_t, \; t = 1, 2, ..., T, \tag{2}$$

where ν is a vector of intercepts, $A_1, A_2, \ldots, A_p$ are $K \times K$ parameter matrices, and u_t is a serially uncorrelated vector process with mean zero and covariance Σ_u. By assumption, the model in (1) is causal, i.e., $\det(A(z)) \neq 0 \,\forall\, |z| \leq 1$, where $A(z) = I_K - A_1 z - A_2 z^2 - \ldots - A_p z^p$, such that the corresponding Wold representation reads as $Y_t = \Phi(L)u_t$, where L is the lag operator, e.g. $Ly_t = y_{t-1}$, and $\Phi(L) = A(L)^{-1}$. While the reduced form parameters ν, Σ, A_i with $i = 1, \ldots, p$ and the residuals u_t can be estimated by means of OLS consistently, the identification of the parameters d_{ij} of the nonsingular $K \times K$ structural mixing matrix D requires external, non-sample information. Accordingly, ϵ_t signify identified structural innovations with mean zero and unit covariance.

We consider three endogenous variables. The first is the US one year treasury bill rate, i_t. 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 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. The second is the log nominal exchange rate in foreign currency per US dollar, respectively, s_t. And third, $\hat{\pi}_t$ serves as an indicator of fiscally induced inflation (Bianchi et al., 2023a) or, alternatively, as an indicator of persistent shifts in monetary policy (Uribe, 2022, see a more detailed discussion on this in section 3.1.3). Like Lukmanova & Rabitsch (2020), we use here the mean of US inflation expectations for the next ten years from the Survey of Professional Forecasters (SPF).[6] The sampled variables i_t, s_t, and $\hat{\pi}_t$ are not cointegrated according to conventional diagnostics, so we estimate the model in first differences. With Δ denoting the first difference operator, i.e. $\Delta = 1 - L$, the vector of endogenous variables is $y_t = (\Delta i_t, \Delta s_t, \Delta \pi_t^e)'$. Based on the AIC criterion, we use VAR orders of $p = 12$.

We also consider $K = 4$ dimensional models including foreign treasury yields, i_t^*. With regard to the three shocks of interest in this work, the informational content of four dimensional systems is similar to the one of trivariate models. For instance, regarding largest available samples for the UK, Japan, and Canada the correlations (i.e., $K = 3$ vs. $K = 4$) between model specific US short-term interest rate shocks are 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.

2.3 Identification based on the uniqueness of the 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

[6] 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.

matrix D in (2) is nonsingular.[7] Hence,

$$\epsilon_t = D^{-1}u_t \text{ and } \mathrm{Cov}[u_t] = DD' =: \Sigma_u. \tag{3}$$

It is well known that, in a Gaussian framework ($u_t \sim N(0, \Sigma_u)$), 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 (3) 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 (2024) explores the non-Gaussian properties of monetary policy shocks and uses independent component analysis 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 appears as a promising solution to achieve identification in a data-based manner.[8]

The data-based approach to identification that we pursue in this study consists of determining the country specific matrices D such 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).[9] While the use of

[7]We also follow the convention to investigate effects of positive structural shocks and assume that the diagonal elements of D are positive.

[8]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.

[9]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).

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 on the theoretical and empirical transmission channels that shape the contemporaneous relationships among short-term US yields, exchange rates, and long-term inflation expectations. This helps us plausibly identify the expected impact of exogenous shocks hitting the dynamic system of three endogenous variables under consideration.

Moreover, we exploit two particular merits that are specific to the present joint analysis of a cross-section of eight structural VARs. First, as the statistical identification scheme is fully agnostic and economic theory is supposed to apply to all advanced economies considered, finding qualitatively similar D matrices for the set of SVARs can be considered as stronger and 'cross-confirming' evidence in favor of a particular causal structure in comparison with single country models. Specifically, the mean group perspective might be used to explicitly test specific (joint) hypotheses on the structural parameters in D. Second, in the present analysis the country specific models comprise common US variables such that (some) economically identical shocks can be expected to drive the observable dynamics in country specific SVARs. In this regard, (very) high empirical correlations among shocks that we retrieve from distinct country specific models provide further support for the chosen economic labels and enhance the cross-confirming informational content of regarding a set of SVARs.

3 Shock labelling

This section begins with a review of the current theoretical and empirical literature that provides evidence on the links between short-term US nominal interest rates, the US dollar exchange rate, and long-term US inflation expectations. This guides us in plausibly determining the impact effects that one expects for the three exogenous shocks considered within the SVAR analysis. Subsequently, we utilize the identification approach described in Section 2.3 to extract the structural shocks that are later used in the impulse response exercises. Furthermore, we provide a more detailed analysis of the identified structural shocks to gain a better understanding of their interpretation and dynamics.

3.1 Expected effect directions based on a literature review

3.1.1 Short-term interest rate shock

The first shock considered is the temporary innovation to the short-term US nominal interest rate (*US IR shock*). The reaction of the US dollar exchange rate to an exogenous increase in US interest rates is not clear-cut in theory. There are theoretical arguments for effects in both directions. In a seminal paper, Dornbusch (2017) puts forth the argument that, in response to a contractionary monetary policy shock, the exchange rate initially appreciates, followed by a depreciation in subsequent periods, a phenomenon that has been dubbed the 'overshooting hypothesis'. Krishnamurthy & Lustig (2019) add the consideration that when the US 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 US Federal Reserve interest rate increases on the US dollar exchange rate. One consequence of higher interest rates is an upward shift in the debt service burden borne by 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 could invoke a depreciation of the US dollar.

The ambiguity of the impact of monetary policy on exchange rates is also reflected in the empirical literature. Several studies find an immediate positive relationship between the US dollar exchange rate (appreciation) and US interest rates (e.g. Müller et al. (2024), Rüth (2020), and Schmitt-Grohé & Uribe (2022)). It is also frequently observed that the exchange rate tends to appreciate further in subsequent periods, in contrast to the predictions of Dornbusch's hypothesis, and only begins to depreciate much later. This pattern is commonly referred to as the 'delayed overshooting puzzle' (Eichenbaum & Evans, 1995; Scholl & Uhlig, 2008). Stavrakeva & Tang (2019) finds even the opposite that the exchange rate depreciates in response to a monetary policy tightening shock. The author attributes this to the signaling effect of monetary policy dominating in times of crisis. An unexpected tightening of US monetary policy signals economic strength, leading to a decline in risk aversion and higher expected inflation in the US. Inoue & Rossi (2019) add 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.

Finally, as argued by Müller et al. (2024) and Nakamura & Steinsson (2018), markets are unable to discern whether an unanticipated increase in the policy rate is the result of a monetary policy shock or a rise in the natural rate of interest. The former has an appreciating effect on the exchange rate, whereas the latter has a depreciating effect. Thus, the overall exchange rate effect is contingent upon market perceptions. Nakamura & Steinsson (2018) identify that one-third of interest rate surprises are attributable to monetary policy shocks, with the remaining two-thirds attributable to innovations in the natural rate.[10] Consequently, without explicitly differentiating between the two, estimates of US interest rate shocks could encompass an aggregation of both structural movements and demonstrate an insignificant exchange rate response. Thus, the response of the US dollar exchange rate to an interest rate shock might be state-dependent and, consequently, the theoretical sign is left open.

Economic theory suggests that the response of long-run inflation expectations to a contractionary interest rate shock is ambiguous. 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, when households can distinguish between monetary and different types of financial shocks, or negative under imperfect information. However, the response could also be positive. In line with the FTPL, monetary tightening destabilizes the present value of future surpluses. In the absence of sufficient fiscal

[10] In a related manner, Kekre & Lenel (2024) find that natural rate shocks dominate the variance of US dollar exchange rates.

adjustment, combined with high public debt levels, concerns about the sustainability of public finances may arise, leading households to reduce their asset holdings and increase their cash holdings. The resulting increase in liquidity puts upward pressure on current inflation, thereby rebalancing the real value of debt with the diminished present value of surpluses. In the presence of price rigidities, this process would translate into an increase in trend inflation (Cochrane, 2023). Moreover, Cochrane (2001, 2023) finds that an increasing average maturity structure of public debt amplifies the response of expected inflation to monetary shocks.

