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NOVEMBER 2021
The analyses, opinions and findings of these papers represent the views of the authors, they are not necessarily those of the Banco de Portugal or the Eurosystem

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# Permanent and temporary monetary policy shocks and the dynamics of exchange rates

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**Abstract**
Over the short run contractionary monetary policy shocks tend to be associated with domestic currency appreciations, which goes against standard interest rate parity conditions. How can this be reconciled with the fact that these conditions tend to be restored over the long run? We show the distinction between permanent and temporary monetary policy shocks is helpful to understand the impacts of monetary policy on exchange rates in the short as well as over the long run. Drawing on monthly data for the United States, Germany, France, Great Britain, Japan, Australia, Switzerland and the euro area from 1971 to 2019, and resorting to a simple structural vector error correction (SVEC) model and mild identifying restrictions, we find that a shock leading to a temporary increase in U.S. nominal interest rates leads to a temporary appreciation of the USD against the other currencies, in line with the literature on the exchange rate effects of monetary shocks and that on the forward premium puzzle. In turn, a monetary policy shock leading to a permanent rise in nominal interest rates - e.g. one associated with a normalisation of monetary policy after a long period at the zero lower bound - has the opposite impact, i.e., in line with interest parity conditions, in the short as well as over the long run. The ensuing depreciation may also contribute to higher (not lower) inflation, also in the short run. We thus confirm, in a simpler setting and for more economies, the results of Schmitt-Grohé and Uribe (2021). This highlights the relevance of differentiating between temporary and permanent monetary policy shocks in interpreting short-run exchange rate movements.

JEL: E52, E58, F31, C32

Keywords: Exchange Rates, Fisher Relation, Monetary Policy Cointegration, Monetary Shocks, Structural VEC Models.

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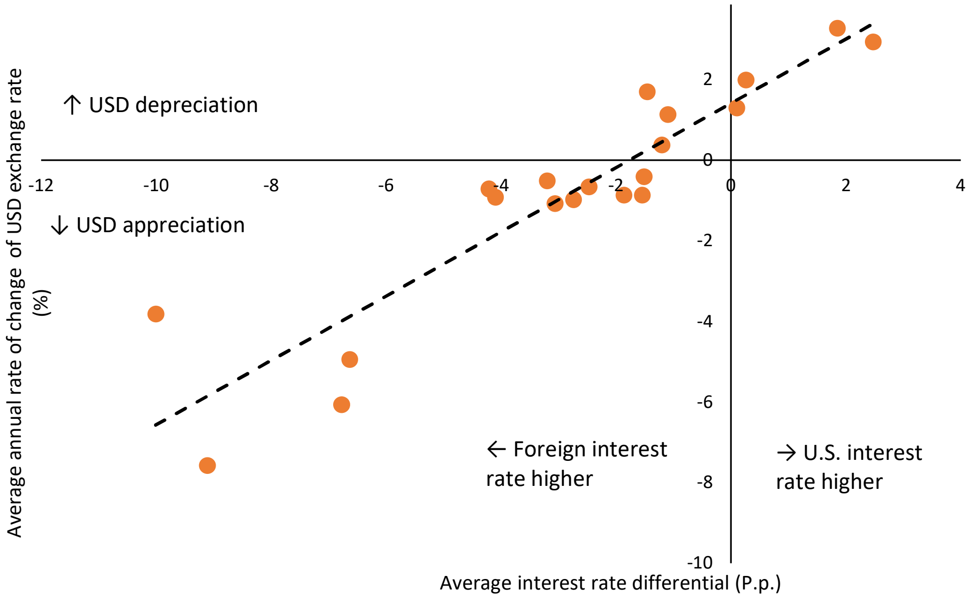
## 1. Introduction

If the Federal Reserve (FED) unexpectedly raises policy interest rates, the U.S. dollar (USD) should appreciate against major currencies, as investors rush to USD denominated higher yielding assets. These are the effect and explanation most often stressed by market participants and policy-makers. At least over the short run, this effect finds support in the empirical literature documenting the impact of identified monetary policy shocks on exchange rates (see Eichenbaum and Evans (1995) for an early reference). Other identification approaches that rely on sign restrictions (see Kim and Roubini (2000) and Scholl and Uhlig (2008)), on more structural frameworks (see Bjornland (2008)) and on changes in short-term interest rates that are exogenous to other economic news (see Zettelmeyer (2004)) also find these results. A similar effect is found if instead of focusing on monetary policy shocks one looks at differences between policy interest rates of major central banks: at least over the short run, a positive difference between domestic and foreign interest rates tends to be associated with an appreciation of the domestic currency. The seminal work of Fama (1984) is one of the earliest references for this so-called forward discount/premium bias. This premium arises when the forward exchange rate, which typically indicates a depreciation of the currency with the higher interest rate - that would ensure an equalisation of nominal returns -, is not an accurate predictor of the future spot exchange rate.[1] In fact, most evidence suggests that, over short horizons, the future spot rate moves opposite to the direction of the forward rate some time before. Several authors (see, e.g., Froot and Frankel (1989) for an earlier reference) try to explain these facts on account of risk premia or even a failure of rational expectations.

Standard interest rate parity conditions would dictate an opposite movement of the exchange rate, particularly over the long run. Take the uncovered interest rate parity (UIRP): under free capital mobility and residual transaction costs, the nominal returns on risk-free assets in different currencies would tend to be equalised when expressed in one of the currencies. An increase in the interest rate of an economy should thus be compensated by a depreciation of its currency to ensure an equalisation of returns. Over long periods, UIRP seems indeed to be recovered, at least approximately. A simple graphical representation of the average monthly interest rate differentials (*vis-à-vis* the U.S. and considering short rates) and the average monthly annual growth rate of the exchange rate of the USD, for a pool of 20 economies with flexible exchange rate regimes, points to a positive relation very much in line with UIRP, as depicted in Figure 1.[2] This evidence hints that, over a large time span, a higher level of the domestic interest rate tends to be associated with domestic currency depreciations, as Lothian (2016) more carefully

1. See Miller (2014) for a detailed review of this topic.

2. In the paper exchange rates are presented in direct terms regarding the USD, that is, one unit of the foreign currency corresponds to $x$ USD. Accordingly, an increase (decrease) in the exchange rate means a depreciation (appreciation) of the USD against the foreign currency.



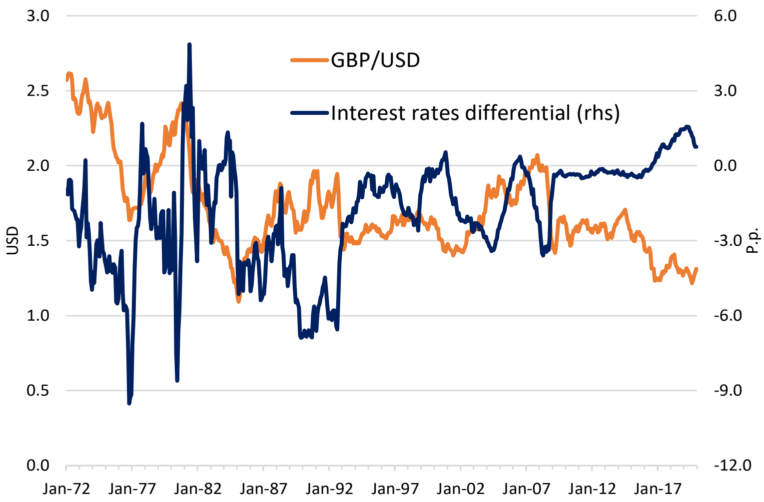

The figure displays for 20 economies the average monthly interest rate differential with the U.S. interest rate (3-month interest rates, annualised) and the monthly average growth rate of the exchange rate of the respective currency against the USD (annualised). The sample period spans from 1971, or whenever data is available, to 2019. Dotted line represents a linear trend.

Figure 1: Average interest rate differential *vs.* average exchange rate against the USD

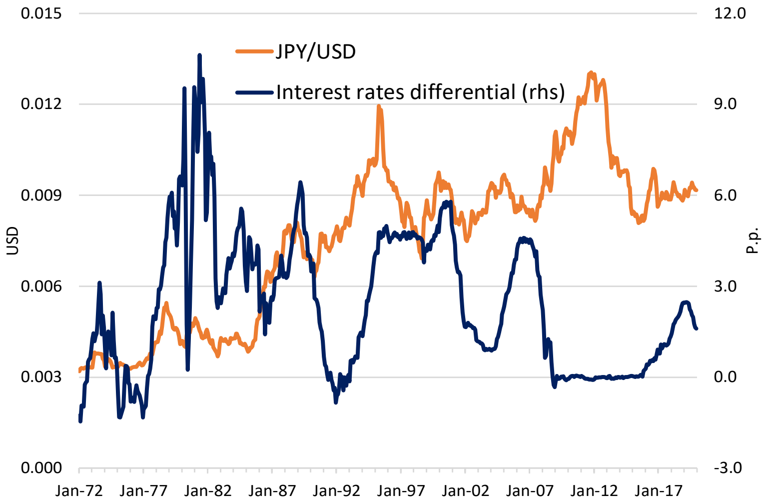
documents. It is worth highlighting that in this figure the regression line does not pass through the origin, indicating a premium associated with USD holdings: after accounting for exchange rate movements, on average, investors seem to tolerate a lower return on USD denominated assets. The focus of this paper will be in the relation between *changes* in monetary policy and *changes* in exchange rates, with no emphasis on these premia.

This positive relation is not always apparent in the short run. Figure 2 plots exchange rates and interest rate differentials, focusing on the pairs U.S. (USD)/Great Britain (GBP) and U.S. (USD)/Japan (JPY). While it is very clear that persistently positive (negative) interest rate differentials are associated with a systematic depreciation (appreciation) of the USD, there are certainly periods in which a rising differential is associated with an appreciation of the USD (in line with the forward discount/premium bias).

How can the short-run and long-run evidence be reconciled? An appealing way is to follow Schmitt-Grohé and Uribe (2021) and assume that monetary shocks come in two types: temporary and permanent shocks. This is warranted if long-run variability, likewise a unit-root behaviour, is found in the data. Then, if temporary shocks have the usual short-term effect and permanent shocks have the usual long-run effects, the two types of evidence are potentially reconcilable. But one




(a) Panel A - G.B./U.S.




(b) Panel B - JP/U.S.

This figure displays, for the period between 1972 and 2019, the evolution of the GBP/USD and the JPY/USD exchange rates in comparison with the respective interest rate differentials computed as the spread between the nominal 3-month U.S. interest rate and the 3-month interest rate of the other economy.

Figure 2: Exchange rate and interest rate differentials

important question remains, in particular for policy-makers: what is the short-run impact of these permanent shocks? Notice that the usual effect of a contractionary monetary policy shock on the exchange rate also affects U.S. inflation through lower import prices, which could be useful if the FED were seeking lower inflation, but not useful at all if deflationary trends persist. However, if such shock is a permanent one - one leading to higher interest rates - and its effects over the short run are similar to the long-run effects, then the ensuing USD depreciation ought to come with higher inflation. Similarly, for investors betting on the intentions of the FED and very sensibly relying on the forward premium puzzle, it may be useful to recognise that the usual overall effect of changes in policy rates on the exchange rate is a combination of the effects of temporary and permanent shocks that may have different impacts over the short run. Hence, the usual effects may be contaminated by the presence of a shock - the permanent one -, with a distinct nature and effects that are the opposite of the standard effects.

This paper seeks to understand and reconcile in the simplest possible setting the short and long-run relation between exchange rates and nominal interest rates. Similarly to Schmitt-Grohé and Uribe (2021), we allow for different responses of these variables to temporary and permanent monetary policy shocks. Their work presents the first evidence that temporary and permanent monetary policy shocks may have different impacts on exchange rates, focusing on the U.S., Great Britain, Japan and Canada.[3] While it is found that the temporary monetary policy shock has the traditional short-run impact (that is, an increase in the policy interest rate leads to an appreciation of the domestic currency), a permanent contractionary monetary policy shock actually results in a depreciation of the domestic currency. This helps reconciling the apparent short-run *vs.* long-run inconsistency found in the literature. The goal of this work is to verify whether the results of Schmitt-Grohé and Uribe (2021) hold by considering a broader set of economies, while employing less stringent identifying restrictions and resorting to a simpler structural parametrisation, a standard SVEC model. Our empirical exercise relies on monthly data for exchange rates, inflation, nominal interest rates and output from 1971 to 2019, considering the U.S. and another advanced economy among the following: Great Britain, Germany, France, Australia, Switzerland, Japan and the euro area. Given the properties of the data, we impose cointegration between nominal interest rates in the two economies and between the nominal interest rate and the inflation rate in the U.S. economy, but unlike Schmitt-Grohé and Uribe (2021) we do not require the coefficients to be one. We can thus account for the observed slow fall in real interest rates over the past decades. Also, unlike Schmitt-Grohé and Uribe (2021), we do not impose that the effects of monetary policy shocks on inflation and output are the usual ones. We are agnostic on these effects, just as in Valle e Azevedo *et al.* (2019). Our identification is facilitated by these cointegrations, which most often find clear support in the data. On top of this, we impose standard long-run monetary neutrality restrictions: the permanent monetary policy shock does not affect permanently output and the permanent output shock does not affect

3. De Michelis and Iacoviello (2016) evaluate the impact of an inflation target shock on the exchange rates for the Japanese and U.S. economies and find some evidence of domestic currency depreciation after an increase in the inflation target, but no distinction is made between temporary and permanent monetary policy shocks.

inflation and nominal rates (one of these latter restrictions can even be relaxed with little change in the results). Finally, in order to distinguish the temporary shocks we further assume that the exchange rate shock has no contemporary effect on nominal interest rates and that the temporary monetary policy shock in the second economy has no contemporaneous effect on U.S. inflation.

Our setup is flexible enough to accommodate deviations from this standard configuration, depending on the specificities of the data. Take a model with the U.S. and Japan: over the whole sample, cointegration between nominal rates is not an adequate hypothesis, even though long-run average differences in those rates are mirrored in systematic exchange rate variations. Also, if the focus is instead on a post 1995 sample, it is more reasonable to assume that interest rates in Japan are stationary, while in the U.S. they are not, potentially implying a drift in the exchange rate variation. In the case of the euro area, given its similar size relative to the U.S. economy, we consider an extended and symmetric version of the model, i.e., each economy is treated in the same way for identification purposes.

Our results show, with remarkable consistency across countries and setups, that permanent and temporary U.S. monetary policy shocks have opposite impacts on the USD exchange rate against the currencies under analysis. A temporary U.S. contractionary monetary policy shock leads to a temporary appreciation of the USD, while the permanent shock leads to a depreciation, even in the short run. In this sense, the forward bias puzzle is not present in the case of a permanent shock, which highlights again the importance of distinguishing between temporary and permanent monetary policy shocks. We also report that the permanent U.S. monetary policy shock does not account for a large share of the forecast error variance decomposition (FEVD) of the USD exchange rate, contrary to Schmitt-Grohé and Uribe (2021), who find that permanent monetary policy shocks explain the majority of the FEVD of the exchange rate. This divergence can result from the assumption of non-stationary of the rate of change of the exchange rate used by Schmitt-Grohé and Uribe (2021) (for which we do not find statistical support) coupled with their imposed cointegration between this variable and the permanent monetary policy shocks.

Overall, our results still hold even if the sample period is restricted to the pre-inflation targeting period. Finally, the “neo-Fisher” effect firstly uncovered by Uribe (2021) and corroborated by Valle e Azevedo *et al.* (2019), i.e., the fact that a permanent increase in policy rates may have a positive impact on inflation even in the short run, survives to this opening of the U.S. economy.

The rest of the paper is organised as follows: Section 2 outlines the methodology and the data used in the empirical analysis. Section 3 presents the main results. Section 4 discusses the model specified for Japanese data, while Section 5 expands the original model for the analysis with the euro area. Section 6 concludes. A Supplementary Material File provides more detailed results for all countries and several robustness analyses.

## 2. Methodology

To measure the impact of temporary and permanent monetary policy shocks on the dynamics of exchange rates we follow closely the SVEC framework and identification methodology exposed in Lütkepohl (2006). We consider the U.S. economy together with Great Britain (G.B.), Germany (DE), France (FR), Australia (AU), Switzerland (CH), Japan (JP) and the euro area (EA).

The data comprises monthly time series from 1971, or whenever they become available, through to 2019. In the main model, the empirical analysis relies on five variables:

- 3-month annualised rate of change of the exchange rate between the USD and the currency of the respective advanced economy - $e_t$;
- 3-month interest rates extracted from Treasury bills or money market instruments, depending on data availability, for both the U.S. and the other advanced economy, respectively $i_t$ and $i_t^*$;
- U.S. core inflation, measured by the year-on-year rate of change of the CPI excluding food and energy - $\pi_t$;
- U.S. industrial production index, as a proxy for output at a monthly frequency - $y_t$.[4]

In the case of the euro area, a more symmetric setting will be employed such that euro area core inflation and industrial production are also included. A complete description of the data can be found in the Appendix. In our sample all variables are found to be non-stationary, with the exception of $e_t$, which is found to be I(0).[5] Next, Table 1 reveals the results of a Johansen trace test for cointegration between the nominal interest rate of the U.S. and that of the other economies. The tests clearly point to cointegration between nominal rates across the economies in our sample, with the exception of Japan.[6] Given this evidence, we henceforth assume that nominal interest rates are cointegrated (except for Japan), though not necessarily with a coefficient equal to unity. This is a departure from the main specification of Schmitt-Grohé and Uribe (2021), where it is assumed that $e_t$ is non-stationary. Finally, as in Valle e Azevedo *et al.* (2019), we assume cointegration between nominal interest rates and inflation rates for the U.S. and the euro area,

4. Industrial production growth correlates well with GDP growth even in the context of a decreasing contribution of manufacturing to output. It has the advantage of being a monthly series available for large time spans.

5. Based on standard ADF tests, there is evidence that both inflation and output are I(1) in all countries. Nominal interest rates also emerge as non-stationary considering the full sample, but not as evidently in the cases of Germany and Switzerland. A remarking conclusion is that the rate of change of the exchange rate of the USD against the other currencies is undoubtedly stationary. In view of including this variable in the SVEC model, it is assumed that it does not have any long-run relation with any other variable in the model.