The empirical literature confirms that the response of long-run inflation expectations to a contractionary monetary policy stance is not clear-cut. Lukmanova & Wouters (2022) find that the (perceived) inflation target proxied by long-term inflation expectations reacts negatively to a restrictive monetary policy shock and returns to its initial level. Thus, the effect on long-run inflation is neutral. Nakamura & Steinsson (2018) estimate that break-even inflation implied by 2-year and 3-year forward rates responds positively, while break-even inflation based on 5-year and 10-year forward rates responds significantly negatively to an unexpected change in interest rates. Beechey et al. (2011) and Gürkaynak et al. (2010) confirm that long-term inflation expectations decline in response to a contractionary monetary policy announcements. Allowing for potential state dependence, we leave the theoretical sign of the expected response of long-run inflation expectations to an unexpected rise in interest rates in Table 1 open.

3.1.2 External shock

The second shock considered in this study measures an unexpected change in the US dollar exchange rate, which we refer to as the *external shock*, analogous to Cormun & De Leo (2022). As we explain in detail in section 3.3.1, the external shock is closely linked to measures for the convenience yield of US treasury securities and, consequently, global risk. This is also confirmed by Corbo & Di Casola (2022), who interpret 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. The theoretical model by Lukmanova & Wouters (2022) supports the negative response of the short-term interest rate to external finance premium shocks. Thus, the theoretical impact effect of an external 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 shock to the US dollar affects long-run inflation expectations. Lukmanova & Wouters (2022) model three types of financial shocks into the Taylor rule, i.e. a risk premium, term premium, and external finance premium shock. Their external finance premium shock comes closest to the external shock identified in our estimations. They show that, under information frictions, when households cannot distinguish between monetary and different types of financial shocks, ten-year inflation expectations immediately and significantly decline after an external financial premium shock, subsequently reverting to its initial level. Orlowski & Soper (2019) and Netsunajev & Winkelmann (2014) analyze the interactions between global market risk, represented by the VIX, and long-term US inflation expectations. Both find that a positive shock to the VIX significantly dampens inflation expectations. This negative relationship becomes particularly pronounced at turbulent market periods or crises, which are accompanied by expectations of disinflation and economic weakness. Therefore, the expected theoretical response of long-run US inflation expectations to an external shock is zero or negative in Table 1.

3.1.3 Inflation expectations shock

The third shock considered is a US long-term inflation expectations shock (*US IE shock*). There are two theories in the academic literature as to the origin of this shock. First, researchers such as Mumtaz & Theodoridis (2018); Uribe (2022); Schmitt-Grohé & Uribe (2022) and Lukmanova & Rabitsch (2020) argue that trend inflation is driven by an inflation targeting shock induced by the monetary authority. This is also pointed out by Nautz et al. (2019), who show that shifts in US long-term inflation expectations are often the result of changes in expectations about the Federal Reserve's inflation target. This view is consistent with the view that "inflation is always and everywhere a monetary phenomenon" (Friedman, 1963). Second, the core principle of the FTPL is that inflation aligns the market value of government debt with the present value of primary surpluses. Consequently, in addition to monetary policy, fiscal policy also plays an important role in determining inflation. Large fiscal imbalances combined with a weakening fiscal (funding) credibility may lead trend inflation to increase, as argued for example, by Leeper (1991); Woodford (1995, 2001); Sims (1994); Dupor (2000); Bassetto (2002); Gómez-Cram et al. (2024) and Cochrane (2001, 2022b,a). Recent studies by Bianchi et al. (2023a) for the US and Barro & Bianchi (2023) for a cross-section of advanced economies support the view that "persistently high inflation is always and everywhere a fiscal phenomenon" (Sargent, 2013), suggesting that shocks to unfunded fiscal transfers are the main drivers of trend inflation and long-run inflation expectations. Building on this perspective, the work by Herwartz & Trienens (2024) shows that inflation-targeting shocks, as identified by Uribe (2022), are closely linked to fiscal policy. In section 3.3.2, we argue that the shock to long-run inflation expectations in the US is significantly related to US fiscal policy.[11]

Theoretical models do not provide a clear indication of the direction in which short-term US interest rates are expected to react to a shock to long-term inflation expectations. In particular, it may depend on the precise design of the monetary reaction function but also on the degree of fiscal backing (e.g., Leeper, 1991; Woodford, 1995; Leeper, 2013; De Michelis & Iacoviello, 2016; Gómez-Cram et al., 2024; Smets & Wouters, 2024). For instance, an independent central bank will raise policy rates in response to a positive shock to long-term inflation expectations (so-called active monetary policy). Herwartz & Trienens (2024) discuss that, under active monetary policy, the resulting downward pressure on the present value of surpluses requires fiscal backing to enable the central bank to set the interest rate above the inflation response. Otherwise, increased interest payments fuel fiscal sustainability concerns and cause households to reduce their bond holdings, thereby amplifying the increase in consumption. The additional rise in inflation, in turn, creates a one-to-one comovement in inflation and yields (i.e. the neo-Fisherian effect), that prevents the central bank from

[11] According to the analysis by Bernanke (2003) and the stepping on a rake hypothesis by Sims (2011), monetary shocks without fiscal backing destabilize the present value of surpluses, thereby also fostering a change in inflation expectations to revert the implied budget constraint violation. Nonetheless, according to the results of Bianchi et al. (2023a), fiscal policy dominates the occurrence of these violations in post-WWII data, leading us to interpret these shocks as fiscally induced US inflation expectations shocks.

stimulating real interest rates. In policy regimes with intermediate active and passive fiscal policies, as discussed by Smets & Wouters (2024), nominal yields rise modestly and effectively reverse the rise in inflation (expectations) when public transfers increase with a high degree of fiscal backing. Conversely if the US Federal Reserve were to respond to a rise in inflation expectations by raising interest rates without significant fiscal backing, this could lead to an uncontrollable inflationary spiral and push the economy into recession, as pointed out by Bianchi & Ilut (2017) and Cochrane (2022c). Conversely, a shock to long-term inflation expectations combined with increased political pressure, serious concerns about fiscal sustainability, economic growth or financial stability may lead the central bank to be more passive, resulting in an immediate neutral to negative policy rate response (so-called passive monetary policy).

The empirical literature is even more inconclusive as to how a shock to inflation expectations affects short-term interest rates. Uribe (2022) and Schmitt-Grohé & Uribe (2022) show an initially neutral reaction of short-term yields to changes in inflation expectations, which then gradually becomes positive. Lukmanova & Rabitsch (2020) find an initial negative impact on short-term nominal US interest rates, which turns positive after a few quarters, when focusing on the post-2008 phase. Using data from 1960 to 2007, Bianchi et al. (2023a) find an immediate negative response of the Fed funds rate to unfunded fiscal shocks that persists for five years.[12] Bianchi & Melosi (2017) show that short-term yields had an immediate positive response to fiscal imbalances before 2008, which turned negative thereafter. Bianchi et al. (2023b) find a negative interest rate response under political pressure before the appointment of Paul Volcker as US Fed Chairman in the early 1980s, as well as during the Donald Trump administration. Herwartz & Trienens (2024) show that a positive interest rate response only occurs in a high yield environment, possibly due to a monetary authority that follows the Taylor rule. In a low interest rate environment, they find, in line with Bianchi & Melosi (2017) and the FTPL, that if a central bank is passive (e.g. due to political pressure, economic growth concerns, financial stability, or fiscal sustainability concerns), this leads to a muted, possibly even negative, reaction of short-term interest rates after inflation expectations shocks. All in all, we conclude from this literature review that

[12] Unfunded fiscal shocks are defined as shocks to transfers that are not backed by future fiscal adjustments.

the reaction of short-term nominal interest rates to an inflation expectations shock is state- and time-dependent and that we cannot make a clear statement about the direction.

The response of exchange rates to a positive shock to US long-term inflation expectations should depend, on the one hand, on the response of currency risk premia and, on the other hand, on the response of interest rates, as suggested by purchasing power parity. Empirical estimates by Schmitt-Grohé & Uribe (2022) find an immediate depreciation, to which they refer as the neo-Fisher effect of the open economy. Using a VAR model for Japan, De Michelis & Iacoviello (2016) find that inflation-targeting shocks result in a temporary real appreciation (depreciation) when interest rates rise faster (slower) than inflation. In the international context, the FTPL literature provides support for nominal and real depreciation effects when inflation expectations shocks are accompanied by public budget concerns (see, e.g., Jiang, 2021a,b). However, since the impact of a long-run inflation expectations 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 patterns behind our shocks. Given that we consider a set of country specific SVARs, it is worth pointing out that the cross-section dimension does not just enable mean group estimation and inference for core parameters of interest. In addition, from the model outline it is evident that two shocks (i.e. the short-term interest rate shock and the inflation expectations shock) should be present in each SVAR. Therefore, empirical correlations among shocks retrieved from the set of SVARs convey useful information to cross-confirm the economic labels attached to empirical shocks.

Table 1: Theoretical sign patterns of structural shocks

Variable	US IR shock	External shock	US IE shock
i_t	$+$	$-$	?
s_t	?	$+$	?
π_t^*	?	$0/-$	$+$

3.2 Estimated effect directions and identification

As determined in this study, the uniqueness of the identified structural shocks $\epsilon_{k,t}$ only holds under informative deviations from the joint Gaussian model. In addition, model implied structural impulse response estimates 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.

While any structural analysis must rely on (debatable) external information - in our case the distributional features of latent shocks - the cross-sectional approach adopted in this work holds particular merits for a reliable solution of the identification problem. Notably, at the level of a *single* VAR, the data based approach to identification is fully agnostic and deserves an additional step of plausible and theory guided shock labelling. Since theoretical insights should not be specific to single models or economies, it is worth underpinning that finding qualitatively similar response profiles from agnostic identification for a *set of economies* cross confirms both, the identifying assumptions and the structural model implications. Noting that our analysis covers cross-sectional results for a total of eight structural models, in a nutshell, it is interesting to unravel in how far the unrestricted data-based estimates allow for a cross-sectionally (almost) uniform interpretation.

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 $h = 1$ up to horizon $h = 3$ (right hand side panel). Note that the impact effects ($h = 0$) along the diagonal are uniformly positive by construction. Thus, it is worth noting that the corresponding effects at horizons $h = 1, 2, 3$ are also throughout positive, on average, although this is not imposed within the identification step.

Several a priori unrestricted structural parameter estimates imply effect directions that are (almost) common for the entire cross-section. This can be interpreted as cross-confirmation of identification outcomes achieved by considering a set of SVARs rather than a single model analysis. The discussion of theoretical impact effect directions (see Table 1) reveals that we consider, in particular, the marginal response of US interest rates and US inflation expectations to an external shock as important for a sound economic labelling of the statistically identified shocks. In line with the theoretical impact effect directions, the empirical estimates of the structural model parameters d_{12} and d_{32} confirm negative impact effects of the external shock on US short-term interest rates and inflation expectations, respectively. Another interesting common result for off-diagonal estimates is that the responses of interest rates to shocks to long-term US inflation expectations are positive for seven and eight economies when looking on impact and within-quarter effects, suggesting that the US Fed is targeting inflation, on average.

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 yields structural shocks featuring sound economic labels.

Table 2: Empirical impact directions of structural shocks

	On Impact ($h = 0$)			Within one quarter ($h = 1, 2, 3$)		
	US IR shock	External shock	US IE shock	US IR shock	External shock	US IE shock
i_t	$+(8) - (0)$	$+(1) - (7)$	$+(7) - (1)$	$+(8) - (0)$	$+(1) - (7)$	$+(8) - (0)$
s_t	$+(7) - (1)$	$+(8) - (0)$	$+(2) - (6)$	$+(7) - (1)$	$+(8) - (0)$	$+(0) - (8)$
π_t^e	$+(3) - (5)$	$+(2) - (6)$	$+(8) - (0)$	$+(8) - (0)$	$+(0) - (8)$	$+(8) - (0)$

Notes: The table shows the absolute number of 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. *US IR shock* stands for short-term US interest rate shock, *External shock* for the exogenous innovations of the US dollar exchange rate and *US IR shock* for the long-term US inflation expectations shock.

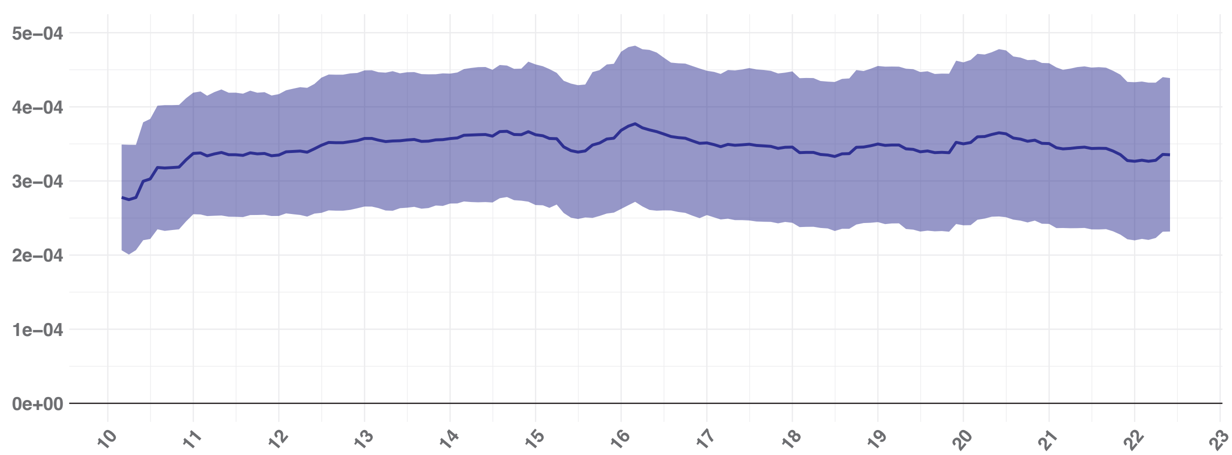
3.3 Interpreting external and inflation expectations shocks

3.3.1 External shocks, global risk and US dollar convenience yields

Shifts in the demand and supply of safe dollar assets have been considered as important drivers of fluctuations in US dollar exchange rates (e.g. Krishnamurthy & Vissing-Jorgensen, 2012; Engel & Wu, 2018; Krishnamurthy & Lustig, 2019; Jiang et al., 2021). 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 factors and safety demand, 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.

In light of the aforementioned considerations, we investigate the extent to which the identified external shock is associated with the attractiveness of US dollar assets to international investors. For this purpose, we consider the return that investors are willing to forego in order to hold safe dollar assets, which is also known as the convenience yield of US dollar securities (Du et al., 2018; Krishnamurthy & Lustig, 2019). Krishnamurthy & Lustig (2019) show that whenever there is a crisis in global financial markets, the convenience yield on dollar safe assets increases persistently. Cormun & De Leo (2022) provide further evidence that the US dollar convenience yield is closely associated with global risk aversion. According to Du et al. (2018) and Krishnamurthy & Lustig (2019), the convenience yield on US Treasury securities can be proxied by the US Treasury basis. Figure 1 shows outcomes of rolling regressions in the spirit of Lilley et al. (2022) for the effects of the external shocks on changes of the US Treasury basis against the G10 economies (ΔTB_t). The treasury basis is

Figure 1: Historical effects of external shocks on the US Treasury basis



Notes: The figure shows the estimated slope coefficients for rolling window regressions

$$-\Delta TB_s = \alpha + \beta \epsilon_{2,s} + u_s, \; s = \tau_1, \tau_1 + 1, \ldots, \tau_2,$$

where $\epsilon_{2,s}$ is the external shock and τ_1 (τ_2) is the lower (upper) bound of rolling samples of size 180 months. Owing to data availability, the regression starts in 1995M3 and ends in 2022M6. The solid line shows the arithmetic mean of point estimates, while the shaded areas represent approximate pointwise confidence intervals with 95% coverage.

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 bonds with same maturity of the G10 countries. To facilitate the interpretation of rolling regression results, we multiply the treasury basis with minus unity ($-\Delta TB$), such that positive values proxy a positive US dollar convenience yield.

We find a relatively stable, significantly positive relationship between our identified external shock and the US dollar convenience yields. Thus, an unexpected US dollar appreciation is also associated with an increase in returns that international investors are willing to forgo to hold safe US dollar assets. We interpret this as evidence that the role of the US dollar as the primary global safe-haven asset, and, hence, also global risk aversion is a strong driver of US dollar exchange rate developments.