6. For further evidence of cointegration in monetary policies, see Belke and Cui (2010) and Arouri *et al.* (2013).

although the coefficients are not restricted to one, again unlike Schmitt-Grohé and Uribe (2021).

| Johansen Trace Tests | U.S./G.B. 3 lags | U.S./DE 2 lags | U.S./FR 3 lags | U.S./AU 3 lags | U.S./CH 7 lags | U.S./JP 9 lags |
|---|---|---|---|---|---|---|
| 0 | 0.02 | 0.02 | 0.00 | 0.01 | 0.00 | 0.46 |
| 1 | 0.80 | 0.52 | 0.49 | 0.58 | 0.72 | 0.73 |

Notes: The Table displays p-values for the null of "at most 0 (or 1) cointegration relations". Specification with constant and trend. AIC for lag choice.

Table 1. Johansen trace tests for monetary policy cointegration

The basic reduced form of the VEC model for each country is given by:

$$\Delta X_t = \alpha_0 + \gamma\beta X_{t-1} + \sum_{j=1}^{K} \beta_j \Delta X_{t-j} + u_t \tag{1}$$

where $X_t := (e_t, \pi_t, i_t^*, i_t, y_t)'$. $\gamma$ and $\beta$ correspond to the loading coefficients and cointegration coefficients matrices, respectively, and $u_t := (u_t^e, u_t^\pi, u_t^{i^*}, u_t^i, u_t^y)'$ is a vector of reduced form serially uncorrelated shocks. Equivalently, focusing on the cointegration part (second and third terms on the right-hand side):

$$\begin{cases} \Delta e_t = \gamma_{11} e_{t-1} + \gamma_{12}(\pi_{t-1} - \beta^F i_{t-1}) + \gamma_{13}(i_{t-1}^* - \beta^{MP} i_{t-1}) + ... + u_t^e \\ \Delta \pi_t = \gamma_{21} e_{t-1} + \gamma_{22}(\pi_{t-1} - \beta^F i_{t-1}) + \gamma_{23}(i_{t-1}^* - \beta^{MP} i_{t-1}) + ... + u_t^\pi \\ \Delta i_t^* = \gamma_{31} e_{t-1} + \gamma_{32}(\pi_{t-1} - \beta^F i_{t-1}) + \gamma_{33}(i_{t-1}^* - \beta^{MP} i_{t-1}) + ... + u_t^{i^*} \\ \Delta i_t = \gamma_{41} e_{t-1} + \gamma_{42}(\pi_{t-1} - \beta^F i_{t-1}) + \gamma_{43}(i_{t-1}^* - \beta^{MP} i_{t-1}) + ... + u_t^i \\ \Delta y_t = \gamma_{51} e_{t-1} + \gamma_{52}(\pi_{t-1} - \beta^F i_{t-1}) + \gamma_{53}(i_{t-1}^* - \beta^{MP} i_{t-1}) + ... + u_t^y \end{cases} \tag{2}$$

where the superscript "$F$" in $\beta^F$ refers to the Fisher relation, whereas "$MP$" refers to monetary policy cointegration. The treatment of $e_t$ follows from its stationarity. To recover the structural form shocks it is necessary to impose some identifying assumptions on a non-singular matrix $B$, such that $u_t = B\varepsilon_t$, where $\varepsilon_t := (\varepsilon_t^e, \varepsilon_t^\pi, \varepsilon_t^{i^*}, \varepsilon_t^i, \varepsilon_t^y)'$ represents five serially and mutually uncorrelated structural shocks. Following the steps in Lütkepohl (2006) (Chapter 9), the VEC model in its reduced form implies the following Beveridge-Nelson decomposition:

$$X_t = X_{0,t}^* + \Xi \sum_{i=1}^{t} u_i + \sum_{j=0}^{\infty} \Xi_j^* u_{t-j} \tag{3}$$

where the first term collects initial values and deterministic trends, the second term accounts for the long-run effects of the shocks and the last term is absolutely summable and thus stationary. Given the relation between the reduced form and the structural form shocks, this decomposition can be expressed as:

$$X_t = X_{0,t}^* + \Xi B \sum_{i=1}^{t} \varepsilon_i + \sum_{j=0}^{\infty} \Xi_j^* B \varepsilon_{t-j} \tag{4}$$

Lütkepohl (2006) (Chapther 6) shows that matrix $\Xi$, a function of the reduced form parameters, has reduced rank in the presence of cointegration. In our baseline model with five variables and three cointegration relations (the stationarity of $e_t$ is conveniently treated as one, together with cointegration between $i_t$ and $i_t^*$ and cointegration between $i_t$ and $\pi_t$), the rank of matrix $\Xi$ is two. Given that matrix $B$ is non-singular, the rank of matrix $\Xi B$ is also two. This is consistent with stating that our model has two stochastic trends driving the data: $(i)$ the trend that gears inflation and the short-term interest rate in the U.S. as well as the short-term interest rate in the other advanced economy and $(ii)$ the trend that drives output in the U.S. economy. The rate of change of the exchange rate has no stochastic trend given its stationary behaviour. All this helps in the identification of $B$, since we can make assumptions on the long-run impact matrix $\Xi B$ to make sure it has rank two. We consider that the structural shock to the inflation rate of the U.S. is a permanent one, and also the one driving the two short-term interest rates in the model. The structural shocks to the interest rates thus only have temporary effects.[7] This assumption, together with the stationarity of $e_t$, implies that the first, third and fourth columns of $\Xi B$ are zeroes. The first row of $\Xi B$ is also zero since the rate of change of the exchange rates cannot be affected by any shock in the model in the long run. Two other assumptions are made in $\Xi B$: $(i)$ the first simply assumes that the permanent monetary policy shock has no long-run impact on output (entry (5,2)), a standard neutrality proposition and $(ii)$ the permanent output shock has no long-run impact on the level of nominal rates or inflation (entries (2,5), (3,5) and (4,5)).[8] Given all this, matrix $\Xi B$ has the following structure:

$$\Xi B = \begin{bmatrix} 0 & 0 & 0 & 0 & 0 \\ 0 & * & 0 & 0 & 0 \\ 0 & * & 0 & 0 & 0 \\ 0 & * & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & * \end{bmatrix} \tag{5}$$

This setting is helpful, but it is not enough to fully identify $B$, since the three temporary shocks must be distinguished and this can only be done directly on $B$. We assume that the structural shock on the rate of change of the exchange rate has no contemporaneous impact on the two nominal interest rates and that the temporary monetary policy shock of the other advanced economy has no contemporaneous impact on the inflation rate of the U.S. economy. This implies a $B$ matrix of the

7. This permanent shock only has permanent effects on both inflation and the two nominal rates. This choice is innocuous and we could have picked either the U.S. or foreign nominal interest shock as the permanent shock and the inflation shock as the other transitory shock; in this way, we would obtain exactly the same impulse response functions as those obtained with the alternative identification, only the labelling of shocks would be switched. The meaningful assumption is that there is only one permanent shock, instead of two or three such shocks with "collinear" effects on these three variables.

8. This actually results in an overidentification of the permanent shocks, but overall the results are similar if we assume there is no long-run effect on nominal rates (only) or on inflation (only).

following form:

$$B = \begin{bmatrix} * & * & * & * & * \\ * & * & 0 & * & * \\ 0 & * & * & * & * \\ 0 & * & * & * & * \\ * & * & * & * & * \end{bmatrix} \tag{6}$$

Combining these short-run and long-run identification restrictions we obtain sufficient conditions for the estimation of the structural model.

## 3. Empirical results

| | Fisher relation U.S. | | | | | |
|---|---|---|---|---|---|---|
| | $\beta^F$ | $\gamma_e$ | $\gamma_{\pi^{US}}$ | $\gamma_{i^*}$ | $\gamma_{i^{US}}$ | $\gamma_{y^{US}}$ |
| U.S. - G.B. | $-0.59^{***}$ | $-0.21$ | $-0.03^{***}$ | $-0.01$ | $0.01$ | $-0.01$ |
| (6 lags) | 0.09 | 0.30 | 0.01 | 0.02 | 0.01 | 0.01 |
| U.S. - DE | $-0.62^{***}$ | $-0.44$ | $-0.02^{***}$ | $-0.01$ | $0.02^{*}$ | $-0.04^{***}$ |
| (2 lags) | 0.08 | 0.36 | 0.01 | 0.01 | 0.01 | 0.01 |
| U.S. - FR | $-0.59^{***}$ | $-0.48$ | $-0.02^{***}$ | $-0.02$ | $0.03^{**}$ | $-0.02$ |
| (4 lags) | 0.09 | 0.32 | 0.01 | 0.01 | 0.01 | 0.01 |
| U.S. - AU | $-0.56^{***}$ | $-0.16$ | $-0.02^{***}$ | $-0.02$ | $0.02^{*}$ | $-0.04^{***}$ |
| (2 lags) | 0.09 | 0.32 | 0.01 | 0.02 | 0.01 | 0.01 |
| U.S. - CH | $-0.60^{***}$ | $-0.12$ | $-0.02^{***}$ | $-0.02$ | $0.02^{*}$ | $-0.05^{***}$ |
| (2 lags) | 0.07 | 0.44 | 0.01 | 0.01 | 0.01 | 0.01 |

Notes: $\gamma_{var}$ corresponds to the adjustment parameter in the equation for variable "$var$". Superscript "US" was added to variables $\pi$, $i$ and $y$ variables to make clear they are U.S. variables. Standard errors displayed below the coefficient estimates. Simple two step estimator (S2S) employed as implemented in JMulti. $***$, $**$ and $*$ denote statistical significance at 1%, 5% and 10% levels, respectively. U.S. - G.B. model: 1971M4-2017M6. U.S. - DE model: 1971M1-2019M6. U.S. - FR model: 1971M1-2019M5. U.S. - AU model: 1971M4-2019M6. U.S. - CH model: 1974M1-2019M6. Hannan-Quinn criterion for lag choice.

Table 2. Estimation of Fisher Relation Parameters

Tables 2 and 3 present the parameter estimates of the VEC model for the cointegration vector parameters ($\beta$) and the adjustment parameters ($\gamma$) for the models with G.B., DE, FR, AU and CH and the two cointegration relations assumed.[9] The analyses for Japan and the euro area are done separately on account of the different specifications required, as discussed in Section 1. Regarding the cointegration between inflation and nominal interest rates in the U.S. (Fisher relation for the U.S. economy), our results point to cointegrating parameters between $0.56$ and $0.62$ in all the models, consistent with Valle e Azevedo *et al.* (2019). Likewise, in the majority of the models, both inflation and nominal interest

9. All the results in this paper regarding estimation of the structural VECM are obtained using JMulti, see http://jmulti.de.

| | Monetary policy Cointegration | | | | | |
|---|---|---|---|---|---|---|
| | $\beta^{MP}$ | $\gamma_e$ | $\gamma_{\pi^{US}}$ | $\gamma_{i^*}$ | $\gamma_{i^{US}}$ | $\gamma_{y^{US}}$ |
| U.S. - G.B. | $-1.18^{***}$ | $0.57^{**}$ | $0.00$ | $-0.07^{***}$ | $0.00$ | $-0.02$ |
| (6 lags) | 0.08 | 0.26 | 0.01 | 0.01 | 0.01 | 0.01 |
| U.S. - DE | $-0.70^{***}$ | $0.09$ | $0.01$ | $-0.02^{***}$ | $-0.01$ | $-0.04^{***}$ |
| (2 lags) | 0.12 | 0.29 | 0.01 | 0.01 | 0.01 | 0.01 |
| U.S. - FR | $-1.19^{***}$ | $-0.08$ | $0.00$ | $-0.06^{***}$ | $-0.02^{*}$ | $-0.03^{***}$ |
| (4 lags) | 0.08 | 0.29 | 0.01 | 0.01 | 0.01 | 0.01 |
| U.S. - AU | $-1.24^{***}$ | $-0.02$ | $0.00$ | $-0.05^{***}$ | $0.00$ | $-0.02^{**}$ |
| (2 lags) | 0.16 | 0.18 | 0.00 | 0.01 | 0.01 | 0.01 |
| U.S. - CH | $-0.58^{***}$ | $0.45$ | $0.00$ | $-0.04^{***}$ | $-0.02^{**}$ | $-0.03^{***}$ |
| (2 lags) | 0.11 | 0.32 | 0.00 | 0.01 | 0.01 | 0.01 |

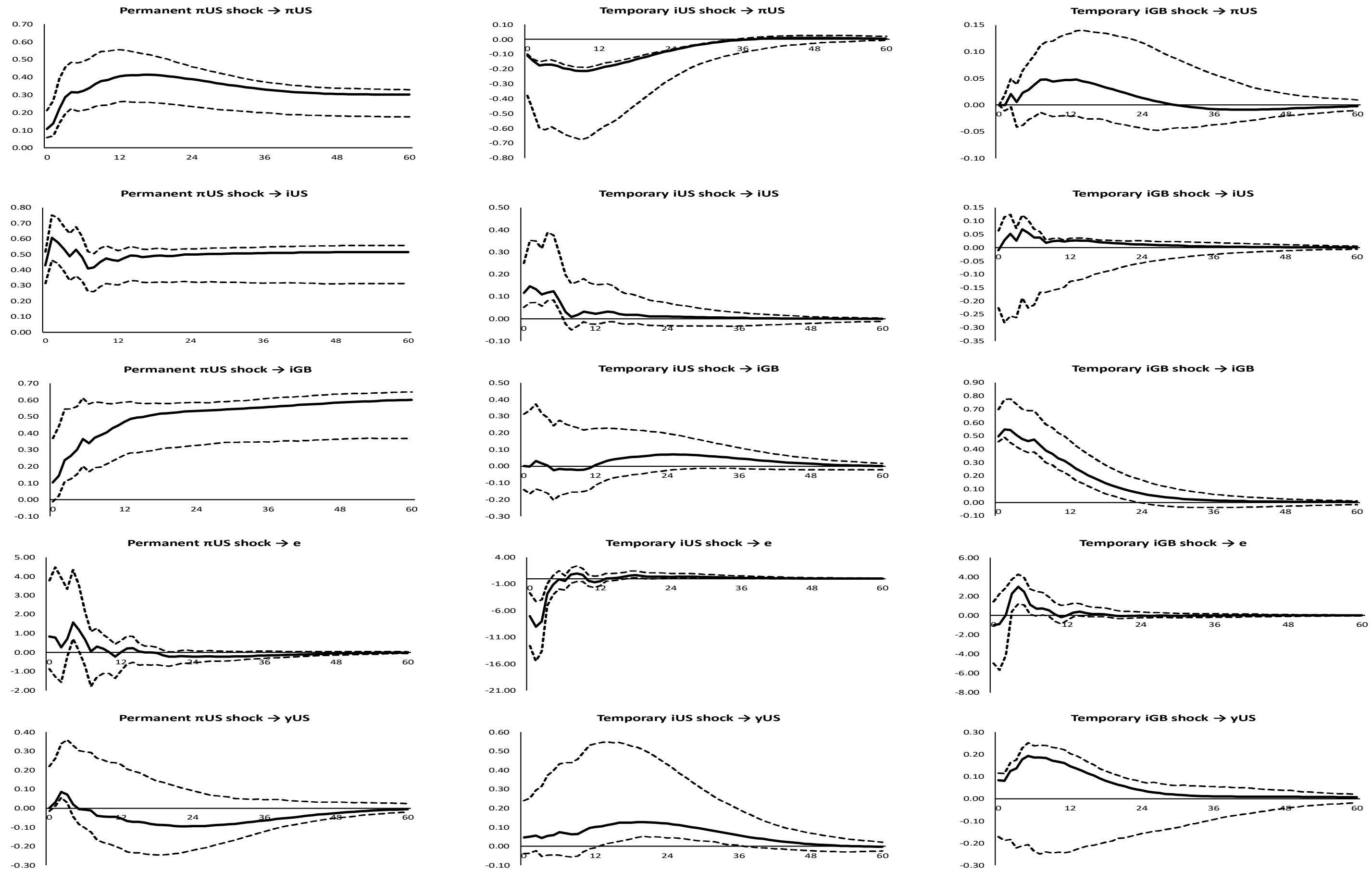
Notes: $\gamma_{var}$ corresponds to the adjustment parameter in the equation for variable "$var$". Superscript "US" was added to variables $\pi$, $i$ and $y$ variables to make clear they are U.S. variables. Standard errors displayed below the coefficient estimates. Simple two step estimator (S2S) employed as implemented in JMulti. $***$, $**$ and $*$ denote statistical significance at 1%, 5% and 10% levels, respectively. U.S. - G.B. model: 1971M4-2017M6. U.S. - DE model: 1971M1-2019M6. U.S. - FR model: 1971M1-2019M5. U.S. - AU model: 1971M4-2019M6. U.S. - CH model: 1974M1-2019M6. Hannan-Quinn criterion for lag choice.

Table 3. Estimation of Monetary Policy Cointegration Parameters

rates in the U.S. adjust to return to the long-run relation, which can be concluded from the significance of $\gamma_{\pi^{US}}$ and $\gamma_{i^{US}}$ in Table 2. The only exception is the model for G.B., where only inflation displays a significant adjustment parameter. As for the cointegration between monetary policies, the parameters hover around one, but range from $0.58$ to $1.24$. Also, the nominal interest rate of the second economy reacts significantly to re-establish the long-run relation, as seen by the significance of $\gamma_{i^*}$ in Table 3. U.S. interest rates also react to re-establish the long-run relation, albeit less pronouncedly, in the models with CH and FR. It is important to recall that in our sample the nominal interest rates of CH are the closest to stationarity, so these results ought to be interpreted with caution. This adjustment behaviour in the cointegration of monetary policies overall suggests a leading behaviour of the U.S. economy *vis-à-vis* the other economies in our sample.[10]

As regards the impulse response functions, we focus here on the results for the G.B. and present figures and more detailed comments for DE, FR, AU and CH in a Supplementary Material File. Figure 4 displays for the G.B. model the responses of U.S. and G.B. nominal interest rates, U.S. inflation, the rate of change of the exchange rate and U.S. output to the identified structural monetary shocks. The first column shows the response to the permanent U.S. monetary policy shock, the second column focuses on the U.S. temporary monetary policy shock and the third column reports the response to the G.B. temporary monetary policy shock.

10. Gray (2013) presents empirical and theoretical evidence of this behaviour between the U.S. and 12 countries from 1980 to 2008 on a panel regression setting and Belke and Gros (2005) provide some evidence specifically between the Federal Reserve and the European Central Bank.




All responses are in percentage points. The first column reports the responses to the permanent U.S. monetary shock, the second column the response to the temporary monetary shock and the third column the response to G.B. temporary monetary shock. Shocks are identified as described in the text. The structural model is estimated by Maximum Likelihood, using the Scoring Algorithm of Amisano and Giannini (1997) as implemented in JMulTi. Response to a one standard deviation shock. 90 % Hall Bootstrap confidence intervals.