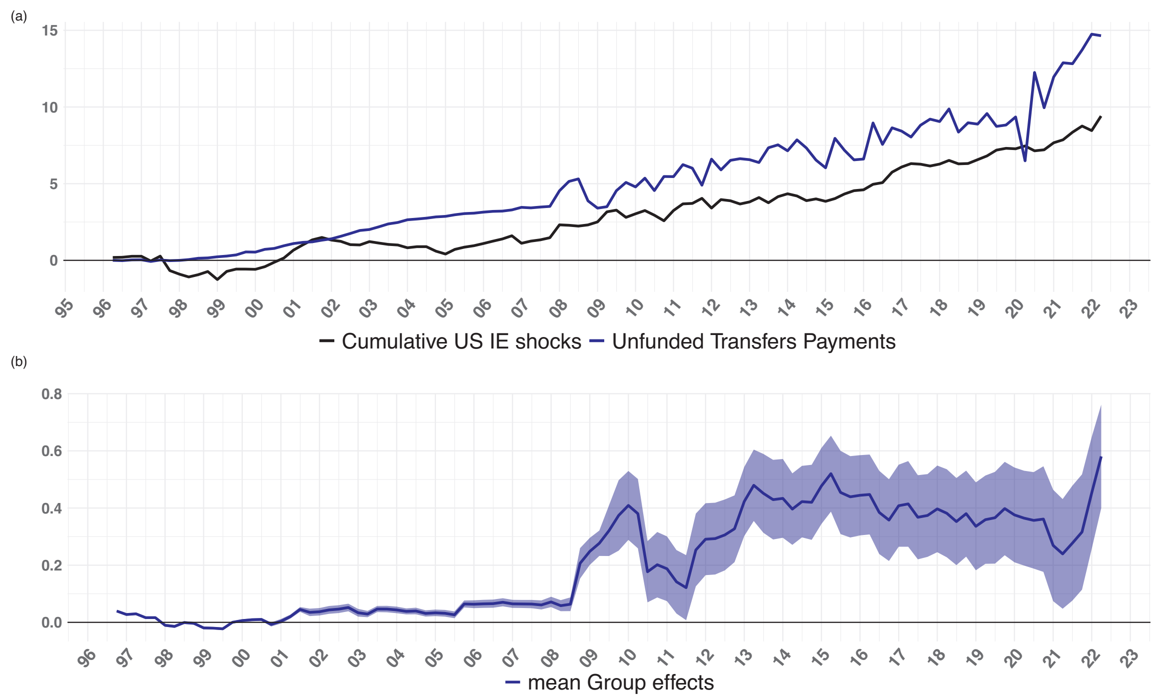
3.3.2 Inflation expectations shocks and unfunded fiscal policy

In this section, we show that the identified shock to US inflation expectations is closely related to uncovered US fiscal expansion. This supports the finding of Bianchi et al. (2023a) that, consistent with the FTPL, shocks in uncovered public transfers have been a key driver of persistent inflation in the US. Our finding also aligns with Bernanke (2003), Leeper & Leith (2016), Cochrane (2023), Herwartz & Trienens (2024), Smets & Wouters (2024), and Gómez-Cram et al. (2024) arguing that trend inflation and inflation expectations are determined not only by monetary policy but also by fiscal policy.

We again consider rolling window regressions to analyze the relationship between shocks to US inflation expectations in period s and the cumulated change in unfunded transfer payments in periods $s-1$, s, and $s+1$. To test robustness, we repeat the regression using shocks to unfunded transfers instead. Both fiscal variables are derived from the theoretical model of Bianchi et al. (2023a).[13] A positive regression coefficient implies that a positive US inflation expectation shock is associated with a cumulative increase in unfunded transfers/unfunded transfer shocks over these three periods. The reason why we also consider the lead and lag of the fiscal variables is to allow that in some episodes US IE shocks lead, while in others unfunded fiscal payments/shocks take precedence.

[13] We thank Francesco Bianchi, Renato Faccini, and Leonardo Melosi for providing us with both time series.

Figure 2: Historical effects of US inflation expectations shocks on unfunded transfer payments



Notes: Panel (a) displays cumulative US IE shocks alongside unfunded transfer payments (in levels and billions of US dollars) from 1996Q2 to 2022Q2. For better comparison, we adjusted the latter by subtracting its initial value. Panel (b) shows estimated slope coefficients for rolling window regressions

$$UF_t = \alpha + \beta\epsilon_{3,s} + u_s,\ \ s = \tau_1, \tau_1 + 1, \ldots, \tau_2,$$

where $\epsilon_{3,s}$ is the US IE shock, aggregated to quarterly data, UF_t is the sum of changes in unfunded transfer payments in $s-1$, s, and $s+1$, and τ_1 (τ_2) is the lower (upper) bound of rolling samples of size 60 quarters. Owing to data availability, the regression starts in 1981Q2 and end in 2022Q2. The solid line shows the arithmetic mean of point estimates across the eight advanced economies, while the shaded areas represent approximate pointwise confidence intervals with 95% coverage.

Panel (a) of Figure 2 shows the cumulative innovations of US IE shocks joint with unfunded transfer payments. Panel (b) displays the rolling window regression results of US IE shocks with changes in unfunded transfer payments.[14] In the late 1990s, we estimate a rather neutral relationship between US IE and changes in unfunded government transfers. This result can be explained by the fact that both US IE shocks and unfunded fiscal transfers showed little

[14]Figure 6 in Appendix E shows the equivalent results using US unfunded transfer shocks instead of unfunded transfer payments.

variation during this period (panel (a)). As argued by Chen et al. (2022), from 1995 to the early 2000s, the fiscal authority adjusted its budget sufficiently to stabilize its debt with surpluses, creating predominantly funded fiscal expansions. Moreover, long-term inflation expectations remained fairly stable during this period (Bernanke, 2003).

Beginning with the early 2000s, both the level of unfunded transfer payments and our cumulative shocks to US long-run inflation expectations show a positive trend (panel (a)). This is consistent with the findings of Chen et al. (2022), who also report a shift toward a more active US fiscal policy and less funded fiscal expansion during this period. This shift is likely influenced by the dot-com crisis and the 9/11 terrorist attacks. After the GFC and the COVID-19 pandemic, also the variation in both series increased. The rolling window regression coefficients (panel (b)) show that since the early 2000s, US IE shocks and changes in unfunded transfer payments exhibit a positive and significant relationship that became considerably stronger after the GFC.[15]

In sum, these results suggest that during periods of substantial variation in US IE shocks and unfunded transfer payments, a significant relationship between these two variables becomes evident.[16]

4 Structural impulse response analysis

Figure 3 shows the cumulative impulse responses functions (IRFs) to the identified shocks, i.e. the US interest rate shock (US IR shock), the external shock, and the shock to long-run US inflation expectations (US IE shock). The shocks are normalized by construction so that the impulse responses shown in this study reflect the effects of a structural shock of magnitude one. The horizontal axis measures time in months. The vertical axis measures the cumulative change in interest rates and inflation in annual percentage points ($0.05 \hat{=} 5\%$) and the cumulative change in the exchange rate (which is defined in log first differences) in percent.

[15] The close relation between IE shocks and fiscal policy is also reflected in an almost complete correlation of 0.96 between the level of unfunded transfer payments and a cumulation of the identified IE shocks. Moreover, cumulations of US IE shocks and unfunded transfer shocks as shown in Figure 6 in Appendix E also exhibit a high correlation of about 0.97

[16] Results in Appendix E further show that this finding holds consistently across US IE and unfunded transfer shocks.

Due to space considerations, we summarize estimation results by providing IRF estimates for all considered economies jointly. Note that this collection of estimates lacks a complementation with model specific confidence bands. Instead, we evaluate 'overall' significance of the displayed dynamics in terms of mean group criteria (Pesaran & Smith, 1995).