Figure 4: Impulse Response Functions - U.S. - G.B. model

The figure shows that the impacts of the permanent and temporary U.S. monetary policy shocks on inflation and nominal rates are, qualitatively, exactly those documented in Valle e Azevedo *et al.* (2019) and Uribe (2021): the permanent shock, associated with a permanent rise in U.S. nominal rates, leads to a permanent increase in inflation, even in the short run, while the temporary shock, associated with a temporary increase in interest rates, has a negative impact on inflation. All adjustments essentially take place within 3 years. The permanent U.S. monetary policy shock also has a permanent effect on the nominal interest rate of G.B., as expected from the cointegration relation between nominal interest rates. The temporary U.S. monetary policy shock appears to have an insignificant effect on the nominal interest rates of the G.B. economy. However, for the other countries considered, this impact is always positive and significant.

Focusing on the impacts of G.B. temporary monetary policy shock, which is associated with an increase in the nominal interest rate of that economy, it is found that most impacts on U.S. variables are not statistically significant at conventional levels (the same is true in the models with other countries, with the exception of the model for France). Looking at the impacts of the structural shocks on U.S. output, the permanent U.S. monetary policy shock has a positive impact in the models that include G.B. and Australia (in the models that include other countries the impact is most often estimated to be positive but not significant, see the Supplementary Material File) while the U.S. temporary monetary policy shock also seems to have a positive impact, but only in the model for G.B. (in the models that include other countries the impact is most often not significant). This contrasts with the negative effect robustly found in Valle e Azevedo *et al.* (2019). Recall that a negative impact is assumed *a priori* in Uribe (2021).

While it is reassuring that the abovementioned effects of permanent and temporary nominal shocks on inflation survive to this opening of the U.S. economy (less so for output), the focus of this paper lies on the fourth row in Figure 4, displaying the impacts of the structural shocks on $e_t$. One easily concludes that the temporary monetary policy shocks (both in the U.S. and in the second economy) lead to an appreciation of the domestic currency in most models. Here, the only exception appears in the model with Australia, where the temporary monetary policy shock of this economy has no significant impact on the rate of change of the exchange rate. When we move to the permanent U.S. monetary policy shock, the impact is the opposite, i.e., a permanent increase in the U.S. nominal interest rate leads to a depreciation of the USD against the five currencies considered in our sample. This result confirms the findings of Schmitt-Grohé and Uribe (2021) who document that a permanent monetary policy shock leads to a depreciation of the domestic currency using data for the U.S., G.B., Japan and Canada. Comparing the impacts of the temporary and permanent U.S. monetary policy shocks, it is notorious that the temporary shock has a much stronger and immediate impact on the exchange rate than the permanent one. In a sense, this outcome could be expected given the strong consensus in the literature pointing to currency appreciations when a distinction between temporary and permanent monetary policy shocks is not made. At the same time, the impact of the permanent U.S. monetary policy shock appears to be immediate but short-lived (less than one year), with the exception being in the model with Australia. Across models, the USD depreciation followed by the permanent U.S. monetary policy shock peaks after two to four months, despite the increase in nominal interest rates being immediate. Recall that a permanent U.S. monetary policy shock eventually triggers a similar response from the other economy on account of cointegration, offsetting the permanent effect on the rate of change of the exchange rate that would be expected due to interest rate parity conditions.

The analysis of the temporary and permanent monetary policy shocks could also be performed by looking at the deviations from UIRP. In the spirit of Schmitt-Grohé and Uribe (2021), the Supplementary Material File presents the impact on these deviations and the results support their main findings: while the temporary monetary policy shock generates a deviation in favour of the USD and against the predictions of UIRP, the permanent monetary policy shock creates a deviation in favour of the foreign currency.

The results displayed so far evaluate the impact of monetary policy shocks on the rate of change of the exchange rate. It is useful to look at the behaviour of the exchange rates. The accumulated impulse response functions (IRFs) can be looked as proxy for this variable. The Supplementary Material File reports the accumulated IRFs of the five models for permanent and temporary U.S. monetary policy shocks on the rate of change of the exchange rate: a permanent U.S. monetary policy shock results in a depreciation of the USD, while a temporary U.S. monetary policy shock has the opposite consequence. As expected, the accumulated impacts from temporary monetary policy shocks are stronger and significant for a longer period of time, while the accumulated effects of permanent monetary policy shocks are smaller and short-lived.

In order to better understand the importance of the identified structural shocks, in Table 4 we look at the FEVD of the five variables in the model U.S. - G.B. (the FEVD tables for the other models are available in the Supplementary Material File). Clearly, the shock that accounts for the majority of the forecast error variance of $e_t$ is the temporary monetary policy shock in the U.S., explaining more than 50% over 60 months, followed by the shock on itself. From the evidence of the remaining countries, some comments stand out: First, the shock that explains the majority of the forecast error variance of $e_t$ is the shock on this same variable and this result is common across models. Second, the share decreases with the forecast horizon, but after 60 months it continues to explain more than 50%. Third, the second most important shock in this dimension is the temporary U.S. monetary policy shock, explaining between 10% and 28% after 60 months. An important conclusion, which is robust across all models, is that the permanent U.S. monetary policy shock, although explaining the bulk of the forecast error variance of U.S. inflation as well as of nominal interest rates, particularly at longer horizons - as previously documented in Valle e Azevedo *et al.* (2019) -, appears to explain only a small fraction for the rate of change of the exchange rate, thereby hinting that the behaviour of the USD is much more driven by the temporary U.S. monetary policy shock than the permanent one. These conclusions are not in line with the findings of Schmitt-Grohé and Uribe (2021) who report that the permanent monetary policy shocks account for the majority of the forecast error variance of the exchange rate. A possible reason regarding these different outcomes can be related to the assumption of non-stationary of the rate of change of the exchange rate in Schmitt-Grohé and Uribe (2021) together with the imposed cointegration between this variable and the permanent monetary policy shocks. Since we find evidence that this variable is stationary, we treated it as such.

| Var. | Shock | Horizon | | | | | | |
|---|---|---|---|---|---|---|---|---|
| | | 1 | 2 | 4 | 12 | 24 | 48 | 60 |
| | $\varepsilon_e$ | 0.35 | 0.32 | 0.30 | 0.28 | 0.28 | 0.28 | 0.28 |
| | $\varepsilon_{\pi_{US}}$ | 0.01 | 0.01 | 0.01 | 0.02 | 0.02 | 0.02 | 0.02 |
| $e - US/GB$ | $\varepsilon_{i_{GB}}$ | 0.01 | 0.01 | 0.02 | 0.07 | 0.07 | 0.07 | 0.07 |
| | $\varepsilon_{i_{US}}$ | 0.56 | 0.59 | 0.61 | 0.56 | 0.56 | 0.56 | 0.56 |
| | $\varepsilon_{y_{US}}$ | 0.07 | 0.07 | 0.07 | 0.07 | 0.07 | 0.08 | 0.08 |
| | $\varepsilon_e$ | 0.41 | 0.39 | 0.32 | 0.23 | 0.17 | 0.11 | 0.10 |
| | $\varepsilon_{\pi_{US}}$ | 0.21 | 0.21 | 0.31 | 0.45 | 0.58 | 0.72 | 0.75 |
| $\pi_{US}$ | $\varepsilon_{i_{GB}}$ | 0.00 | 0.00 | 0.00 | 0.01 | 0.01 | 0.00 | 0.00 |
| | $\varepsilon_{i_{US}}$ | 0.22 | 0.23 | 0.18 | 0.16 | 0.11 | 0.08 | 0.07 |
| | $\varepsilon_{y_{US}}$ | 0.16 | 0.17 | 0.18 | 0.16 | 0.12 | 0.08 | 0.07 |
| | $\varepsilon_e$ | 0.00 | 0.00 | 0.00 | 0.01 | 0.02 | 0.01 | 0.01 |
| | $\varepsilon_{\pi_{US}}$ | 0.04 | 0.05 | 0.12 | 0.33 | 0.58 | 0.78 | 0.83 |
| $i_{GB}$ | $\varepsilon_{i_{GB}}$ | 0.89 | 0.88 | 0.81 | 0.61 | 0.37 | 0.18 | 0.14 |
| | $\varepsilon_{i_{US}}$ | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 | 0.01 | 0.00 |
| | $\varepsilon_{y_{US}}$ | 0.07 | 0.07 | 0.06 | 0.05 | 0.03 | 0.02 | 0.01 |
| | $\varepsilon_e$ | 0.00 | 0.00 | 0.00 | 0.01 | 0.00 | 0.00 | 0.00 |
| | $\varepsilon_{\pi_{US}}$ | 0.79 | 0.83 | 0.87 | 0.92 | 0.96 | 0.98 | 0.98 |
| $i_{US}$ | $\varepsilon_{i_{GB}}$ | 0.00 | 0.00 | 0.00 | 0.01 | 0.00 | 0.00 | 0.00 |
| | $\varepsilon_{i_{US}}$ | 0.06 | 0.05 | 0.05 | 0.03 | 0.02 | 0.01 | 0.01 |
| | $\varepsilon_{y_{US}}$ | 0.15 | 0.11 | 0.07 | 0.03 | 0.02 | 0.01 | 0.01 |
| | $\varepsilon_e$ | 0.11 | 0.10 | 0.06 | 0.01 | 0.01 | 0.01 | 0.01 |
| | $\varepsilon_{\pi_{US}}$ | 0.00 | 0.00 | 0.01 | 0.00 | 0.00 | 0.00 | 0.00 |
| $y_{US}$ | $\varepsilon_{i_{GB}}$ | 0.03 | 0.03 | 0.04 | 0.04 | 0.02 | 0.01 | 0.01 |
| | $\varepsilon_{i_{US}}$ | 0.01 | 0.01 | 0.01 | 0.01 | 0.01 | 0.01 | 0.00 |
| | $\varepsilon_{y_{US}}$ | 0.85 | 0.85 | 0.87 | 0.95 | 0.96 | 0.98 | 0.98 |

Table 4. Forecast error variance decomposition - U.S. - G.B. model

This analysis was repeated for sub-samples with end dates corresponding to the period when the second economy in each model adopted an explicit inflation target/aim or 1999 for France, Germany and Switzerland. Against the backdrop of inflation targeting regimes, it is arguably harder to identify permanent nominal shocks, as inflation (and nominal rates) would tend to display a stationary behaviour. The sample periods considered are thus: G.B. from 1971M4 to 1992M1; Germany and France from 1971M1 to 1999M1; Australia from 1971M4 to 1996M9 and Switzerland from 1974M1 to 1999M12. The results can be found in the Supplementary Material File. In general, the results are quite similar to those obtained in the full sample. The cointegration parameter for the Fisher relation in the U.S. economy, albeit smaller, continues to be close to, but below, one across all models, with the adjustment to this relation occurring via both inflation and nominal interest rates. The cointegration parameter for the relation between the two nominal interest rates continues to show some dispersion (values from $0.41$ to $098$) and the adjustment to the long-run relation is exclusively made by the nominal interest rate of the second economy. It is important to underline that, in the models with DE and CH, this coefficient is not statistically significant, which can be due to the (almost) stationary behaviour of interest rates even in this sub-sample. Despite these complications, overall, the conclusions from the impulse response analysis broadly confirm the results for the whole sample. The temporary and permanent U.S. monetary policy shocks have similar impacts on inflation and interest rates of the U.S. economy and the impacts on the exchange rate are also much in line with those obtained in the full sample model: the temporary U.S. monetary policy shock leads to an appreciation of the USD against the other currencies, while the permanent U.S. shock results in a USD depreciation. Nevertheless, in the model with CH, the temporary U.S. monetary policy shock appears to have no significant impact on the exchange rate. In the other models this impact is sizeable compared with the impact of the permanent monetary policy shock. In the model with DE, the permanent monetary policy shock has a positive but insignificant impact on the exchange rate, again in contrast to the significance found in the full sample results. In terms of the FEVD, the main conclusions from the full sample model are still valid: the forecast error variance for $e_t$ continues to be mostly explained by the shock to this variable, with the exception of the model with G.B., where the temporary monetary policy shock in the U.S. explains 43% of the variance after 60 months. The permanent U.S. monetary policy shock continues to explain the majority of the forecast error variance of U.S. inflation and nominal interest rates across all models, but the contribution to the forecast error of $e_t$ remains more muted, ranging from 5% to 17% at longer horizons.

## 4. No Cointegration in Monetary Policies – Japan

As discussed in Section 2, Japan's nominal interest rates are not cointegrated with those of the U.S. economy, which requires adjustments in our benchmark model. At the same time, the prolonged experience of Japan in a low inflation and low interest rate environment makes it a relevant case to analyse, particularly after an ample degree of monetary accommodation. A closer look into Japan's interest rates reveals that, from the beginning of the sample until roughly 1994, this variable displays a non-stationary behaviour, which contrasts with the last 25 years, where it becomes fairly stable. The outcome from ADF tests supports this distinction and therefore we treat the two samples separately.

### *4.1. Non-stationary sample*

The model for the sub-sample from 1971M4 to 1994M12 for the Japanese economy is fairly similar to the model discussed in Section 2. The key distinction will be on the cointegration relations. Given the lack of statistical evidence regarding cointegration between interest rates in Japan and in the U.S. economy, also in this sub-sample, we do not impose it. We thus rely on only two cointegration relations: the Fisher relation in the U.S. and the one related to the stationarity of $e_t$. In this setting, Japan's interest rate is included in the model just as a non-stationary variable with no cointegration relations with the other variables. Given the relation between the imposed long-run restrictions and the cointegration relations, this change also implies a modification in the long-run restrictions used to identify the structural shocks. The long-run identification matrix will have the following form:

$$\Xi B = \begin{bmatrix} 0 & 0 & 0 & 0 & 0 \\ 0 & * & 0 & 0 & 0 \\ 0 & * & * & 0 & 0 \\ 0 & * & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & * \end{bmatrix} \tag{7}$$

The only difference with respect to the original long-run identification rests on the third column of matrix $\Xi B$ that now allows Japan's monetary policy shock to have permanent effects only on the nominal interest rate of Japan. This is a necessary assumption given the non-stationary behaviour of this variable and it implies that this monetary policy shock will be considered a permanent shock implying that no distinction is made between temporary and permanent monetary shocks in Japan in this version of the model. Regarding the short-run restriction matrix, given that only two temporary shocks must be distinguished, even fewer restrictions are needed to fully identify the model. It suffices to consider the following specification for $B$:

$$B = \begin{bmatrix} * & * & * & * & * \\ * & * & * & * & * \\ * & * & * & * & * \\ 0 & * & * & * & * \\ * & * & * & * & * \end{bmatrix} \tag{8}$$

Thus, the shock on the exchange rate has no immediate impact on the U.S. interest rate. We could instead assume no immediate impact on Japan's interest rate without significant changes in the main results.

Table 5 displays the estimates of the cointegration parameter and the loading coefficients of the SVEC model. As in the main model in Section 3, the coefficient of the Fisher relation in the U.S. remains statistically significant, with both inflation and U.S. interest rate adjusting to the long-run relation.

Figure 5 presents the estimated impulse response functions. As in the main model, the permanent U.S. monetary policy shock increases permanently inflation and interest rates in the U.S., while the temporary monetary policy shock,

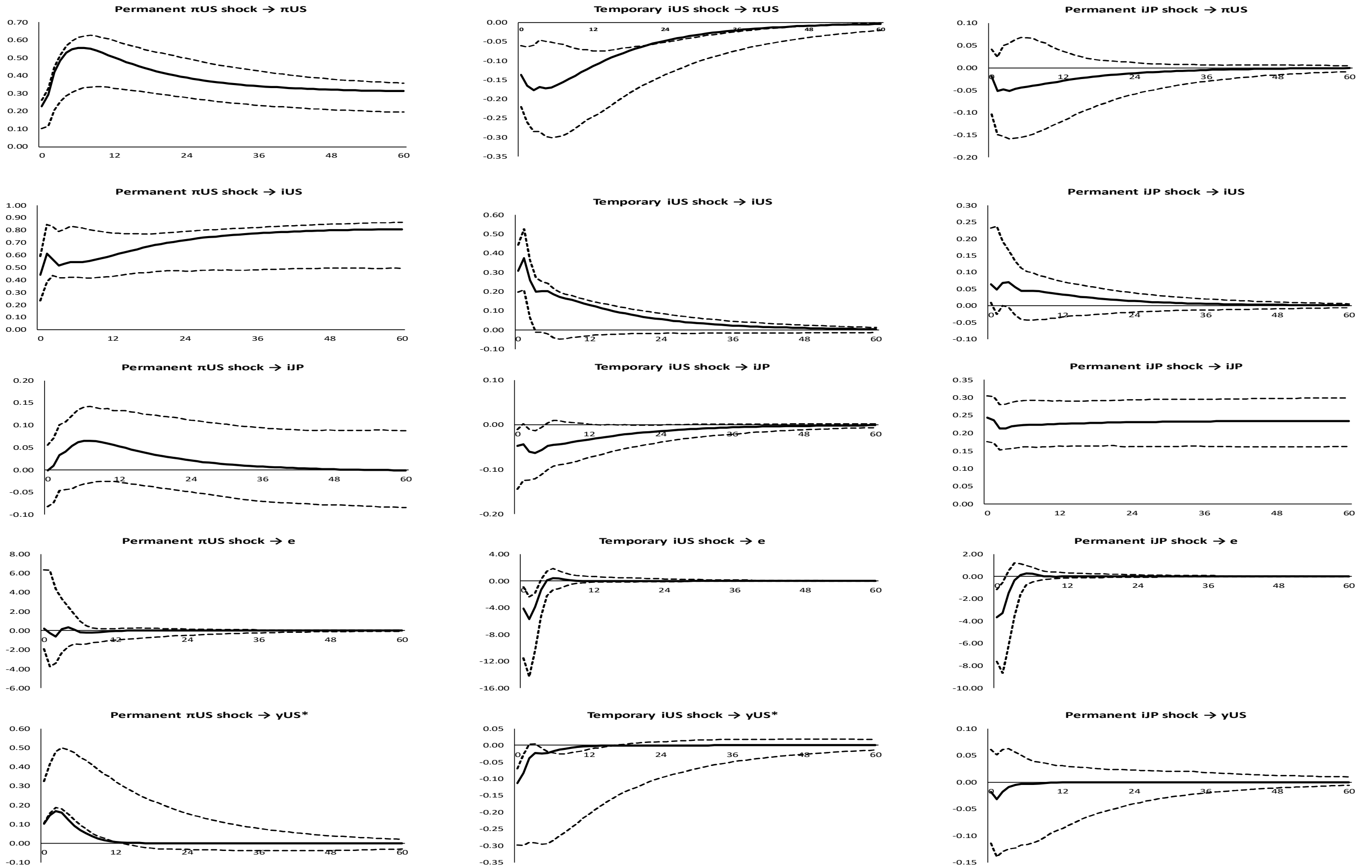
| Non-stationary sample - Fisher relation U.S. | | | | | | |
|---|---|---|---|---|---|---|
| | $\beta^F$ | $\gamma_e$ | $\gamma_{\pi^{US}}$ | $\gamma_{i^*}$ | $\gamma_{i^{US}}$ | $\gamma_{y^{US}}$ |
| U.S. - JP | $-0.38^{***}$ | $0.06$ | $-0.03^{**}$ | $-0.01$ | $0.05^{***}$ | $-0.02^{*}$ |
| (2 lags) | 0.13 | 0.49 | 0.01 | 0.01 | 0.02 | 0.01 |

Notes: $\gamma_{var}$ corresponds to the adjustment parameter in the equation for variable "$var$". Superscript "US" was added to variables $\pi$, $i$ and $y$ variables to make clear they are U.S. variables. Standard errors displayed below the coefficient estimates. Simple two step estimator (S2S) employed as implemented in JMulti. $***$, $**$ and $*$ denote statistical significance at 1%, 5% and 10% levels, respectively. U.S. - Japan model: 1971M4-1994M12. Hannan-Quinn criterion for lag choice.