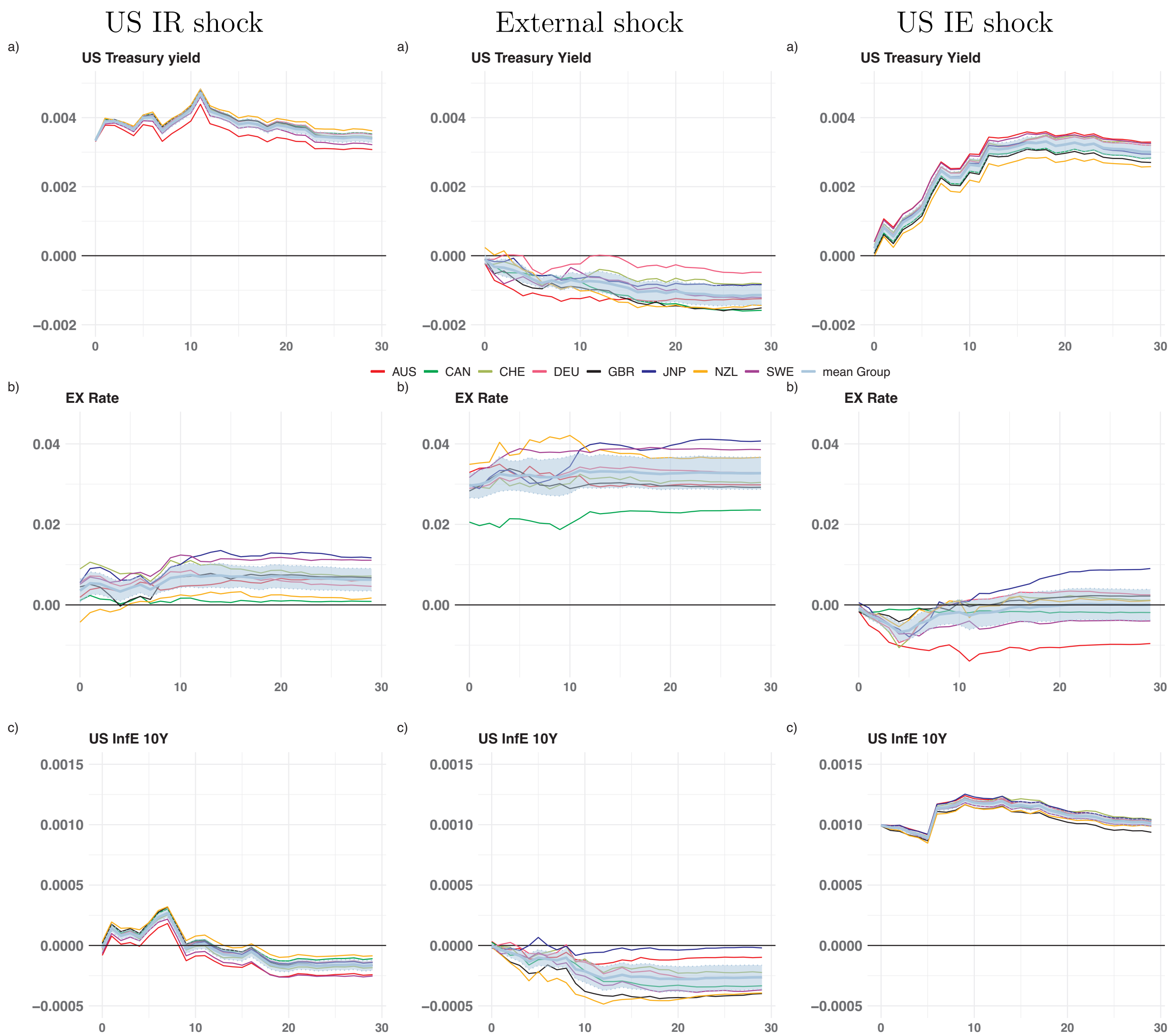
The first column of Figure 3 shows the response to a contractionary US interest rate shock. As previously noted, an increase in value represents an appreciation of the US dollar. We find that the US dollar appreciates immediately by, on average, half a percentage point (Panel (b)). After approximately one year does the US dollar starts to depreciate, albeit at a slow rate. This pattern corresponds to the delayed overshooting result established, for example, by Eichenbaum & Evans (1995). However, as will be demonstrated in Section 5, this result is significantly dependent on the analyzed time period.

We estimate that US inflation expectations initially increase in the first year following a contractionary US interest rate shock (panel (c)). However, they subsequently decline in accordance with our prediction and the findings of previous studies, including those by Beechey et al. (2011), Gürkaynak et al. (2010), Nakamura & Steinsson (2018), and Lukmanova & Wouters (2022). The initial rise may be attributed to the necessity for markets to ascertain whether the interest rate shock is indicative of a monetary policy shock or a shock to the neutral interest rate (Müller et al., 2024). The former is presumed to exert a negative impact on inflation expectations, whereas the latter is expected to have a positive effect.

The second column of Figure 3 shows the responses to an external shock in the form of an unexpected appreciation of the US dollar exchange rate. In line with our prediction, the cross-sectional analysis shows a significant and immediate decline in US short-term interest rates, which becomes more pronounced in the first few months and persists thereafter (panel (a)). The explanation is that the external shock is closely associated with an increase in the willingness of international investors to hold US dollar-denominated assets and that depresses US yields. Thus, we confirm the theoretical finding of Lukmanova & Wouters (2022).

US dollar exchange rates remain at the elevated level following an external shock as shown in panel (b). Inflation expectations in the US decline significantly and persistently in response to a positive external shock (panel (c)), which is consistent with the findings of Lukmanova & Wouters (2022), Orlowski & Soper (2019), and Netsunajev & Winkelmann (2014).

Figure 3: Full-sample mean group IRFs, 1980M1-2022M6



Notes: Cumulative responses to US interest rate shocks (*US IR shocks*, external shocks, and US IE shocks) up to an horizon of $H = 30$ months. Endogenous variables are US Treasury yields (12M), nominal EX rates, and US inflation expectations (10Y). Sample periods are 1980M1-2022M6 (AUS, CAN, CHE, DEU, GBR, JPN, NZL, SWE). Shaded areas show inferential results for mean group estimators, i.e. pointwise intervals covering arithmetic means $\pm$ 2 standard deviations of the mean group estimator. The shocks are normalized by construction so that the impulse responses shown in this figure reflect the effects of a structural shock of magnitude 1. The horizontal axis measures time in months. The vertical axis measures changes in interest rates and inflation in annual percentage points (e.g., $0.05 \mathrel{\hat{=}} 5\%$), and changes in the exchange rate in percent.

Thus, as explained by Orlowski & Soper (2019), the increased global risk associated with the external shock raises concerns about economic weakness and disinflationary tendencies.

The third column of Figure 3 shows the responses to a positive shock to long-term inflation expectations that we consider to be associated with unfunded fiscal expansion and the fiscal authority's credibility to stabilize large debt. We find that US short-term interest rates rise in the first 12 months and remain at elevated levels thereafter (panel (a)). Similar as Gómez-Cram et al. (2024), we interpret this as an indication of an active monetary policy stance, i.e. a central bank with an inflation targeting reaction function. However, as explained later in Section 5, we find that the response of short-term interest rates varies considerably for the different sub-samples considered. This is in line with Bianchi & Melosi (2017), Bianchi & Ilut (2017), Herwartz & Trienens (2024), Gómez-Cram et al. (2024), and Smets & Wouters (2024), who show that the monetary reaction function changes over time in the wake of different political and institutional constellations.

As shown in panel (b), we find, on average, a significant immediate depreciation of bilateral US dollar exchange rates in response to a positive shock to US inflation expectations, which is in line with the result of Schmitt-Grohé & Uribe (2022) and Lukmanova & Rabitsch (2020). After a few months, this depreciation disappears. Given the above evidence that long-run inflation expectations are significantly influenced by, among other things, unfunded fiscal expansion and the fiscal authority's credibility in stabilizing large public debt, our result confirms the hypothesis that the US dollar exchange rate also has an important fiscal policy component.

In conclusion, US dollar exchange rates are impacted not only by US monetary policy, reflected in US interest rates, they are also affected by external/global factors such as the convenience yield of US dollar assets and long-run inflation expectations that are determined by, among other factors, fiscal sustainability considerations.

In terms of information efficiency, it is worth noting that we have attributed the main drivers of US dollar exchange rates to structural shocks that are well defined and confirmed in the cross-section. Therefore, it is unlikely that the marginal reaction profiles discussed suffer from biases resulting from contamination of the identified shocks, which may be one reason for the rather inconclusive empirical results of the drivers of the US exchange rate in the existing literature so far. Moreover, beyond the stylized orthogonality conditions, these shocks can also be interpreted as independent components. Therefore, the described identification of the main determinants of US dollar rates is also well immunized against higher order dependencies between the structural origins of interest rates, inflation expectations, and international yield opportunities and associated risks.

5 Sensitivity tests - Sub-sample results

There is evidence in the literature that the responses of interest rates, inflation expectations, and exchange rates to macroeconomic shocks depend on the sample period. In the context of short-run interest rate responses, the existing literature on the FTPL suggests that the response of interest rates to inflation expectations shocks is contingent upon the monetary reaction function, resulting in state-dependent interest rate responses (Sargent & Wallace, 1981; Leeper, 1991; Bianchi & Melosi, 2017; Herwartz & Trienens, 2024; Smets & Wouters, 2024; Gómez-Cram et al., 2024). Further, there is evidence that the response of exchange rates to macroeconomic shocks depends on the sample period. For example, Eichenbaum & Evans (1995), using data from January 1974 to 1990, report significant delayed overshooting of exchange rates following monetary shocks, while Scholl & Uhlig (2008), using a sample from 1975 to 2002, find a more subdued pattern of exchange rates. This finding is consistent with the conclusion of Kim et al. (2017) that the delayed overshooting puzzle is a phenomenon that emerged predominantly during the 1980s, and is related to the Volcker era. Likewise, Bernoth et al. (2022) find a structural break in the magnitude of US dollar excess returns in 2007 with the onset of the GFC, which may also indicate a change in the response of the US dollar exchange rate to macroeconomic shocks.