Table 5. Estimation of Vector Error Correction Model and Cointegration Parameters - Non-stationary sample

associated with a temporary increase in U.S. interest rates, has a negative impact on inflation. But here, given the lack of cointegration between the two interest rates, the permanent U.S. monetary policy shock only increases temporarily Japan's interest rate and not in a statistically significant way. Next, the shock to Japan's interest rate is here a permanent shock associated with a permanent and positive impact on Japan's interest rate and, at the same time, with a positive impact on U.S. interest rates and an insignificant effect on U.S. inflation. As regards the effects on $e_t$, and unlike the original model, the permanent U.S. monetary policy shock seems to have no significant impact. In contrast, the permanent monetary policy shock in Japan leads to an appreciation of the USD. This result remains consistent with the findings of the baseline model where a permanent monetary policy shock that raises interest rates results in a depreciation of the domestic currency. Finally, the temporary U.S. monetary policy shock still has the usual impact, leading to an appreciation of the USD. Overall, these results indicate that our findings of differentiated impacts from temporary and permanent monetary policy shocks are robust to a setting not characterised by monetary policy cointegration.

Regarding the FEVD, the main results from the original model remain valid, as the results in the Supplementary Material File document: most of the forecast error variance of the rate of change of the exchange rate continues to be explained by its own shock, followed by the contribution of the temporary monetary policy shock in the U.S. economy, while the permanent monetary policy shock in the U.S. explains a small share. As before, the permanent U.S. monetary policy shock is the main driver of the forecast error variance of both inflation and interest rates in the U.S. economy. Unlike the original model, the forecast error variance of interest rates in Japan is now mostly explained by Japan's monetary policy shock. Recall that cointegration between interest rates is not imposed in this version of the model.




All responses are in percentage points. The first column has the responses to the permanent U.S. monetary shock, the second column the response to the temporary monetary shock and the third column the response to JP temporary monetary shock. Sample ranges from 1974M1 to 1994M12. Shocks are identified as described in the text. The structural model is estimated by Maximum Likelihood, using the Scoring Algorithm of Amisano and Giannini (1997) as implemented in JMulTi. Response to a one standard deviation shock. 90 % Hall Bootstrap confidence intervals. *Some IRFs lay outside the respective confidence intervals, particularly the response of *y*, which can be attributed to imprecise point estimates. This is bound to occur with bootstrapped confidence intervals, but we highlight that our focus is on the impact of the identified shocks on the remaining variables.

Figure 5: Impulse Response Functions - U.S. - JP model - Non-stationary sample

### *4.2. Stationary sample*

For the sub-sample from 1995M1 to 2017M6, even further changes must be made to the original model to account for the particular nature of the Japanese data. As mentioned before, in this period the Japan's interest rate is well described as stationary, which implies that not only it will not be cointegrated with the U.S. interest rate, but also that it will not be impacted or impact any other variable in the model permanently. To accommodate these features, Japan's interest rate will be treated just like the rate of change of the exchange rate $e_t$. In this vein, the model will continue to have three cointegration relations: the Fisher relation in the U.S., the cointegration relation for the stationary of $e_t$ and another one for Japan's interest rate. These changes imply a different form for the identification matrices. We specify the following long-run matrix:

$$\Xi B = \begin{bmatrix} 0 & 0 & 0 & 0 & 0 \\ 0 & * & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 \\ 0 & * & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & * \end{bmatrix} \tag{9}$$

Compared with the main model, the striking difference resides on the third row of the matrix, which is completely restricted to zero. This follows from the stationarity of Japan's interest rate. Moving on to the short-run identifying restrictions, we impose the following:

$$B = \begin{bmatrix} * & * & * & * & * \\ * & * & 0 & * & * \\ 0 & * & * & * & * \\ 0 & * & * & * & * \\ * & * & * & * & * \end{bmatrix} \tag{10}$$

Given that, as in the original model, three shocks still need to be distinguished, we assume, as before, that the exchange rate shock has no contemporaneous impact on U.S. and Japan's interest rates. We further assume that Japan's (temporary) monetary policy shock has no contemporaneous impact on U.S. inflation (entry (2,3) set to zero). Table 6 displays the estimated cointegration and loading coefficients. In this model, the cointegration coefficient for the Fisher relation is significantly smaller when compared with the main model, which can be attributed to the slow fall in real interest rates observed during this period. Also, the only variable that adjusts to the long-run relation is U.S. inflation.

Figure 6 displays the impulse response functions for this model. Similarly to the model of Section 3, the permanent U.S. monetary policy shock leads to a permanent increase in both inflation and interest rates in the U.S. economy. The temporary U.S. monetary policy shock has no significant impact on U.S. inflation and it decreases slightly Japan's interest rate. As for the temporary Japanese monetary policy shock, associated with a temporary increase in Japan's interest rate, it has no impact on U.S. inflation, while leading to an increase in U.S. interest

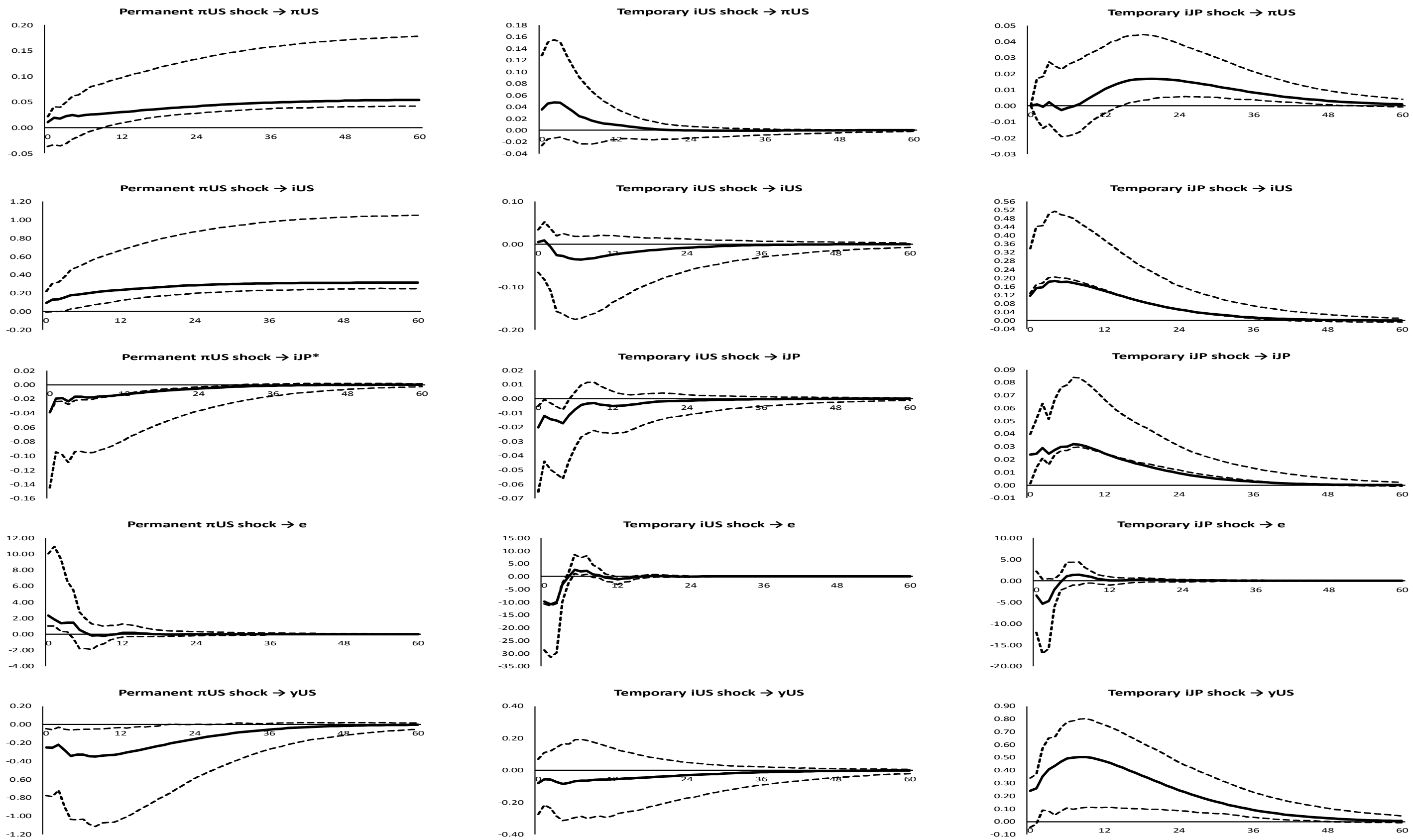
| | Stationary - Fisher relation U.S. | | | | | |
|---|---|---|---|---|---|---|
| | $\beta^F$ | $\gamma_e$ | $\gamma_{\pi^{US}}$ | $\gamma_{i^*}$ | $\gamma_{i^{US}}$ | $\gamma_{y^{US}}$ |
| U.S. - JP | $-0.17^{***}$ | $-4.26^{*}$ | $-0.08^{***}$ | $0.00$ | $0.00$ | $-0.10$ |
| (4 lags) | 0.03 | 2.23 | 0.02 | 0.01 | 0.03 | 0.11 |

Notes: $\gamma_{var}$ corresponds to the adjustment parameter in the equation for variable "$var$". Superscript "US" was added to variables $\pi$, $i$ and $y$ variables to make clear they are U.S. variables. Standard errors displayed below the coefficient estimates. Simple two step estimator (S2S) employed as implemented in JMulti. $***$, $**$ and $*$ denote statistical significance at 1%, 5% and 10% levels, respectively. U.S. - JP model: 1995M1-2017M6. Hannan-Quinn criterion for lag choice.

Table 6. Estimation of Vector Error Correction Model and Cointegration Parameters - Stationary sample

rates. Finally, the impacts on $e_t$ are exactly the ones found in the main model for most economies. The permanent contractionary U.S. monetary policy shock leads to a depreciation of the USD, while the temporary shocks results in the appreciation of the U.S. currency. The temporary Japanese monetary policy shock has no statistically significant impact on $e_t$. These findings, although possibly less reliable on account of estimation concerns, continue to suggest that the findings of the original model remain valid in a more recent sub-sample with a stationary interest rate.

This version of the model also presents some differences regarding the FEVD compared with the original model (Table available in the Supplementary Material File). The forecast error variance of $e_t$ is now mostly explained by the temporary monetary policy shock in the U.S. economy, even after three years, followed by the temporary monetary policy shock in Japan. The shock on $e_t$ itself now only accounts for a residual share. The permanent U.S. monetary policy shock continues to explain a small fraction of the forecast error variance of $e_t$, as concluded in the other models.




All responses are in percentage points. The first column has the responses to the permanent U.S. monetary shock, the second column the response to the temporary monetary shock and the third column the response to JP temporary monetary shock. Sample ranges from 1995M1 to 2017M6. Shocks are identified as described in the text. The structural model is estimated by Maximum Likelihood, using the Scoring Algorithm of Amisano and Giannini (1997) as implemented in JMulTi. Response to a one standard deviation shock. 90 % Hall Bootstrap confidence intervals. * Some IRFs lay outside the respective confidence intervals, which can be attributed to imprecise point estimates. This is bound to occur with bootstrapped confidence intervals. It is important to recall that in this case important variables display little variation.

Figure 6: Impulse Response Functions - U.S. - JP model - Stationary sample

## 5. Extended model – Euro area

So far, the baseline model evaluates the impacts of U.S. temporary and permanent monetary policy shocks on $(i)$ interest rates in both the U.S. economy and the second economy; $(ii)$ U.S. inflation; $(iii)$ U.S. output and $(iv)$ the exchange rate of the USD against the currency of the second economy. A possible extension of the model could be envisaged to account for the distinction between temporary and permanent monetary policy shocks also in the second economy. Accordingly, this Section expands the baseline model by using data for the U.S. and the euro area from 1971M1 to 2019M6 (before 1999 data for Germany is used as a proxy to extend the sample) in a more symmetric SVEC model with seven variables: the five variables in the original model plus inflation and industrial production in the euro area.

This version of the model warrants a reassessment of the cointegration relations to be considered. In the original model the two relations are the Fisher relation in the U.S. and the cointegration of monetary policies (interest rates) of the U.S. and the second economy. When including the inflation rate of the euro area in the model another cointegration ought to be considered: the Fisher relation in the euro area. This feature of the data is important to distinguish between temporary and permanent monetary policy shocks in the euro area. Here, the cointegration between monetary policies is dropped in order to simplify the identification and to allow the distinction between permanent and temporary monetary policy shocks in the two economies.[11]

As a result, the specification of the model in this Section treats the Fisher equations in both the U.S. and the euro area as cointegration relations, and treats the non-stationarity of $e_t$ as another cointegration relation (as in the original model). Based on this, the new vector of variables is $X_t := (e_t, \pi_t^{US}, \pi_t^{EA}, i_t^{US}, i_t^{EA}, y_t^{US}, y_t^{EA})'$, while the vector of structural shocks is $\varepsilon_t := (\varepsilon_t^e, \varepsilon_t^{\pi^{US}}, \varepsilon_t^{\pi^{EA}}, \varepsilon_t^{i^{US}}, \varepsilon_t^{i^{EA}}, \varepsilon_t^{y^{US}}, \varepsilon_t^{y^{EA}})'$, which represents seven serially and mutually uncorrelated structural shocks.

The expanded dimension of the model and the different cointegration relations necessarily imply different identification restrictions. We try as much as possible to give a more symmetric structure to the imposed restrictions, while keeping features of the identification of the original model, using both short and long-run restrictions. The short-run identification $B$ matrix has the following representation:

11. A version of the model with monetary policy cointegration can be considered, but the main results do not change substantially.

$$B = \begin{bmatrix} * & * & * & * & * & * & * \\ * & * & * & * & 0 & * & * \\ * & * & * & 0 & * & * & * \\ 0 & * & 0 & * & * & * & * \\ 0 & 0 & * & * & * & * & * \\ * & * & * & * & * & * & * \\ * & * & * & * & * & * & * \end{bmatrix} \tag{11}$$

In turn, the long-run identification matrix $\Xi B$ is as follows:

$$\Xi B = \begin{bmatrix} 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & * & * & 0 & 0 & 0 & 0 \\ 0 & * & * & 0 & 0 & 0 & 0 \\ 0 & * & * & 0 & 0 & 0 & 0 \\ 0 & * & * & 0 & 0 & 0 & 0 \\ 0 & * & * & 0 & 0 & * & 0 \\ 0 & * & * & 0 & 0 & 0 & * \end{bmatrix} \tag{12}$$

These specifications mean that both a permanent and a temporary monetary policy shock are considered for the U.S. and the euro area (columns 2 to 5 of $\Xi B$). Similarly to the original model, the permanent shocks are the structural shocks on inflation, while the temporary ones are the structural shocks on nominal interest rates. Furthermore, it is assumed that the permanent monetary policy shocks do not have a short-term impact on the interest rate of the other economy (the zeroes in columns 2 and 3 of matrix $B$). Next, the structural shocks to output, just as in the original model, are assumed to have only long-run impacts on the output of the respective economy (columns 6 and 7 of matrix $\Xi B$). Regarding the temporary monetary policy shocks and the shock on $e_t$, the short-run identification matrix has exactly the same structure as in the original model: the shock on $e_t$ has no immediate impact on interest rates (first column of $B$), the temporary U.S. monetary policy shock has no immediate impact on euro area inflation and the euro area temporary monetary policy shock has no immediate impact on U.S. inflation (columns 4 and 5 of $B$).[12]

Tables 7 and 8 display the estimated cointegration and adjustment parameters. As in the original model, the cointegration parameters for the Fisher relation are both statistically significant and close to, but below one, particularly in the U.S. economy. In the euro area, this parameter stands slightly lower at $0.57$. The adjustment parameters in the case of the U.S. suggests that both inflation and the interest rate adjust to re-establish the long-run relation, while in the case of the euro area only inflation seems to react. Other estimates for the adjustment parameters are also statistically significant at the conventional levels, particularly

12. The restrictions related to money neutrality are not needed and were not imposed in this version of the model to facilitate the convergence of the estimation process.

| | Fisher relation U.S. | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| | $\beta^{F_{US}}$ | $\gamma_e$ | $\gamma_{\pi^{US}}$ | $\gamma_{\pi^{EA}}$ | $\gamma_{i^{US}}$ | $\gamma_{i^{EA}}$ | $\gamma_{y^{US}}$ | $\gamma_{y^{EA}}$ |
| U.S. - EA | $-0.78^{***}$ | $-0.60$ | $-0.02^{**}$ | $0.03^{***}$ | $0.03^{**}$ | $-0.02^{**}$ | $-0.05^{***}$ | $-0.01$ |
| (2 lags) | 0.06 | 0.43 | 0.01 | 0.01 | 0.01 | 0.01 | 0.02 | 0.04 |

Notes: $\gamma_{var}$ corresponds to the adjustment parameter in the equation for variable "$var$". Superscript "US" or "EA" was added to variables $\pi$, $i$ and $y$ variables to make clear they are U.S. or euro area variables. Standard errors displayed below the coefficient estimates. Simple two step estimator (S2S) employed as implemented in JMulti. $***$, $**$ and $*$ denote statistical significance at 1%, 5% and 10% levels, respectively. U.S. - EA model: 1971M1-2019M6, data for Germany before 1999. Hannan-Quinn criterion for lag choice.

Table 7. Estimation of Vector Error Correction Model and Cointegration Parameters

| | Fisher relation EA | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| | $\beta^{F_{EA}}$ | $\gamma_e$ | $\gamma_{\pi^{US}}$ | $\gamma_{\pi^{EA}}$ | $\gamma_{i^{US}}$ | $\gamma_{i^{EA}}$ | $\gamma_{y^{US}}$ | $\gamma_{y^{EA}}$ |
| U.S. - EA | $-0.57^{***}$ | $-0.70$ | $-0.01$ | $-0.09^{***}$ | $-0.02$ | $0.02$ | $0.04$ | $0.01$ |
| (2 lags) | 0.04 | 0.70 | 0.01 | 0.01 | 0.02 | 0.02 | 0.03 | 0.07 |

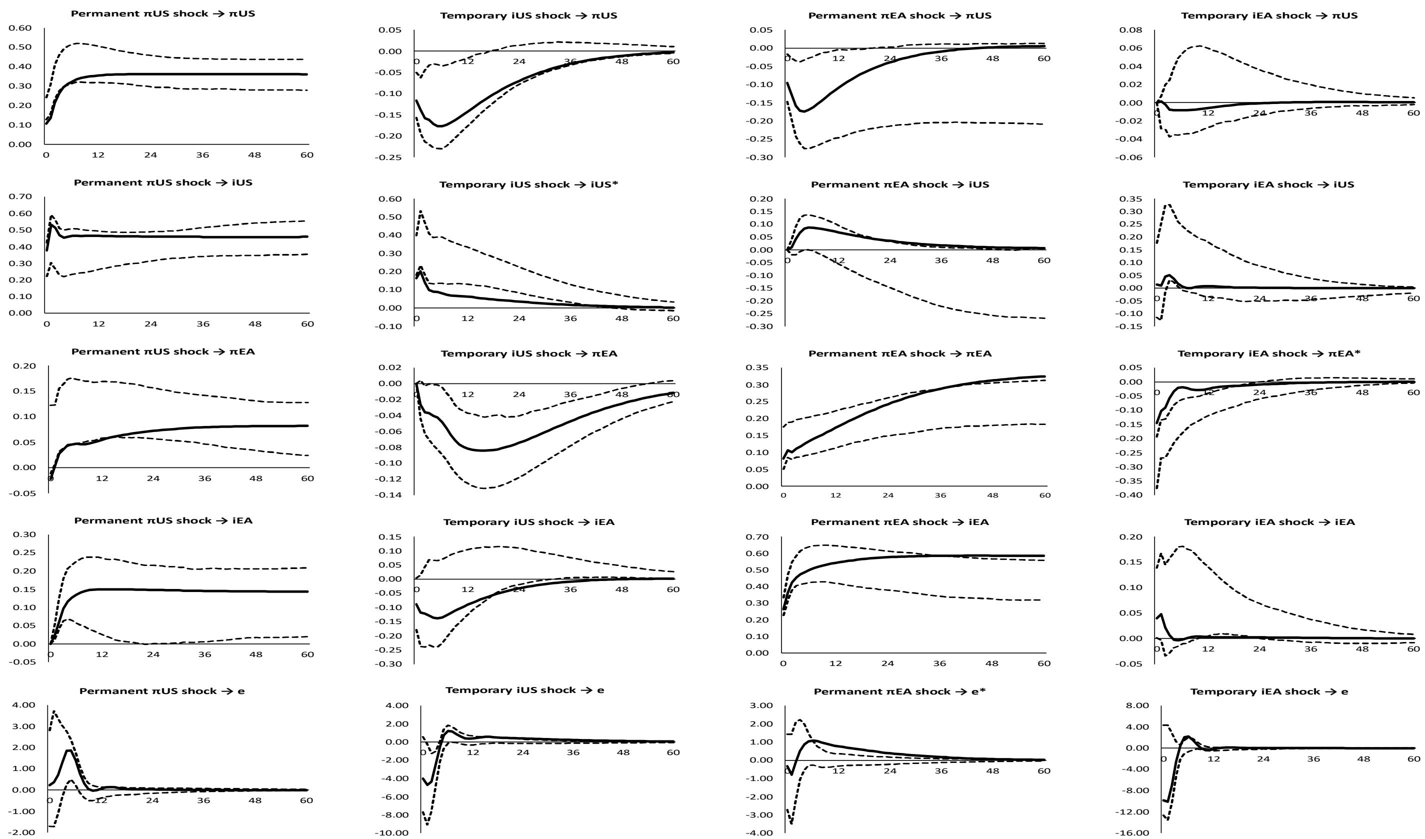
Notes: $\gamma_{var}$ corresponds to the adjustment parameter in the equation for variable "$var$". Superscript "US" or "EA" was added to variables $\pi$, $i$ and $y$ variables to make clear they are U.S. or euro area variables. Standard errors displayed below the coefficient estimates. Simple two step estimator (S2S) employed as implemented in JMulti. $***$, $**$ and $*$ denote statistical significance at 1%, 5% and 10% levels, respectively. U.S. - EA model: 1971M1-2019M6, data for Germany before 1999. Hannan-Quinn criterion for lag choice.

Table 8. Estimation of Vector Error Correction Model and Cointegration Parameters

in response to deviations from the Fisher relation in the U.S. economy, even for variables assumed to have a zero coefficient in the cointegration equation.

Figure 7 displays the IRFs from the U.S. - EA model with the four columns exhibiting the impacts of the permanent and temporary U.S. monetary policy shocks, followed by the impacts of the permanent and temporary euro area monetary policy shocks. The inspection of this figure shows that the temporary and permanent monetary policy shocks have the expected effects on each region's inflation and interest rates: the temporary monetary policy shock increases the domestic interest rate and has a negative impact on domestic inflation, while the permanent monetary policy shock has a positive impact on both interest rates and inflation, also in the short run. Regarding the impact on the foreign economy the results are mixed. While the temporary monetary policy shock in the euro area has no impact on U.S. interest rates or inflation, the temporary shock in the U.S. has a negative impact on euro area inflation. The permanent U.S. monetary policy shock has a limited impact on the nominal variables in the euro area, while the permanent euro area shock has an immediate positive impact on the U.S. interest rate. As for the impact of the shocks on $e_t$, there is evidence that the impacts of the monetary policy shocks in the U.S. (both permanent and temporary) are consistent with the results in the baseline model: the temporary U.S. monetary policy shock leads to an appreciation of the USD, while the permanent U.S. monetary policy shock has the opposite effect, resulting in an depreciation of the USD. As in the baseline model, the impact of the temporary shock appears to be stronger than the impact of the permanent monetary policy shock. The impacts of the euro area monetary policy shocks on the USD are not statistically significant, but in contrast with the generic findings from the baseline model the point estimates indicate a negative impact of the temporary monetary policy shock on the USD.

Looking at the FEVD (Table available in the Supplementary Material File), the forecast error variance of $e_t$ is strongly explained by the temporary monetary policy shocks (together they account for 73% after three years), with the shock on this variable explaining together 18% after three years. The permanent monetary policy shocks continue to explain only a residual amount (6% after three years), in line with the results of the baseline model. The permanent monetary policy shocks in both economies continues to be the main driver of the FEVD of the respective interest rate and inflation at longer horizons.




All responses are in percentage points. The first column has the responses to the permanent U.S. monetary shock, the second column the response to the temporary U.S. monetary shock, the third column the response to the EA permanent shock and the fourth column the response to the EA temporary monetary shock. Sample ranges from 1971M1 to 2019M6. Shocks are identified as described in the text. The structural model is estimated by Maximum Likelihood, using the Scoring Algorithm of Amisano and Giannini (1997) as implemented in JMulTi. Response to a one standard deviation shock. 90 % Hall Bootstrap CI. *Some IRFs lay outside the respective confidence intervals, which can be attributed to imprecise point estimates. This is bound to occur with bootstrapped confidence intervals. We recall that in this version of the model a large number of parameters needs to be estimated.

Figure 7: Impulse Response Functions - U.S. - EA model

## 6. Conclusion

We find compelling evidence that permanent and temporary monetary policy shocks in the U.S. economy have opposite impacts on the USD against several other currencies: a shock associated with a temporary increase in U.S. interest rates leads to an appreciation of the USD, whereas a shock associated with a permanent increase in U.S. interest rates gears a depreciation of the USD, i.e., in the direction of UIRP. This helps reconciling the literature that investigates the effects of monetary policy shocks on the behaviour of exchange rates - which tends to report an appreciation of the domestic currency after a contractionary monetary shock -, with theoretical arguments and empirical literature that finds exchange rate movements consistent with interest rare parity conditions. Ignoring the difference between the two types of shocks may lead to puzzles. Also, we confirm the "neo-Fisher" effect in models that open the U.S. economy, as a permanent monetary policy shock associated with a permanent increase in interest rates can also generate higher inflation - along with the depreciation of the domestic currency - even in the short run, just as long-run forces would dictate.

We thus confirm the results of Schmitt-Grohé and Uribe (2021), who first explored the distinction between permanent and temporary shocks in explaining exchange rate dynamics. We employ less stringent identifying restrictions and resort to a fairly simple structural model. We take advantage of two cointegration relations that characterise the data and help identifying the structural shocks: $(i)$ the Fisher relation and $(ii)$ the cointegration between monetary policies (or interest rates) across two advanced economies. We additionally impose standard neutrality restrictions and some other short-run (i.e. on impact) weak identifying assumptions. Unlike Schmitt-Grohé and Uribe (2021), we do not need to impose the standard effects of temporary monetary policy shocks to identify the various shocks. We perform various robustness checks and analyse several departures from this standard setting, dictated by the properties of the data (e.g. the model with Japan) or by the desirability of a more symmetric treatment of the currencies (e.g. when we consider the U.S. and the euro area). The main results are remarkably stable over these different specifications and also across sub-samples.

## References

Amisano, G. and C. Giannini (1997). *Topics in Structural VAR Econometrics*. Springer, 2nd edn., Berlin.

Arouri, M., F. Jawadi, and N. Nguyen (2013). "What can we tell about monetary policy synchronization and interdependence over the 2007-2009 global financial crisis?" *Journal of Macroeconomics*, 36, 175–187.

Belke, A. and Y. Cui (2010). "US–Euro Area Monetary Policy Interdependence: New Evidence from Taylor Rule-based VECMs." *The World Economy*, 33(5), 778–797.

Belke, A. and D. Gros (2005). "Asymmetries in Transatlantic Monetary Policymaking:Does the ECB Follow the Fed?" *Journal of Common Market Studies*, 43(5), 921–946.

Bjornland, H. C. (2008). "Monetary Policy and Exchange Rate Interactions in a Small Open Economy." *The Scandinavian Journal of Economics*, 110(1), 197–221.

De Michelis, A. and M. Iacoviello (2016). "Raising an inflation target: The Japanese experience with Abenomics." *European Economics Review*, 88, 67–87.

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

Fama, E. F. (1984). "Forward and spot exchange rates." *Journal of Monetary Economics*, 14(3), 319–338.

Froot, K. and J. Frankel (1989). "Forward Discount Bias: Is It an Exchange Risk Premium?" *The Quarterly Journal of Economics*, 104(1), 139–161.

Gray, C. (2013). "Responding to a Monetary Superpower: Investigating the Behavioral Spillovers of U.S. Monetary Policy." *Atlantic Economic Journal*, 41(2), 173–184.

Inoue, A. and B. Rossi (2019). "The effects of conventional and unconventional monetary policy on exchange rates." *Journal of International Economics*, 118, 419–447.

Kim, S. and N. Roubini (2000). "Exchange rate anomalies in the industrial countries: A solution with a structural VAR approach." *Journal of Monetary Economics*, 45(3), 561–586.

Lothian, J. R. (2016). "Uncovered interest parity: The long and the short of it." *Journal of Empirical Finance*, 36(1), 1–7.

Lütkepohl, H. (2006). *New Introduction to Multiple Time Series Analysis*. Springer, Berlin and Heidelberg.

Miller, N. C. (2014). *Exchange rate economics: the uncovered interest parity puzzle and other anomalies*. Edward Elgar Publishing.

Schmitt-Grohé, S. and M. Uribe (2021). "Exchange Rates and Uncovered Interest Differentials: The Role of Permanent Monetary Shocks." *Journal of International Economics*, (forthcoming).

Scholl, A. and H. Uhlig (2008). "New evidence on the puzzles: Results from agnostic identification on monetary policy and exchange rates." *Journal of International Economics*, 76(1), 1–13.

Uribe, M. (2021). "The Neo-Fisher Effect: Econometric Evidence from Empirical and Optimizing Models." *American Economic Journal: Macroeconomics*, (forthcoming).

Valle e Azevedo, J., J. Ritto, and P. Teles (2019). "The Neutrality of Nominal Rates: How Long is the Long Run?" *Banco de Portugal Working Papers*, (11).

Zettelmeyer, J. (2004). "The impact of monetary policy on the exchange rate: evidence from three small open economies." *Journal of Monetary Economics*, 51(3), 635–652.

## Appendix

### *Data description*

We rely on information for the U.S. and seven open advanced economies, namely G.B., Germany, France, Australia, Switzerland, Japan and the euro area[13]. The data is collected from the Federal Reserve Bank of St. Louis' FRED website, Eurostat and the OECD. The period considered for each country is dependent on the respective availability, ranging from 1971 - the year the Bretton Woods exchange rates system came to an end - until 2019. Table 9 describes the variables considered in the empirical analysis.

| Variable | Symbol | Description | Source | Time-span |
|---|---|---|---|---|
| Output | $y_t$ | U.S. : Monthly index of production in total industry | FRED | 1971M1 - 2019M6 |
| | | EA : Monthly index of production in total industry | Eurostat | 1994M1 - 2019M6 |
| Inflation | $\pi_t$ | U.S. : Year-on-year growth rate of the CPI excluding food and energy | FRED | 1971M1 - 2019M6 |
| | | EA : Year-on-year growth rate of the HICP excluding food and energy | Eurostat | 2001M12 - 2019M6 |
| Interest rates | $i_t$ | U.S. : 3-month treasury bills rates in the secondary market | FRED | 1971M1 - 2019M6 |
| | | G.B. : 3-month treasury securities rates | FRED | 1971M1 - 2017M6 |
| | | DE : 3-month money market rate | FRED | 1971M1 - 2019M6 |
| | | FR : 3-month money market rate | FRED | 1971M1 - 2019M5 |
| | | AU : 3-month bank bills rates | FRED | 1971M1 - 2019M6 |
| | | CH : 3-month eurodollar deposit rate | FRED | 1974M1 - 2019M6 |
| | | JP : 3-month treasury bills | FRED | 1974M1 - 2017M6 |
| | | EA : 3-month money market rate | FRED | 1994M1 - 2019M6 |
| Exchange rate | $e_t$ | 3-month annualised growth rate based on a monthly series | FRED/ OECD | 1971M1 - 2019M6 |

Note: Data for the DE and FR exchange rate was collected from the OECD currency conversions.

Table 9. Description and sources of the variables used in the empirical analysis

13. Initially, the analysis has also considered data for Canada, which was later discarded given it is recognised in the literature as having unexpected results from monetary policy shocks on the exchange rate (see, e.g., Inoue and Rossi (2019)).

# Permanent and temporary monetary policy shocks and the dynamics of exchange rates

## Supplementary Material File – Further results

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**Abstract**
This appendix discusses in more detail the estimation of the structural VEC model for Germany, France, Australia and Switzerland. We also consider the estimation of the model, for all the economies under analysis, using a pre- inflation targeting sub-sample. In the various cases we provide tables with Forecast Error Variance Decompositions. Finally, for the full sample we present results for the accumulated impulse reponse functions (to gauge the effects of shocks on exchange rates themselves, rather than on the variation in exchange rates) as well as for the effects of shocks on deviations from uncovered interest rate parity.

JEL: E52, E58, F31, C32

Keywords: Exchange Rates, Fisher Relation, Monetary Policy Cointegration, Monetary Shocks, Structural VEC Models.

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## 1. Methodology

To measure the impact of temporary and permanent monetary policy shocks on the dynamics of exchange rates we follow closely the SVEC model methodology and identification strategy exposed in Lütkepohl (2006). We consider the U.S. economy together with the United Kingdom (G.B.), Germany (DE), France (FR), Australia (AU), Switzerland (CH), Japan (JP) and the euro area (EA). In most specifications the empirical analysis relies on five variables:

- 3-month annualised rate of change of the exchange rate between the USD and the currency of the respective advanced economy - $e_t$;
- 3-month interest rates extracted from Treasury bills or money market instruments, depending on data availability, for both the U.S. and for the other advanced economy, respectively $i_t$ and $i_t^*$;
- U.S. core inflation, measured by the year-on-year rate of change of the CPI excluding food and energy - $\pi_t$;
- U.S. industrial production index, as a proxy for output at a monthly frequency - $y_t$.[1]

The data set is collected from the Federal Reserve Bank of St. Louis' FRED website, Eurostat and the OECD. The period considered for each country is dependent on the respective availability, ranging from 1971 – the year the Bretton Woods exchange rates system came to an end – until 2019. A complete description of the data set can be found in Table 1.

In our sample all variables are non-stationary, with the exception of $e_t$, which is found to be I(0).[2]

Next, Table 2 , also reported in the main text, reveals the results of a Johansen trace test for cointegration between the nominal interest rate of the U.S. and that of the other economies. The outcomes of the tests clearly point to a cointegration in monetary policies (i.e. of policy interest rates *vis-à-vis* the U.S. nominal rate) across the economies in our sample, with the exception of JP.[3] Given this evidence we henceforth assume that nominal interest rates are cointegrated with those of the U.S. (except for JP), though not necessarily with a coefficient equal to unity. As previously mentioned, our model is similar to the one developed by Schmitt-Grohé and Uribe (2021) but in their main model it is assumed that $e_t$ is non-stationary.

1. Industrial production growth correlates well with GDP growth even in the context of a decreasing contribution of manufacturing to output. Is has the advantage of being available for large time spans.

2. Based on standard ADF tests there is evidence that both inflation and output are I(1) in all countries. The nominal interest rates also emerge as non-stationary, but not as evidently in the case of DE and CH. A remarking conclusion is that the rate of change of the exchange rate of the USD against the other currencies is undoubtedly stationary. In view of including this variable in the VECM methodology, it is assumed that it does not have any long-term relation with any other variables in the model.