Thus, this section examines whether the shock responses of US interest rates, inflation expectations, and US dollar exchange rates of the eight considered economies are indeed time varying. For this purpose, we repeat our SVAR estimations for three sub-samples: the Volcker-period from 1980M1 to 1987M7, a pre-crisis period from 1987M8 to 2007M4, and a post-crisis period from 2008M4 to 2022M6. Figure 4 shows the results.[17]

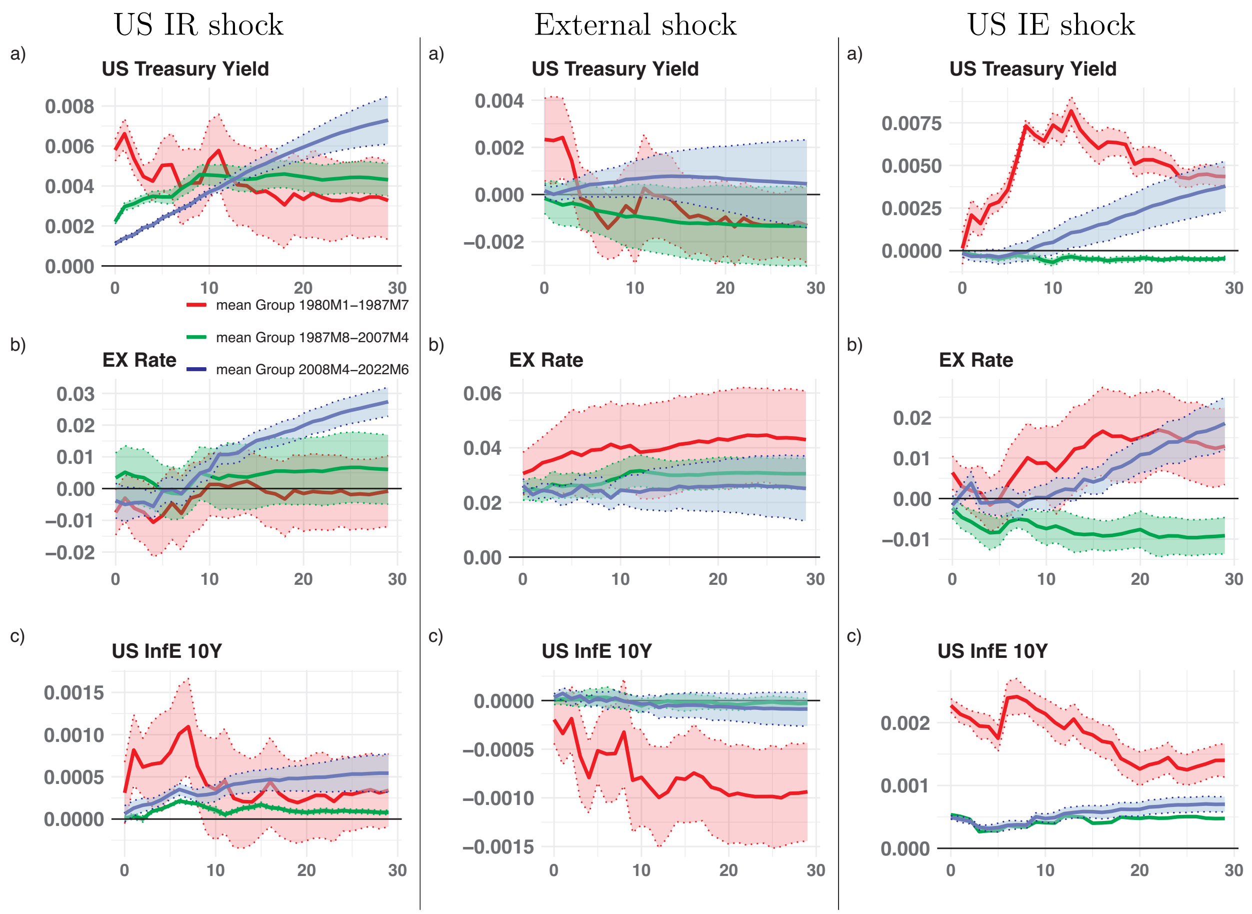
Due to space considerations, we describe only those estimation results for which we find significant differences across the three sub-periods under examination. As shown in the first column of Figure 4, only in the post-crisis period do we confirm our earlier result that the US dollar significantly appreciates in response to a contractionary US interest rate shock. For the pre-crisis period, we also find an US dollar appreciation, however, this response is not significant. During the Volcker era, the exchange rate response fluctuates around zero and is insignificant throughout. A suitable explanation is discussed above. According to Müller et al. (2024), the US interest rate shock can be interpreted as an aggregate of monetary and natural interest rate shocks. While the former is generally considered to have an appreciating effect on exchange rate, the latter has a depreciating effect. Our sub-sample IRFs indicate that, in the post-crisis period, the US interest rate shock is dominated by monetary policy shocks, while it represents a mix of the two in the two other sub-samples.

As illustrated in the second column of Figure 2, the response pattern of inflation expectations to external shocks also differs significantly across the three considered time periods (panel (c)). This finding is consistent with that of Orlowski & Soper (2019), who also report that the impact of market risk on US inflation expectations varies over time. Restricting observations to the pre- and post-crisis periods reveals that the reaction of US long-run inflation expectations to external shocks is insignificant over the full 30-month horizon. However, during the Volcker period, the full-sample result is confirmed, namely that inflation expectations largely declined in response to external shocks. An explanation is that, in this period, an increased demand in safe US dollar denominated assets was accompanied by expectations of disinflation and economic weakness.

We also find time dependence in the response of US short-term nominal yields and the US dollar exchange rate to US inflation expectations shocks (panels (a) and (b), third column). During the Volcker era and after the GFC, US short-term yields responded positively

[17] As we demonstrate in Section 3.3.1, the nature of exchange rate dynamics is subject to a transition process starting from the late 1980s as well as mid-2007. To capture only the new external variation, we exclude the transition period and leave a one-year gap between the two sub-samples. Following the findings of Eichenbaum & Evans (1995) for the Volcker sample, we truncate our sample just before 1990.

Figure 4: Sub-samples mean group IRFs, 1980M1-2022M6



Notes: The figure displays mean group results conditional on three sub-samples for responses to US interest rate shocks (first column), external shocks (second), and US inflation expectations shocks (third). Red-colored results are based on the sub-samples 1980M6-1987M7; green-colored results are conditional on subsamples 1987M78-2007M4; blue-colored results are conditional on sub-samples 2008M4-2022M6 (AUS, CAN, CHE, DEU, GBR, JPN, NZL, SWE). The endogenous variables considered are the US treasury yields with a one-year maturity, the bilateral exchange rates against the USD, and US inflation expectations (10Y).

to shocks to inflation expectations, reflecting a Taylor rule-based reaction function (Bianchi & Ilut, 2017; Herwartz & Trienens, 2024), albeit with a lag of a few months in the post-GFC period. Accordingly, the US dollar exchange rate appreciated significantly. However, in the pre-crisis period between 1987 and 2007, the estimates show a negative, albeit weak, response of US Treasury yields, on average. This is in line with Romer & Romer (2023), who argue that a recession in the early 1990s fostered a shift toward more passive monetary policies. Similarly, Cieslak et al. (2023) and Herwartz & Trienens (2024) also document a passive monetary stance in the early 1990s and 2000s. In line with the arguments put forth by Herwartz & Trienens (2024) and Gómez-Cram et al. (2024), this outcome is largely shaped by brief events like the dot-com crisis and 9/11, which were accompanied by a swift surge in budget deficits. During these periods, there is evidence that the US monetary authority mitigated an increase in the risk of government debt by decreasing interest payments. Accordingly, the median group estimates also point to a significant depreciation of the US dollar during the pre-crisis period. In the post-crisis period, the reaction of short-term interest rates to a surge in inflation expectations is once again significantly positive, and the US dollar also demonstrates a significant appreciation, albeit with a lag of several months. This suggests a return to a more active monetary policy stance.

6 Historical and forecast error variance decomposition

In this section, we analyze historical decompositions and forecast error variance decompositions (FEVD) to assess the relevance of the US interest rate shocks, external shocks, and US inflation expectations shocks for the variation in the bilateral US dollar exchange rate against the eight advanced economies under investigation. Given the finding of time-dependent effects in the previous section, we conduct both decompositions for each of the three considered sub-samples.