3. For further evidence of cointegration in monetary policies, see Belke and Cui (2010) and Arouri *et al.* (2013).

| Variable | Symbol | Description | Source | Time-span |
|---|---|---|---|---|
| Output | $y_t$ | U.S. : Monthly index of production in total industry | FRED | 1971M1 - 2019M6 |
| | | EA : Monthly index of production in total industry | Eurostat | 1994M1 - 2019M6 |
| Inflation | $\pi_t$ | U.S. : Year-on-year growth rate of the CPI excluding food and energy | FRED | 1971M1 - 2019M6 |
| | | EA : Year-on-year growth rate of the HICP excluding food and energy | Eurostat | 2001M12 - 2019M6 |
| Interest rates | $i_t$ | U.S. : 3-month treasury bills rates in the secondary market | FRED | 1971M1 - 2019M6 |
| | | G.B. : 3-month treasury securities rates | FRED | 1971M1 - 2017M6 |
| | | DE : 3-month money market rate | FRED | 1971M1 - 2019M6 |
| | | FR : 3-month money market rate | FRED | 1971M1 - 2019M5 |
| | | AU : 3-month bank bills rates | FRED | 1971M1 - 2019M6 |
| | | CH : 3-month eurodollar deposit rate | FRED | 1974M1 - 2019M6 |
| | | JP : 3-month treasury bills | FRED | 1974M1 - 2017M6 |
| | | EA : 3-month money market rate | FRED | 1994M1 - 2019M6 |
| Exchange rate | $e_t$ | 3-month annualised growth rate based on a monthly series | FRED/ OECD | 1971M1 - 2019M6 |

Note: Data for the DE and FR exchange rate was collected from the OECD currency conversions.

Table 1. Description and sources of the variables used in the empirical analysis

Finally, as documented in Valle e Azevedo *et al.* (2019), nominal interest rates and inflation rates in advanced economies appear to be cointegrated, so we take this assumption for U.S. inflation and nominal interest rates, although the coefficient is not restricted to one, again unlike Schmitt-Grohé and Uribe (2021).

| Johansen Trace Tests | U.S./G.B. 3 lags | U.S./DE 2 lags | U.S./FR 3 lags | U.S./AU 3 lags | U.S./CH 7 lags | U.S./JP 9 lags |
|---|---|---|---|---|---|---|
| 0 | 0.02 | 0.02 | 0.00 | 0.01 | 0.00 | 0.46 |
| 1 | 0.80 | 0.52 | 0.49 | 0.58 | 0.72 | 0.73 |

Note: The Table displays p-values for the null of "at most 0 (or 1) cointegration relations". Specification with constant and trend. AIC for lag choice.

Table 2. Johansen trace tests for monetary policy cointegration

The basic reduced form of the VEC model for all countries is given by:

$$\Delta X_t = \alpha_0 + \gamma\beta X_{t-1} + \sum_{j=1}^{K} \beta_j \Delta X_{t-j} + u_t \tag{1}$$

where $X_t := (e_t, \pi_t, i_t^*, i_t, y_t)'$. $\gamma$ and $\beta$ correspond to the loading coefficients and cointegration coefficients matrices, respectively, and $u_t := (u_t^e, u_t^\pi, u_t^{i^*}, u_t^i, u_t^y)'$ is a vector of reduced form serially uncorrelated shocks. Equivalently, focusing on the cointegration part (second term):

$$\begin{cases}\Delta e_t = \gamma_{11}e_{t-1} + \gamma_{12}(\pi_{t-1} - \beta^F i_{t-1}) + \gamma_{13}(i^*_{t-1} - \beta^{MP} i_{t-1}) + ... + u^e_t \\ \Delta \pi_t = \gamma_{21}e_{t-1} + \gamma_{22}(\pi_{t-1} - \beta^F i_{t-1}) + \gamma_{23}(i^*_{t-1} - \beta^{MP} i_{t-1}) + ... + u^\pi_t \\ \Delta i^*_t = \gamma_{31}e_{t-1} + \gamma_{32}(\pi_{t-1} - \beta^F i_{t-1}) + \gamma_{33}(i^*_{t-1} - \beta^{MP} i_{t-1}) + ... + u^{i^*}_t \\ \Delta i_t = \gamma_{41}e_{t-1} + \gamma_{42}(\pi_{t-1} - \beta^F i_{t-1}) + \gamma_{43}(i^*_{t-1} - \beta^{MP} i_{t-1}) + ... + u^i_t \\ \Delta y_t = \gamma_{51}e_{t-1} + \gamma_{52}(\pi_{t-1} - \beta^F i_{t-1}) + \gamma_{53}(i^*_{t-1} - \beta^{MP} i_{t-1}) + ... + u^y_t \end{cases} \tag{2}$$

where the superscript "$F$" in $\beta^F$ refers to the Fisher relation whereas "$MP$" refers to monetary policy cointegration. The treatment of $e_t$ follows from its stationarity. To recover the structural form shocks it is necessary to impose some identifying assumptions on a non-singular matrix $B$, such that $u_t = B\varepsilon_t$, where $\varepsilon_t := (\varepsilon^e_t, \varepsilon^\pi_t, \varepsilon^{i^*}_t, \varepsilon^i_t, \varepsilon^y_t)'$ represents five serially and mutually uncorrelated structural shocks. Following the steps in Lütkepohl (2006), (Chapter 9), the VEC model in its reduced form implies the following Beveridge-Nelson decomposition:

$$X_t = X^*_{0,t} + \Xi \sum_{i=1}^{t} u_i + \sum_{j=0}^{\infty} \Xi^*_j u_{t-j} \tag{3}$$

where the first term collects initial values and deterministic trends, the second term accounts for the long-run effects of the shocks and the last term is absolutely summable and thus stationary. Given the relation between the reduced form and the structural form shocks, this decomposition can be expressed as:

$$X_t = X^*_{0,t} + \Xi B \sum_{i=1}^{t} \varepsilon_i + \sum_{j=0}^{\infty} \Xi^*_j B \varepsilon_{t-j} \tag{4}$$

Lütkepohl (2006) (Chapther 6) shows that matrix $\Xi$, a function of the reduced form parameters, has reduced rank in the presence of cointegration. In our baseline model with five variables and three cointegration relations (the stationarity of $e_t$ is conveniently treated as one, together with cointegration between $i_t$ and $i^*_t$ and cointegration between $i_t$ and $\pi_t$), the rank of matrix $\Xi$ is two. Given that matrix $B$ is non-singular, the rank of matrix $\Xi B$ is also two. This is consistent with stating that our model has two stochastic trends driving the data: $(i)$ the trend that gears inflation and the short-term interest rate in the U.S. as well as the short-term interest rate in the other advanced economy and $(ii)$ the trend that drives output in the U.S. economy. The rate of change of the exchange rate has no stochastic trend given its stationary behaviour. All this helps in the identification of $B$, since we can make assumptions on the long-run impact matrix $\Xi B$ to make sure it has rank two. We consider that the structural shock to the inflation rate of the U.S. is a permanent one, and also the one driving the two short-term interest rates in the model. The structural shocks to the interest rates thus only have temporary effects.

[4] This assumption, together with the stationarity of $e_t$, implies that the first, third and fourth columns of $\Xi B$ are zeroes. The first row of $\Xi B$ is also zero since the rate of change of the exchange rates cannot be affected by any shock in the model in the long run. Two other assumptions are made in $\Xi B$: $(i)$ the first simply assumes that the permanent monetary policy shock has no long-run impact on output (entry (5,2)), a standard neutrality proposition and; $(ii)$ the permanent output shock has no long-run impact on the level of nominal rates or inflation (entries (2,5), (3,5) and (4,5)).[5] Given all this, matrix $\Xi B$ has the following structure:

$$\Xi B = \begin{bmatrix} 0 & 0 & 0 & 0 & 0 \\ 0 & * & 0 & 0 & 0 \\ 0 & * & 0 & 0 & 0 \\ 0 & * & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & * \end{bmatrix} \tag{5}$$

This setting is helpful, but it is not enough to fully identify $B$, since the three temporary shocks must be distinguished and this can only be done directly on $B$. We assume that the structural shock on the rate of change of the exchange rate has no contemporaneous impact on the two nominal interest rates and that the temporary monetary policy shock of the other advanced economy has no contemporaneous impact on the inflation rate of the U.S. economy. This implies a $B$ matrix of the following form:

$$B = \begin{bmatrix} * & * & * & * & * \\ * & * & 0 & * & * \\ 0 & * & * & * & * \\ 0 & * & * & * & * \\ * & * & * & * & * \end{bmatrix} \tag{6}$$

Combining these short-run and long-run identification restrictions we obtain sufficient conditions for the estimation of the structural model.

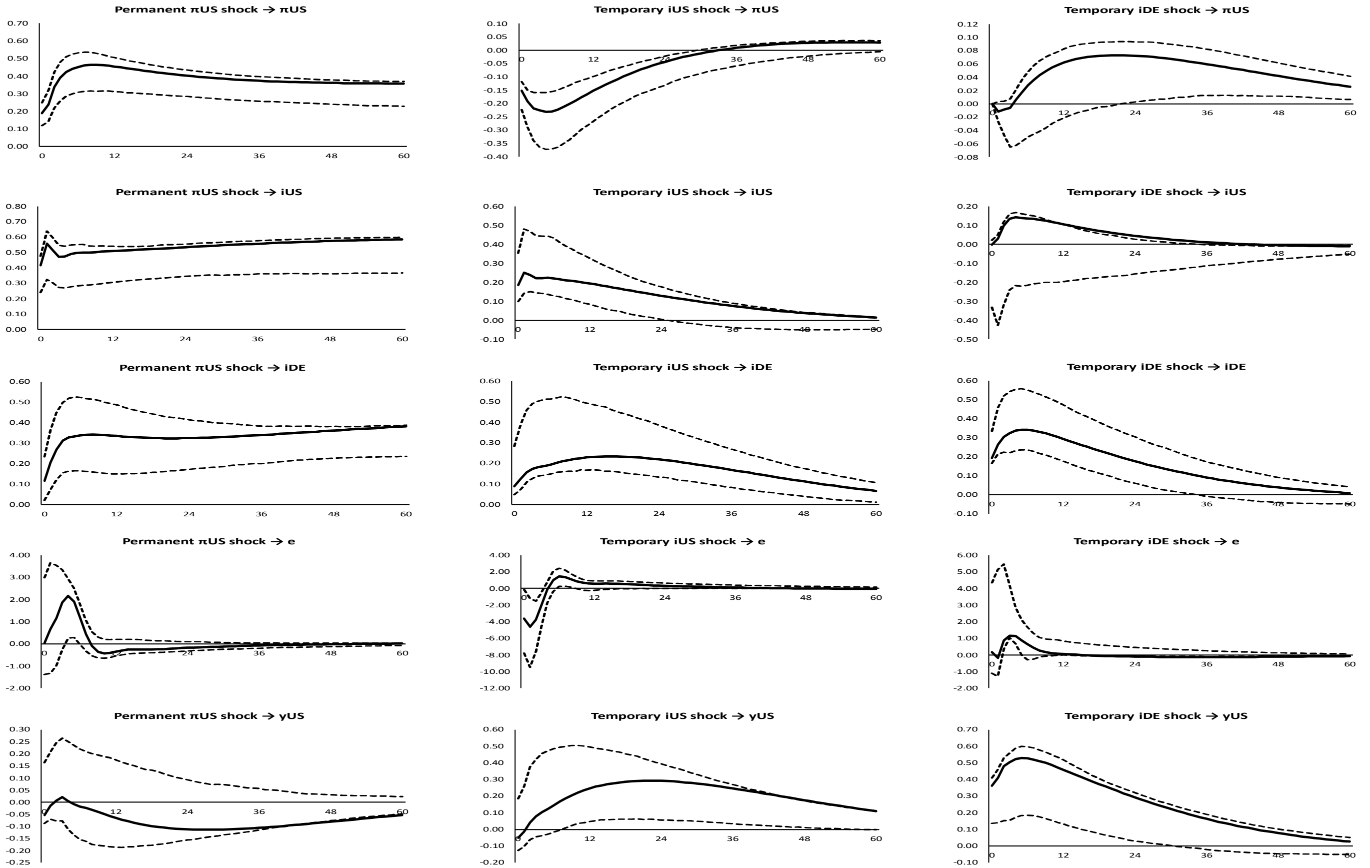
4. This permanent shock only has permanent effects on both inflation and the two nominal rates. This choice is innocuous and we could have picked either the U.S. or foreign nominal interest shock as the permanent shock and the inflation shock as the other transitory shock; in this way we would obtain exactly the same impulse response functions as those obtained with the alternative identification, only the labelling of shocks would be switched. The meaningful assumption is that there is only one permanent shock, instead of two or three such shocks with "collinear" effects on these three variables.

5. This actually results in overidentification of the permanent shocks, but overall the results are similar if we assume there is no long-run effect on nominal rates (only) or on inflation (only).

## 2. Robustness analysis

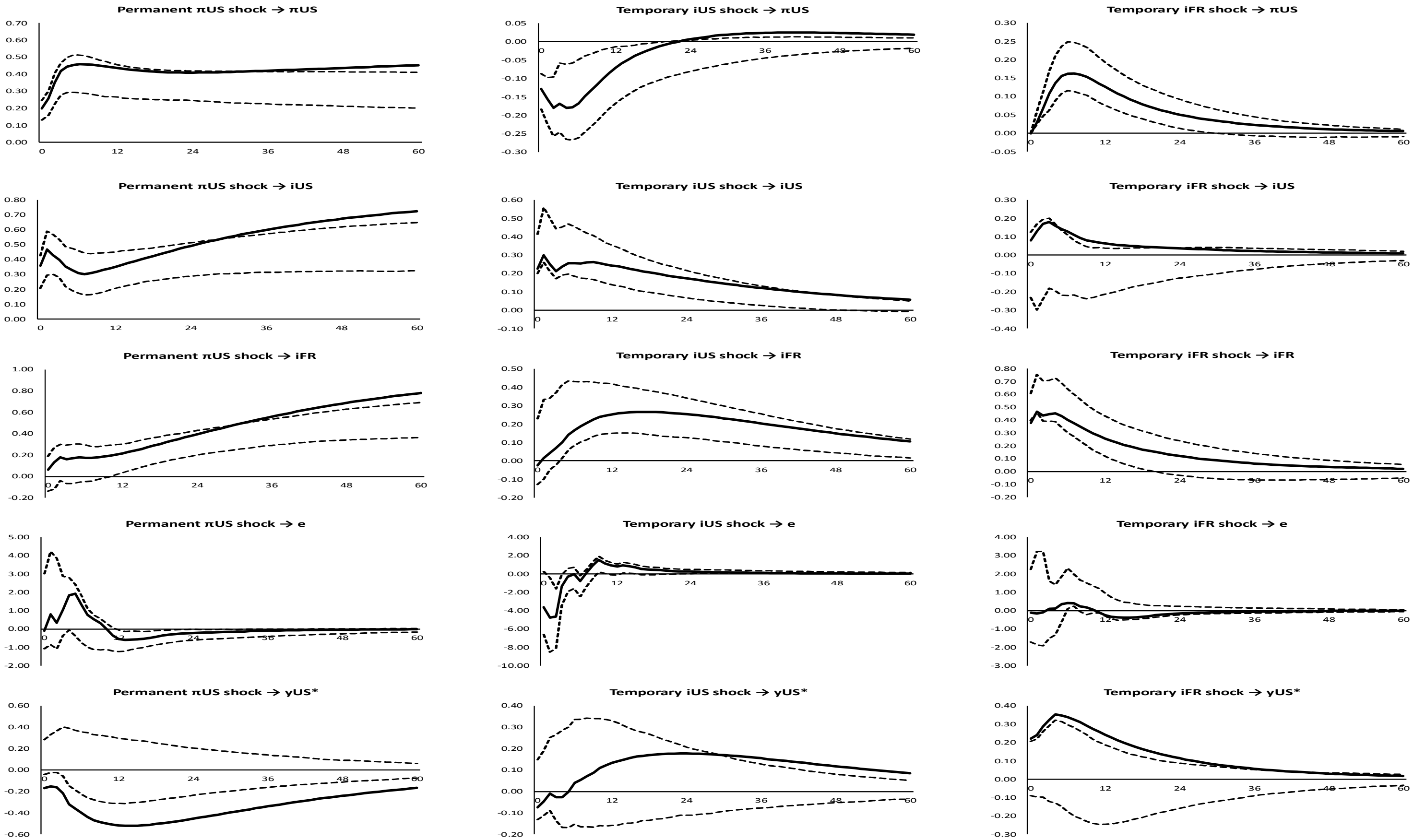
### *2.1. Other Countries - Full Sample*

Figures 1, 2, 3 and 4 display for DE, FR, AU and CH the responses of nominal interest rates, U.S. inflation, the rate of change of the exchange rate and U.S. output to the identified structural shocks. The first column shows the response to the permanent U.S. monetary policy shock, the second column focuses on the U.S. temporary monetary policy shock and the third column reports the response to temporary monetary policy shock on the other advanced economy. We recall that the dataset for DE and FR covers 1971M1-2019M6, for AU it covers 1971M4-2019M6 and for CH 1974M1-2019M6.



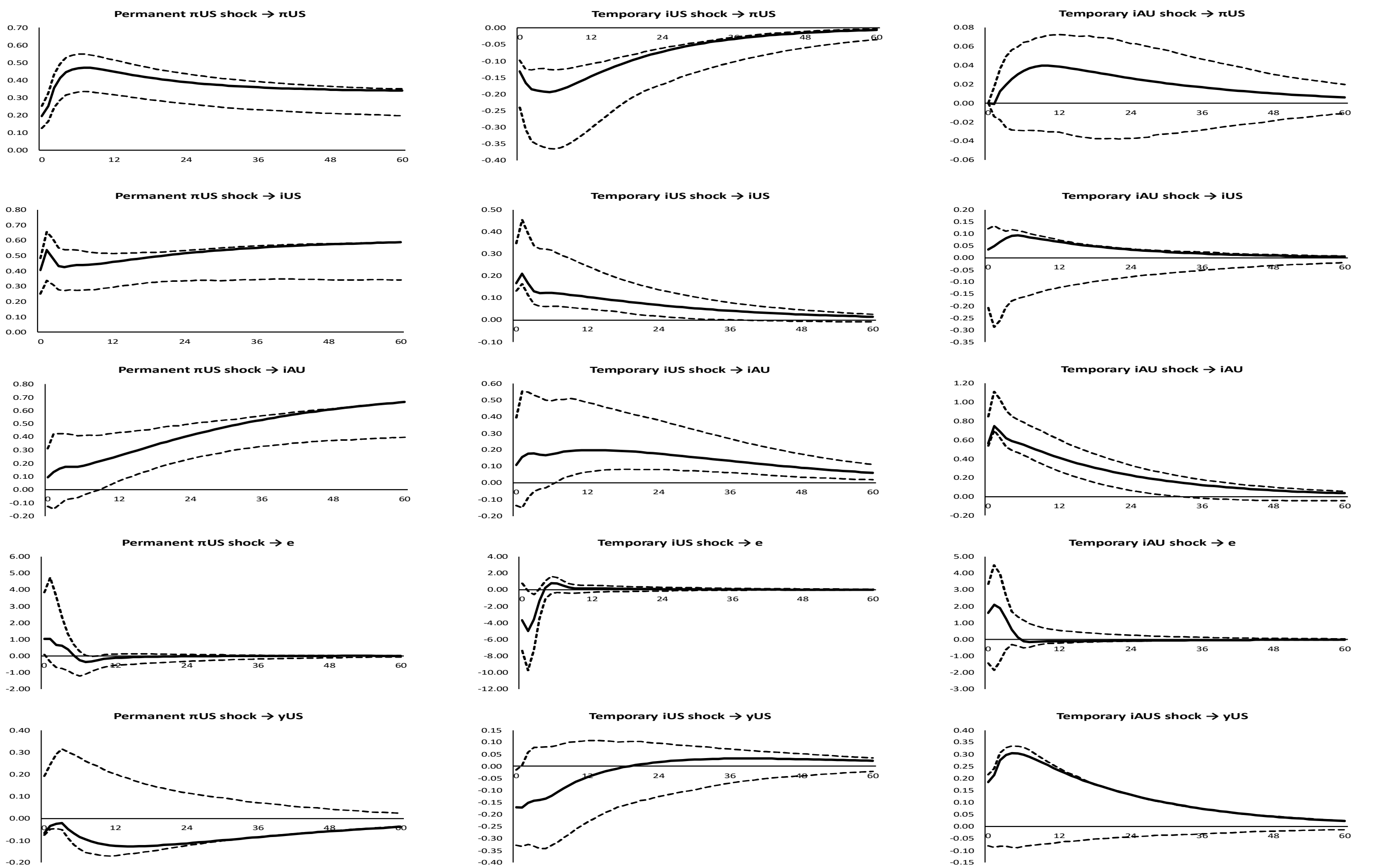

All responses are in percentage points. The first column has the responses to the permanent U.S. monetary shock, the second column the response to the temporary monetary shock and the third column the response to DE temporary monetary shock. Shocks are identified as described in the text. The structural model is estimated by Maximum Likelihood, using the Scoring Algorithm of Amisano and Giannini (1997) as implemented in JMulTi. Response to a one standard deviation shock. 90 % Hall Bootstrap confidence intervals. Sample: 1971M1-2019M6

Figure 1: Impulse Response Functions - U.S. - DE model



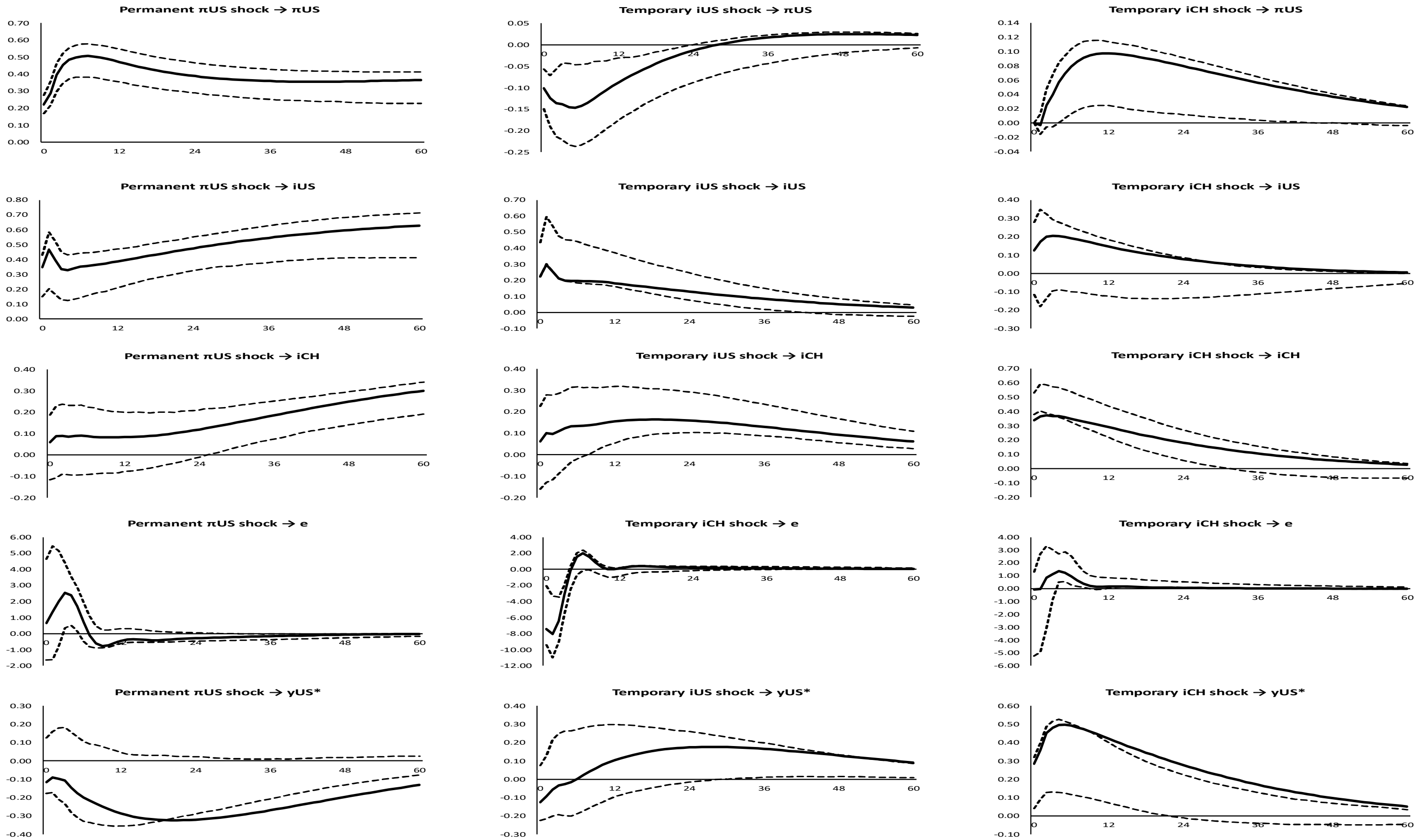

All responses are in percentage points. The first column has the responses to the permanent U.S. monetary shock, the second column the response to the temporary monetary shock and the third column the response to FR temporary monetary shock. Shocks are identified as described in the text. The structural model is estimated by Maximum Likelihood, using the Scoring Algorithm of Amisano and Giannini (1997) as implemented in JMulTi. Response to a one standard deviation shock. 90 % Hall Bootstrap confidence intervals. Sample: 1971M1-2019M6. *Some IRFs lay outside the respective confidence intervals, particularly the response of *y*, which can be attributed to imprecise point estimates. This is bound to occur with bootstrapped confidence intervals, but we highlight that our focus is on the impact of the identified shocks on the remaining variables.

Figure 2: Impulse Response Functions - U.S. - FR model




All responses are in percentage points. The first column has the responses to the permanent U.S. monetary shock, the second column the response to the temporary monetary shock and the third column the response to AU temporary monetary shock. Shocks are identified as described in the text. The structural model is estimated by Maximum Likelihood, using the Scoring Algorithm of Amisano and Giannini (1997) as implemented in JMulTi. Response to a one standard deviation shock. 90 % Hall Bootstrap confidence intervals. Sample: 1971M1-2019M6

Figure 3: Impulse Response Functions - U.S. - AU model




All responses are in percentage points. The first column has the responses to the permanent U.S. monetary shock, the second column the response to the temporary monetary shock and the third column the response to CH temporary monetary shock. Shocks are identified as described in the text. The structural model is estimated by Maximum Likelihood, using the Scoring Algorithm of Amisano and Giannini (1997) as implemented in JMulTi. Response to a one standard deviation shock. 90 % Hall Bootstrap confidence intervals. Sample: 1974M1-2019M6. *Some IRFs lay outside the respective confidence intervals, particularly the response of *y*, which can be attributed to imprecise point estimates. This is bound to occur with bootstrapped confidence intervals, but we highlight that our focus is on the impact of the identified shocks on the remaining variables.

Figure 4: Impulse Response Functions - U.S. - CH model

As in the case of the model with G.B. in the main text, the figures show that the impacts of the permanent and temporary U.S. monetary policy shocks on inflation and nominal rates are, qualitatively, exactly those documented in Valle e Azevedo *et al.* (2019) and Uribe (2021): the permanent shock, associated with a permanent rise in U.S. nominal rates, leads to a permanent increase in inflation, even in the short run, while the temporary shock, associated with a temporary increase in nominal interest rates, has a negative impact on inflation. All the adjustments essentially take place within 3 years. The permanent U.S. monetary policy shock also has a permanent effect on the nominal interest rate of the second economy, as expected from the cointegration relation between the two nominal interest rates. Also, the temporary U.S. monetary policy shock also appears to have a positive impact on the nominal interest rate of the secondary economy.

Focusing on the temporary monetary policy shock of the second economy, it is found that the impacts on the U.S. interest rate are positive but not statistically significant at the conventional levels. In the models with DE, FR and CH the temporary monetary policy shock in the second economy appears to have a positive impact on U.S. inflation. From the penultimate row in each Figure, displaying the impacts of the structural shocks on the exchange rate, one easily concludes that the temporary monetary policy shocks (both in the U.S. and in the second economy) lead in most models to an appreciation of the domestic currency. Here, the only exception appears in the model with AU data, where the temporary monetary policy shock of this economy has no significant impact on the rate of change of the exchange rate. When we move to the permanent U.S. monetary policy shock, the impact is the opposite, i.e., a permanent increase in the U.S. nominal interest rate leads to a depreciation of the USD against the five currencies considered in our sample. This result confirms the findings of Schmitt-Grohé and Uribe (2021) who document that a permanent monetary policy shock leads to a depreciation of the domestic currency using data for the U.S., G.B. and JP. Comparing the impacts of the temporary and permanent U.S. monetary policy shocks, it is notorious that the temporary shock has a much stronger and immediate impact on the exchange rate than the permanent one. In a sense this outcome could be expected given the strong consensus in the literature pointing to currency appreciations (i.e. when a distinction is not made between temporary and permanent monetary policy shocks). At the same time, the impact of the permanent U.S. monetary policy shock appears to be short-lived and not immediate, with the exception of the model for AU. Across models, the USD exchange rate variation followed by the permanent U.S. monetary policy shock is only significant for some months and and often not on impact, despite the increase in nominal interest rates being immediate.

In order to better understand the importance of the identified structural shocks, in Table 3, 4, 5 and 6 we look at the forecast error variance decomposition (FEVD) of the five variables in the models. Based on these results, some comments stand out: First, the shock that explains the majority of the forecast error variance of $e_t$ is the shock on this same variable and this fact is common across models for different countries. Second, the share decreases with the forecast horizon, but after

60 months it continues to explain more than 50% for all models. Third, the second most significant variable is the temporary U.S. monetary policy shock, explaining between 13% and 30% after 60 months. Fourth, the permanent U.S. monetary policy shock, although explaining the bulk of the forecast error variance of the inflation rate in the U.S. and of both interest rates, particularly at longer horizons, as previously documented in Valle e Azevedo *et al.* (2019), appears to explain only a small fraction for the rate of change of the exchange rate, hinting that the behaviour of the USD is much more driven by the temporary U.S. monetary policy shock than the permanent one. Again, these conclusions are not in line with the findings of Schmitt-Grohé and Uribe (2021) who conclude that the permanent monetary policy shocks account for the majority of the forecast error variance of the exchange rate. A possible reason regarding these different outcomes can be related to the assumption of non-stationary of the rate of change of the exchange rate in Schmitt-Grohé and Uribe (2021), together with the imposed cointegration between this variable and the permanent monetary policy shocks. We take this variable to be stationary.

| Var. | Shock | Horizon | | | | | | |
|---|---|---|---|---|---|---|---|---|
| | | 1 | 2 | 4 | 12 | 24 | 48 | 60 |
| | $\varepsilon_e$ | 0.91 | 0.89 | 0.84 | 0.80 | 0.79 | 0.79 | 0.79 |
| | $\varepsilon_{\pi_{US}}$ | 0.00 | 0.00 | 0.01 | 0.03 | 0.03 | 0.03 | 0.03 |
| $e - US/DE$ | $\varepsilon_{i_{DE}}$ | 0.00 | 0.00 | 0.00 | 0.01 | 0.01 | 0.01 | 0.01 |
| | $\varepsilon_{i_{US}}$ | 0.08 | 0.10 | 0.12 | 0.12 | 0.13 | 0.13 | 0.13 |
| | $\varepsilon_{y_{US}}$ | 0.00 | 0.01 | 0.02 | 0.03 | 0.04 | 0.04 | 0.04 |
| | $\varepsilon_e$ | 0.03 | 0.04 | 0.03 | 0.01 | 0.01 | 0.00 | 0.00 |
| | $\varepsilon_{\pi_{US}}$ | 0.48 | 0.48 | 0.56 | 0.65 | 0.72 | 0.79 | 0.82 |
| $\pi_{US}$ | $\varepsilon_{i_{DE}}$ | 0.00 | 0.00 | 0.00 | 0.00 | 0.01 | 0.02 | 0.02 |
| | $\varepsilon_{i_{US}}$ | 0.31 | 0.31 | 0.25 | 0.16 | 0.11 | 0.07 | 0.06 |
| | $\varepsilon_{y_{US}}$ | 0.17 | 0.17 | 0.17 | 0.17 | 0.16 | 0.12 | 0.10 |
| | $\varepsilon_e$ | 0.00 | 0.00 | 0.00 | 0.01 | 0.01 | 0.01 | 0.01 |
| | $\varepsilon_{\pi_{US}}$ | 0.10 | 0.14 | 0.19 | 0.26 | 0.32 | 0.45 | 0.52 |
| $i_{DE}$ | $\varepsilon_{i_{DE}}$ | 0.29 | 0.27 | 0.26 | 0.27 | 0.25 | 0.19 | 0.16 |
| | $\varepsilon_{i_{US}}$ | 0.06 | 0.06 | 0.06 | 0.09 | 0.14 | 0.15 | 0.14 |
| | $\varepsilon_{y_{US}}$ | 0.54 | 0.54 | 0.49 | 0.38 | 0.28 | 0.19 | 0.17 |
| | $\varepsilon_e$ | 0.00 | 0.00 | 0.00 | 0.01 | 0.01 | 0.00 | 0.00 |
| | $\varepsilon_{\pi_{US}}$ | 0.80 | 0.80 | 0.79 | 0.79 | 0.84 | 0.90 | 0.92 |
| $i_{US}$ | $\varepsilon_{i_{DE}}$ | 0.00 | 0.00 | 0.02 | 0.04 | 0.03 | 0.02 | 0.01 |
| | $\varepsilon_{i_{US}}$ | 0.16 | 0.16 | 0.16 | 0.15 | 0.12 | 0.07 | 0.06 |
| | $\varepsilon_{y_{US}}$ | 0.05 | 0.03 | 0.02 | 0.01 | 0.01 | 0.01 | 0.01 |
| | $\varepsilon_e$ | 0.00 | 0.01 | 0.00 | 0.01 | 0.01 | 0.00 | 0.00 |
| | $\varepsilon_{\pi_{US}}$ | 0.01 | 0.01 | 0.00 | 0.00 | 0.01 | 0.01 | 0.01 |
| $y_{US}$ | $\varepsilon_{i_{DE}}$ | 0.54 | 0.55 | 0.55 | 0.41 | 0.23 | 0.09 | 0.07 |
| | $\varepsilon_{i_{US}}$ | 0.01 | 0.01 | 0.01 | 0.04 | 0.06 | 0.05 | 0.04 |
| | $\varepsilon_{y_{US}}$ | 0.44 | 0.43 | 0.44 | 0.54 | 0.70 | 0.85 | 0.89 |

Table 3. Forecast error variance decomposition - U.S. - DE model

| Var. | Shock | Horizon | | | | | | |
|---|---|---|---|---|---|---|---|---|
| | | 1 | 2 | 4 | 12 | 24 | 48 | 60 |
| $e - US/FR$ | $\varepsilon_e$ | 0.88 | 0.86 | 0.84 | 0.80 | 0.78 | 0.77 | 0.77 |
| | $\varepsilon_{\pi_{US}}$ | 0.00 | 0.00 | 0.00 | 0.03 | 0.03 | 0.03 | 0.03 |
| | $\varepsilon_{i_{FR}}$ | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 |
| | $\varepsilon_{i_{US}}$ | 0.12 | 0.13 | 0.15 | 0.16 | 0.16 | 0.16 | 0.16 |
| | $\varepsilon_{y_{US}}$ | 0.00 | 0.00 | 0.00 | 0.01 | 0.02 | 0.03 | 0.03 |
| $\pi_{US}$ | $\varepsilon_e$ | 0.04 | 0.03 | 0.02 | 0.01 | 0.01 | 0.00 | 0.00 |
| | $\varepsilon_{\pi_{US}}$ | 0.59 | 0.58 | 0.61 | 0.65 | 0.76 | 0.86 | 0.89 |
| | $\varepsilon_{i_{FR}}$ | 0.00 | 0.00 | 0.03 | 0.06 | 0.05 | 0.03 | 0.03 |
| | $\varepsilon_{i_{US}}$ | 0.25 | 0.22 | 0.15 | 0.08 | 0.05 | 0.03 | 0.02 |
| | $\varepsilon_{y_{US}}$ | 0.12 | 0.16 | 0.19 | 0.19 | 0.13 | 0.08 | 0.06 |
| $i_{FR}$ | $\varepsilon_e$ | 0.00 | 0.00 | 0.00 | 0.00 | 0.01 | 0.01 | 0.01 |
| | $\varepsilon_{\pi_{US}}$ | 0.02 | 0.04 | 0.07 | 0.12 | 0.26 | 0.60 | 0.71 |
| | $\varepsilon_{i_{FR}}$ | 0.65 | 0.63 | 0.64 | 0.61 | 0.41 | 0.16 | 0.11 |
| | $\varepsilon_{i_{US}}$ | 0.00 | 0.00 | 0.01 | 0.11 | 0.21 | 0.15 | 0.11 |
| | $\varepsilon_{y_{US}}$ | 0.33 | 0.33 | 0.28 | 0.16 | 0.12 | 0.08 | 0.06 |
| $i_{US}$ | $\varepsilon_e$ | 0.00 | 0.00 | 0.00 | 0.02 | 0.02 | 0.01 | 0.01 |
| | $\varepsilon_{\pi_{US}}$ | 0.67 | 0.67 | 0.67 | 0.57 | 0.64 | 0.81 | 0.86 |
| | $\varepsilon_{i_{FR}}$ | 0.03 | 0.04 | 0.08 | 0.07 | 0.04 | 0.02 | 0.01 |
| | $\varepsilon_{i_{US}}$ | 0.27 | 0.27 | 0.24 | 0.28 | 0.23 | 0.11 | 0.08 |
| | $\varepsilon_{y_{US}}$ | 0.02 | 0.01 | 0.01 | 0.06 | 0.09 | 0.05 | 0.04 |
| $y_{US}$ | $\varepsilon_e$ | 0.01 | 0.01 | 0.01 | 0.00 | 0.00 | 0.00 | 0.00 |
| | $\varepsilon_{\pi_{US}}$ | 0.09 | 0.08 | 0.07 | 0.12 | 0.11 | 0.07 | 0.06 |
| | $\varepsilon_{i_{FR}}$ | 0.16 | 0.16 | 0.17 | 0.08 | 0.03 | 0.01 | 0.01 |
| | $\varepsilon_{i_{US}}$ | 0.02 | 0.01 | 0.00 | 0.00 | 0.01 | 0.01 | 0.01 |
| | $\varepsilon_{y_{US}}$ | 0.72 | 0.73 | 0.75 | 0.80 | 0.85 | 0.91 | 0.92 |