Figure 5 shows the average percentage contribution of the three shocks to absolute changes in the exchange rates. During the Volcker period, all three shocks considered are roughly equally important in explaining the development of US dollar exchange rates, even if there are slight variations over time (panel (a)).

In the pre-crisis period (panel b), the impact of external shocks has risen in comparison to the Volcker period, with a continued increase over the considered period. This shock accounts for approximately 50% of the variation in US dollar exchange rates. This points to the growing importance of external/global factors, such as the global demand for safe and liquid US dollar-denominated assets, in explaining the value of the US dollar, as evidenced in related studies (Krishnamurthy & Vissing-Jorgensen, 2012; Engel & Wu, 2018; Krishnamurthy & Lustig, 2019). The second most significant shock in the pre-crisis period is the US interest rate shock, which explains, on average, approximately 30% of exchange rate movements, while shocks to long-run inflation expectations account for about 20% of exchange rate movements.

Figure 5: Sub-samples avg. historical decomposition of exchange rate changes, 1980M1-2022M6

On average, external shocks were also the most important factor in explaining exchange rate dynamics in the post-crisis period. However, their relative importance has declined since 2020. Moreover, there were also shorter phases in which shocks to long-term inflation expectations had the strongest impact on the value of the US dollar. This can be observed, for instance, in the period following 2010, when uncertainty regarding the duration of the zero interest rate phase increased in the majority of advanced economies, thereby intensifying concerns about the long-term sustainability of public finances. Similarly, it is observable in the period following 2022, when there was a considerable rise in public deficits and debt, primarily in response to the energy crisis. Between 2008 and 2020, the importance of US interest rate shocks declines somewhat compared to previous subsamples, explaining only about 20% of exchange rate fluctuations. Since 2020, US interest rate shocks have regained importance in explaining US dollar exchange rates, which can be attributed to the rapid rise in official central bank interest rates in response to the rise in inflation in the countries analyzed.

Figure 7 in Appendix F shows subsample-specific mean group FEVD results for the bilateral exchange rates of the eight advanced economies under investigation. During the Volcker period, external shocks initially explained on average around 75% of the forecast error variance of US dollar exchange rates, dropping to around 55% after 30 months. US interest rate shocks accounted for more than 25% on average. On-impact, the contribution of shocks to US inflation expectations was rather small during in Volcker era, but the contribution increased at longer horizons to account for about 20% of the forecast error variances. In the pre-crisis period, external shocks continued to increase in relative importance in explaining US dollar exchange rates, while the contribution of inflation expectation shocks weakened in relative terms. For the post-crisis period, we detect a growing importance of external shocks, while US interest rate shocks lost importance. US inflation expectation shocks regained relevance for US dollar exchange rates, most likely due to the increased risk of US government debt and inflation following unfunded government spending during COVID-19 (see also Gómez-Cram et al., 2024).

7 Conclusion

Using monthly data for the period 1980M1 to 2022M6 and a cross-section of eight advanced economies – Australia, Canada, Germany, Japan, New Zealand, Sweden, Switzerland and the

United Kingdom – we analyze the exogenous drivers of US dollar exchange rates by means of agnostically identified structural VARs.

Our results indicate that a tightening of US interest rates results in an appreciation of the US dollar. Furthermore, we observe that the US dollar exchange rate remains at an elevated level following an external shock, measured as an unexpected appreciation of the US dollar. We show that this identified external shock is strongly related to the convenience yield - the premium that investors are willing to pay for holding US dollar assets. Consequently, lower global risk appetite and greater demand for safe US dollar assets are associated with a sustained appreciation of the US dollar. Finally, an unexpected rise in long-run inflation expectations in the US results in a depreciation of the US dollar. We demonstrate that US inflation expectations shocks are significantly correlated with US fiscal policy, particularly with a rise in unfunded public transfer payments. In this sense, the US dollar exchange rate also contains a significant fiscal component.

We find that all three identified shocks, the US interest rate shock, the external shock, and the long-run US inflation expectations shock, are important determinants of US dollar exchange rate dynamics. Among these, the external shock emerges as the predominant driver on average.

We observe a notable time-dependent pattern in how US short-term interest rates and dollar exchange rates respond to inflation expectation shocks, suggesting evolving monetary reaction functions over different periods. During the Volcker era, an unexpected rise in long-run inflation expectations typically led to increased US short-term interest rates and US dollar appreciation, indicative of an active monetary policy stance. In contrast, from the late 1980s to the early 2000s, comparable shocks gave rise to significant, albeit modest, reductions in US interest rates and considerable dollar depreciation. This shift in response patterns suggests a transition toward a more passive US monetary policy in the pre-crisis period. In the post-crisis period, this pattern has reversed, suggesting a return to an active monetary policy stance.

These results underscore the intricate relationships between monetary policy, inflation expectations, and exchange rate dynamics, while highlighting the importance of historical context in interpreting economic data and policy responses. The complex interplay between monetary and fiscal policies, coupled with shocks to the demand for safe dollar assets, explains a considerable portion of the variation in the US dollar exchange rate. This insight could prove instrumental in addressing the exchange rate disconnect puzzle and informing the optimal design of both fiscal and monetary policies.

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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 (2) have been suggested (e.g., Moneta et al., 2013; Matteson & Tsay, 2017; Lanne et al., 2017; Gouriéroux et al., 2017).[18] 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),

$$\mathcal{B} = \int_{(0,1)^K} \left[\sqrt{T} \left(C(\tilde{\epsilon}) - \prod_{k=1}^{K} U(\tilde{\epsilon}_k) \right) \right]^2 d\tilde{\epsilon}, \tag{4}$$

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

$$\widehat{D} = \widetilde{D}_{\hat{\theta}}, \text{with } \hat{\theta} = \text{argmin}_{\theta}\{\mathcal{B}|\tilde{\epsilon}_t = \widetilde{D}_{\theta}^{-1} u_t\}. \tag{5}$$

To implement (5), we use rotation matrices that structure the space of potential decompositions of the reduced form residual covariance estimates $\widehat{\Sigma}_u = G R_\theta R_\theta' G' = \widetilde{D}_\theta \widetilde{D}_\theta'$, where G is a lower triangular Cholesky factor of $\widehat{\Sigma}_u$ and $R_\theta R_\theta'$ is the identity matrix. Hence, $\widehat{D} = G R_{\hat{\theta}}$. Random vectors $\tilde{\epsilon}$ are determined from orthogonalized reduced form model disturbances ($\tilde{\epsilon}_t = \widetilde{D}_\theta^{-1} \hat{u}_t$), and

[18] 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 (5) can be achieved by means of nonlinear optimization.[19]

It is worth noting that the point estimate $\widehat{D}$ that solves (5) 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,** i_t, i_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*.

[19]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.

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 2022M6 for all eight economies. The time series used are 12-month US Treasury yields, bilateral spot exchange rates, and 10-year US inflation expectations.

- We restrict the end of the *pre-crisis* sub-sample for all economies to 2007M4. The *Volcker* (*post-crisis*) sub-sample is homogeneous for all economies from 1980M1 to 1987M7 (2008M4 to 2022M6).

- 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 from 1995M2 to 2022M6 and data from the codeset by Uribe (2022) on the CPC from 1981Q2 to 2018Q2.

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 IR	External	US IE	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 interest rate shock, external shock, and US inflation expectations 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 12.

Table 4: Testing fundamentalness of VAR residuals

p-max	1	2	3	4	5	6	7	8
AUS	0.62	0.59	0.55	0.52	0.50	0.48	0.46	0.45
CAN	0.82	0.79	0.77	0.75	0.73	0.72	0.71	0.69
CHE	0.38	0.35	0.32	0.29	0.27	0.26	0.25	0.25
DEU	0.71	0.68	0.65	0.63	0.61	0.6	0.59	0.59
GBR	0.34	0.32	0.30	0.27	0.26	0.25	0.24	0.24
JPN	0.45	0.41	0.37	0.35	0.33	0.31	0.31	0.30
NZL	0.60	0.60	0.59	0.59	0.58	0.58	0.58	0.58
SWE	0.84	0.83	0.82	0.81	0.81	0.80	0.79	0.79

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 12 using the Parzen Kernel. Results for alternative kernels and residuals from VAR(12) models are similar and 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:[20]

$$\widehat{D}_{AUS} = \begin{bmatrix} \underset{(0.325;5.204)}{0.331} & \underset{(0.005;0.046)}{-0.023} & \underset{(0.018;1.758)}{0.041} \\ \underset{(-0.243;-0.088)}{0.19} & \underset{(3.177;13.153)}{3.291} & \underset{(-0.072;-0.02)}{-0.158} \\ \underset{(-0.043;-1.4)}{-0.008} & \underset{(0.009;-0.479)}{-0.003} & \underset{(2.335;8.224)}{0.100} \end{bmatrix}, \widehat{D}_{CAN} = \begin{bmatrix} \underset{(0.329;5.187)}{0.338} & \underset{(-0.011;-0.847)}{-0.011} & \underset{(0.018;1.556)}{0.008} \\ \underset{(0.106;0.922)}{0.110} & \underset{(1.996;9.619)}{2.057} & \underset{(-0.136;-1.358)}{-0.167} \\ \underset{(-0.039;-1.035)}{0.001} & \underset{(0.02;0.354)}{0.003} & \underset{(2.329;8.273)}{0.1} \end{bmatrix}$$

$$\widehat{D}_{CHE} = \begin{bmatrix} \underset{(0.33;5.244)}{0.334} & \underset{(0.003;-0.024)}{-0.026} & \underset{(0.017;1.397)}{0.030} \\ \underset{(0.662;1.764)}{0.894} & \underset{(3.045;17.767)}{2.994} & \underset{(-0.185;-0.814)}{-0.126} \\ \underset{(-0.019;-0.6)}{-0.004} & \underset{(-0.021;-0.052)}{-0.002} & \underset{(2.34;8.27)}{0.099} \end{bmatrix}, \widehat{D}_{DEU} = \begin{bmatrix} \underset{(0.33;5.213)}{0.335} & \underset{(0.01;-0.214)}{-0.011} & \underset{(0.02;1.386)}{0.024} \\ \underset{(0.431;1.618)}{0.557} & \underset{(2.884;17.521)}{2.897} & \underset{(-0.105;-0.278)}{-0.078} \\ \underset{(-0.03;-0.575)}{-0.003} & \underset{(-0.005;0.025)}{-0.001} & \underset{(2.342;8.286)}{0.099} \end{bmatrix}$$

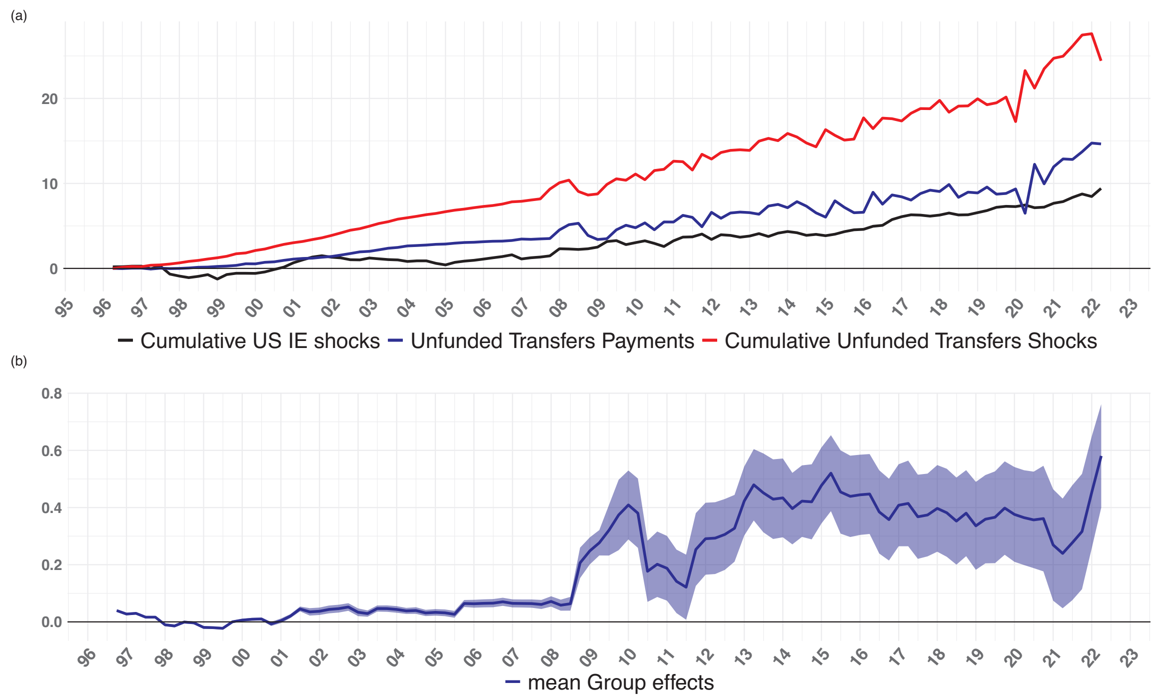
[20]For inferential purposes we use a Moving Block Bootstrap as suggested by Brüggemann et al. (2016). According to their recommendation the block length is set to 25 ($\approx 5.03\ T^{1/4}$). To improve the scaling of documented estimation results structural parameter estimates and bootstrap means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

$$\widehat{D}_{GBR} = \begin{bmatrix} \underset{(0.327;5.304)}{0.335} & \underset{(0.008;-0.518)}{-0.009} & \underset{(0.019;1.671)}{0.006} \\ \underset{(0.255;1.327)}{0.448} & \underset{(2.773;15.067)}{2.828} & \underset{(-0.114;-0.596)}{-0.17} \\ \underset{(-0.042;-1.115)}{0.002} & \underset{(-0.052;-1.088)}{-0.001} & \underset{(2.338;8.379)}{0.099} \end{bmatrix}, \widehat{D}_{JPN} = \begin{bmatrix} \underset{(0.328;5.281)}{0.336} & \underset{(0.028;-0.199)}{-0.010} & \underset{(0.023;1.854)}{0.023} \\ \underset{(0.239;1.607)}{0.553} & \underset{(3.022;16.842)}{2.970} & \underset{(-0.056;0.152)}{0.056} \\ \underset{(-0.055;-1.422)}{-0.002} & \underset{(-0.035;-0.513)}{-0.002} & \underset{(2.347;8.184)}{0.100} \end{bmatrix},$$

$$\widehat{D}_{NZL} = \begin{bmatrix} \underset{(0.322;4.883)}{0.332} & \underset{(0.018;0.453)}{0.024} & \underset{(0.015;1.741)}{-0.002} \\ \underset{(-0.392;-0.325)}{-0.430} & \underset{(3.347;10.91)}{3.491} & \underset{(-0.076;-0.017)}{0.017} \\ \underset{(-0.016;-1.159)}{-0.003} & \underset{(0.022;-0.178)}{-0.001} & \underset{(2.328;8.371)}{0.099} \end{bmatrix}, \widehat{D}_{SWE} = \begin{bmatrix} \underset{(0.329;5.376)}{0.332} & \underset{(0.018;-0.516)}{-0.021} & \underset{(0.022;1.673)}{0.040} \\ \underset{(0.122;1.479)}{0.513} & \underset{(3.113;17.186)}{3.165} & \underset{(-0.069;-0.33)}{-0.003} \\ \underset{(-0.037;-1.106)}{-0.007} & \underset{(-0.035;-0.111)}{-0.001} & \underset{(2.346;8.289)}{0.099} \end{bmatrix}$$

E - Estimation results with unfunded fiscal transfer shocks

Figure 6: Historical effects of US inflation expectations shocks on unfunded transfer shocks



Notes: Panel (a) displays cumulative US IE shocks as identified in this work joint with cumulative unfunded fiscal transfer shocks, and unfunded transfer payments (in levels and billions of US dollars) identified by Bianchi et al. (2023a) from 1996Q2 to 2022Q2. Panel (b) shows estimated slope coefficients for rolling window regressions

$$UF_t = \alpha + \beta\epsilon_{3,s} + u_s,\ s = \tau_1, \tau_1 + 1, \ldots, \tau_2,$$

where $\epsilon_{3,s}$ is the US IE shock, aggregated to quarterly data, UF_t is the sum of unfunded transfer shocks in $s-1$, s, and $s+1$, and τ_1 (τ_2) is the lower (upper) bound of rolling samples of size 60 quarters. Owing to data availability, the regression starts in 1981Q2 and ends in 2022Q2. The solid line shows the arithmetic mean of point estimates across the eight advanced economies, while the shaded areas represent approximate pointwise confidence intervals with 95% coverage.

F - Forecast-error variance decompositions

Figure 7: Sub-samples mean group FEVDs for exchange rate changes, 1980M1-2022M6.