Table 4. Forecast error variance decomposition - U.S. - FR model

| Var. | Shock | Horizon | | | | | | |
|---|---|---|---|---|---|---|---|---|
| | | 1 | 2 | 4 | 12 | 24 | 48 | 60 |
| | $\varepsilon_e$ | 0.85 | 0.83 | 0.81 | 0.79 | 0.79 | 0.79 | 0.79 |
| | $\varepsilon_{\pi_{US}}$ | 0.01 | 0.01 | 0.01 | 0.01 | 0.01 | 0.01 | 0.01 |
| $e - US/AU$ | $\varepsilon_{i_{AU}}$ | 0.02 | 0.02 | 0.03 | 0.03 | 0.03 | 0.03 | 0.03 |
| | $\varepsilon_{i_{US}}$ | 0.11 | 0.13 | 0.14 | 0.14 | 0.14 | 0.14 | 0.14 |
| | $\varepsilon_{y_{US}}$ | 0.01 | 0.01 | 0.02 | 0.02 | 0.03 | 0.03 | 0.03 |
| | $\varepsilon_e$ | 0.03 | 0.02 | 0.02 | 0.01 | 0.01 | 0.01 | 0.00 |
| | $\varepsilon_{\pi_{US}}$ | 0.45 | 0.45 | 0.52 | 0.62 | 0.70 | 0.79 | 0.81 |
| $\pi_{US}$ | $\varepsilon_{i_{AU}}$ | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 |
| | $\varepsilon_{i_{US}}$ | 0.20 | 0.20 | 0.15 | 0.11 | 0.09 | 0.06 | 0.05 |
| | $\varepsilon_{y_{US}}$ | 0.32 | 0.33 | 0.32 | 0.26 | 0.21 | 0.14 | 0.13 |
| | $\varepsilon_e$ | 0.00 | 0.00 | 0.00 | 0.02 | 0.03 | 0.02 | 0.01 |
| | $\varepsilon_{\pi_{US}}$ | 0.02 | 0.02 | 0.04 | 0.08 | 0.21 | 0.51 | 0.62 |
| $i_{AU}$ | $\varepsilon_{i_{AU}}$ | 0.73 | 0.74 | 0.75 | 0.74 | 0.61 | 0.34 | 0.26 |
| | $\varepsilon_{i_{US}}$ | 0.03 | 0.03 | 0.04 | 0.07 | 0.10 | 0.08 | 0.06 |
| | $\varepsilon_{y_{US}}$ | 0.22 | 0.20 | 0.16 | 0.08 | 0.06 | 0.05 | 0.04 |
| | $\varepsilon_e$ | 0.00 | 0.00 | 0.00 | 0.01 | 0.00 | 0.00 | 0.00 |
| | $\varepsilon_{\pi_{US}}$ | 0.83 | 0.85 | 0.86 | 0.85 | 0.89 | 0.94 | 0.95 |
| $i_{US}$ | $\varepsilon_{i_{AU}}$ | 0.01 | 0.01 | 0.01 | 0.02 | 0.02 | 0.01 | 0.01 |
| | $\varepsilon_{i_{US}}$ | 0.14 | 0.14 | 0.11 | 0.08 | 0.05 | 0.03 | 0.02 |
| | $\varepsilon_{y_{US}}$ | 0.02 | 0.01 | 0.01 | 0.04 | 0.03 | 0.02 | 0.02 |
| | $\varepsilon_e$ | 0.11 | 0.09 | 0.05 | 0.01 | 0.00 | 0.00 | 0.00 |
| | $\varepsilon_{\pi_{US}}$ | 0.02 | 0.01 | 0.00 | 0.01 | 0.01 | 0.01 | 0.01 |
| $y_{US}$ | $\varepsilon_{i_{AU}}$ | 0.13 | 0.14 | 0.16 | 0.11 | 0.06 | 0.03 | 0.02 |
| | $\varepsilon_{i_{US}}$ | 0.12 | 0.10 | 0.07 | 0.02 | 0.01 | 0.00 | 0.00 |
| | $\varepsilon_{y_{US}}$ | 0.62 | 0.65 | 0.71 | 0.85 | 0.92 | 0.96 | 0.97 |

Table 5. Forecast error variance decomposition - U.S. - AU model

| Var. | Shock | Horizon | | | | | | |
|---|---|---|---|---|---|---|---|---|
| | | 1 | 2 | 4 | 12 | 24 | 48 | 60 |
| $e-US/CH$ | $\varepsilon_e$ | 0.69 | 0.69 | 0.64 | 0.62 | 0.62 | 0.61 | 0.61 |
| | $\varepsilon_{\pi_{US}}$ | 0.00 | 0.01 | 0.02 | 0.04 | 0.04 | 0.04 | 0.04 |
| | $\varepsilon_{i_{CH}}$ | 0.00 | 0.00 | 0.00 | 0.01 | 0.01 | 0.01 | 0.01 |
| | $\varepsilon_{i_{US}}$ | 0.28 | 0.29 | 0.32 | 0.30 | 0.30 | 0.30 | 0.30 |
| | $\varepsilon_{y_{US}}$ | 0.02 | 0.01 | 0.02 | 0.03 | 0.03 | 0.03 | 0.03 |
| $\pi_{US}$ | $\varepsilon_e$ | 0.06 | 0.05 | 0.03 | 0.02 | 0.01 | 0.01 | 0.01 |
| | $\varepsilon_{\pi_{US}}$ | 0.55 | 0.55 | 0.60 | 0.68 | 0.74 | 0.81 | 0.84 |
| | $\varepsilon_{i_{CH}}$ | 0.00 | 0.00 | 0.00 | 0.02 | 0.03 | 0.03 | 0.02 |
| | $\varepsilon_{i_{US}}$ | 0.11 | 0.11 | 0.08 | 0.05 | 0.04 | 0.02 | 0.02 |
| | $\varepsilon_{y_{US}}$ | 0.28 | 0.29 | 0.28 | 0.23 | 0.19 | 0.13 | 0.11 |
| $i_{CH}$ | $\varepsilon_e$ | 0.00 | 0.00 | 0.01 | 0.01 | 0.01 | 0.01 | 0.01 |
| | $\varepsilon_{\pi_{US}}$ | 0.02 | 0.03 | 0.03 | 0.04 | 0.06 | 0.20 | 0.32 |
| | $\varepsilon_{i_{CH}}$ | 0.58 | 0.56 | 0.59 | 0.65 | 0.64 | 0.48 | 0.40 |
| | $\varepsilon_{i_{US}}$ | 0.02 | 0.03 | 0.04 | 0.09 | 0.15 | 0.18 | 0.16 |
| | $\varepsilon_{y_{US}}$ | 0.38 | 0.38 | 0.33 | 0.21 | 0.14 | 0.12 | 0.11 |
| $i_{US}$ | $\varepsilon_e$ | 0.00 | 0.00 | 0.00 | 0.01 | 0.00 | 0.00 | 0.00 |
| | $\varepsilon_{\pi_{US}}$ | 0.62 | 0.64 | 0.61 | 0.61 | 0.70 | 0.84 | 0.88 |
| | $\varepsilon_{i_{CH}}$ | 0.08 | 0.09 | 0.13 | 0.14 | 0.10 | 0.05 | 0.03 |
| | $\varepsilon_{i_{US}}$ | 0.26 | 0.26 | 0.25 | 0.21 | 0.15 | 0.08 | 0.06 |
| | $\varepsilon_{y_{US}}$ | 0.03 | 0.02 | 0.02 | 0.04 | 0.05 | 0.03 | 0.03 |
| $y_{US}$ | $\varepsilon_e$ | 0.10 | 0.10 | 0.08 | 0.03 | 0.01 | 0.00 | 0.00 |
| | $\varepsilon_{\pi_{US}}$ | 0.05 | 0.03 | 0.02 | 0.05 | 0.06 | 0.05 | 0.04 |
| | $\varepsilon_{i_{CH}}$ | 0.28 | 0.33 | 0.37 | 0.26 | 0.15 | 0.07 | 0.05 |
| | $\varepsilon_{i_{US}}$ | 0.05 | 0.04 | 0.02 | 0.01 | 0.01 | 0.01 | 0.01 |
| | $\varepsilon_{y_{US}}$ | 0.52 | 0.50 | 0.52 | 0.66 | 0.77 | 0.86 | 0.89 |

Table 6. Forecast error variance decomposition - U.S. - CH model

## 3. Pre-inflation targeting

With the purpose of providing some robustness to the results reported in Section 3 of the main text, next we analyse results for a pre-inflation targeting sample. We re-estimate the VEC model for the sample up to the period when the second economy in each models adopts an explicit inflation target or 1999 for Germany and France (introduction of the euro). The new sample periods considered are as follows: G.B. from 1971M4 to 1992M1; DE and FR from 1971M1 to 1999M1; AU from 1971M4 to 1996M9 and CH from 1974M1 to 1999M12.

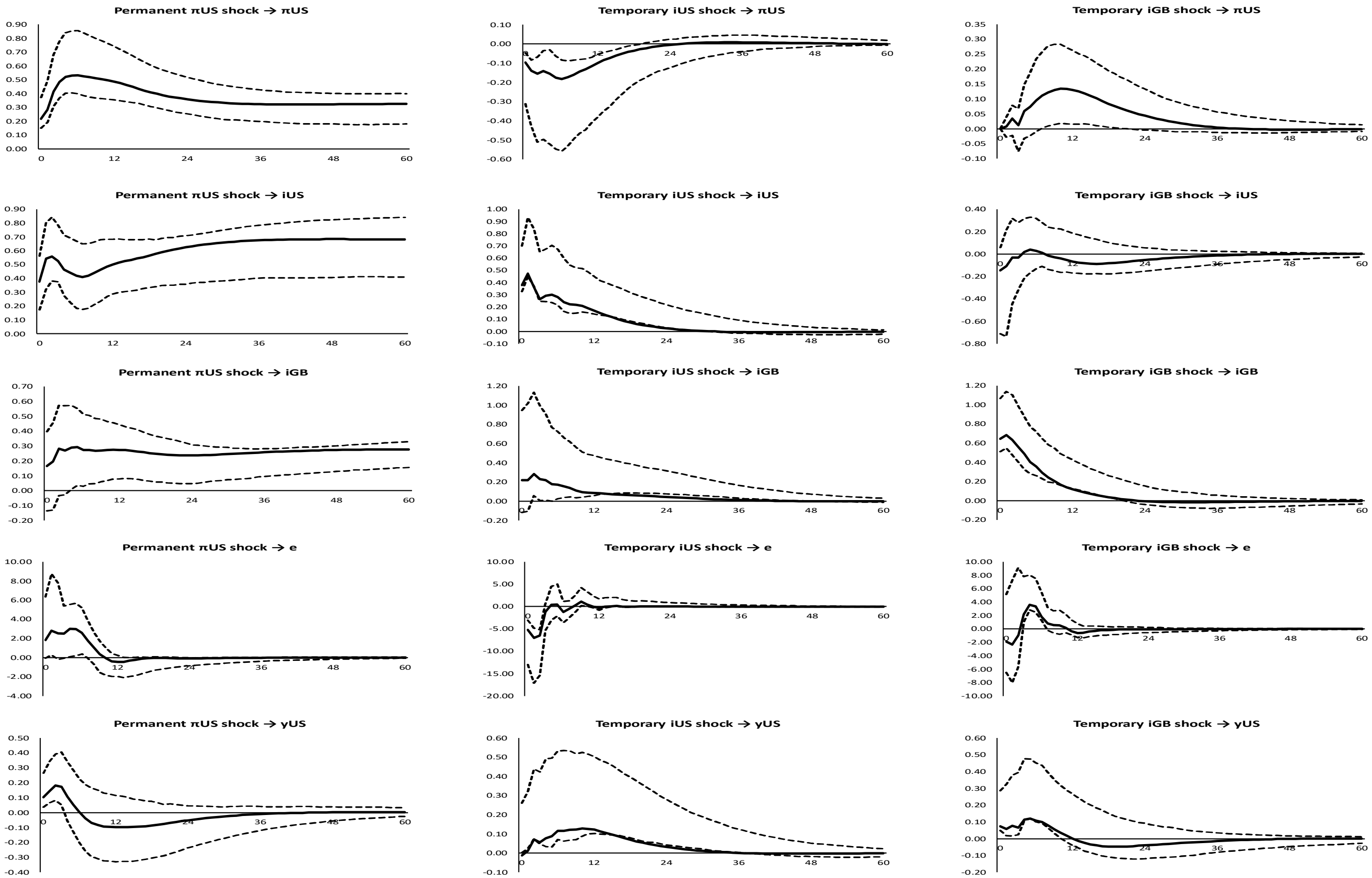
The estimation results of the cointegration relations and the IRFs are presented in table 7 and Figures 5, 6, 7, 8 and 9, respectively. In general, the results are quite similar to those from the baseline model considering the full sample. However the are differences that are important to highlight. The cointegration parameter for the Fisher relation in the US continues to be below unity across all models, with the adjustment to this relation occurring from both the inflation and nominal interest rates. These parameters are now smaller than the ones with the full sample, ranging from $0.33$ to $0.48$, which may be justified by the greater difficulty in identifying a Fisher relation in view of the low relative variability of inflation and interest rates in this sub-sample, compared with the full sample. For the cointegration of monetary policies the parameters continue to show some dispersion and are once again smaller (from $0.41$ to $0.98$) and the adjustment to the long-run relation is exclusively made by the nominal interest rate of the second economy. It's important to underline that in the models with DE and CH this coefficient is not statistically significant, which can be due to the (almost) stationary behaviour of interest rates even in this sub-sample. We recall that in DE and CH inflation (and interest rates) remained relatively low throughout the sample, in contrast with the other economies, that have witnessed the 1970's great inflation and the subsequent disinflation. Despite these complications, and overall, the conclusions from the impulse response analysis also broadly support the results of the model estimated over the whole sample. The temporary and permanent U.S. monetary policy shocks have the impact on inflation and interest rates in the U.S. that was documented for the whole sample. The impacts on the USD exchange rate are also mostly consistent with the original model: the temporary U.S. monetary policy shock leads to an appreciation of the dollar against the other currencies, while the permanent U.S. shock results in a USD depreciation. Nevertheless, in the model with CH the temporary U.S. monetary policy shock appears to have no significant impact on the exchange rate. In the other models that impact is sizeable compared with the impact of the permanent monetary policy shock. In the model with DE the permanent monetary policy shock has a positive but insignificant impact on the exchange rate, again in contrast to the significance found in the full sample results.

| Fisher relation U.S. | | | | | | |
|---|---|---|---|---|---|---|
| | $\beta^F$ | $\gamma_e$ | $\gamma_{\pi^{US}}$ | $\gamma_{i^*}$ | $\gamma_{i^{US}}$ | $\gamma_{y^{US}}$ |
| U.S. - G.B. | $-0.48^{***}$ | $0.18$ | $-0.03^{**}$ | $0.02$ | $0.05^{**}$ | $-0.01$ |
| (6 lags) | 0.18 | 0.41 | 0.01 | 0.03 | 0.02 | 0.01 |
| U.S. - DE | $-0.36^{***}$ | $-0.64$ | $-0.02^{**}$ | $-0.01$ | $0.05^{***}$ | $-0.03^{***}$ |
| (2 lags) | 0.14 | 0.46 | 0.01 | 0.01 | 0.02 | 0.01 |
| U.S. - FR | $-0.33^{**}$ | $-0.45$ | $-0.02^{*}$ | $-0.04^{*}$ | $0.04^{**}$ | $-0.02$ |
| (4 lags) | 0.17 | 0.40 | 0.01 | 0.02 | 0.02 | 0.01 |
| U.S. - AU | $-0.40^{**}$ | $-0.33$ | $-0.02^{**}$ | $-0.08^{**}$ | $0.05^{***}$ | $-0.02$ |
| (2 lags) | 0.16 | 0.34 | 0.01 | 0.03 | 0.02 | 0.01 |
| U.S. - CH | $-0.36^{***}$ | $-0.42$ | $-0.02^{**}$ | $-0.01$ | $0.05^{***}$ | $-0.04^{***}$ |
| (2 lags) | 0.12 | 0.57 | 0.01 | 0.02 | 0.02 | 0.01 |

| Monetary policy Cointegration | | | | | | |
|---|---|---|---|---|---|---|
| | $\beta^{MP}$ | $\gamma_e$ | $\gamma_{\pi^{US}}$ | $\gamma_{i^*}$ | $\gamma_{i^{US}}$ | $\gamma_{y^{US}}$ |
| U.S. - G.B. | $-0.41^{***}$ | $0.43$ | $0.01$ | $-0.12^{***}$ | $-0.01$ | $-0.01$ |
| (6 lags) | 0.17 | 0.41 | 0.01 | 0.03 | 0.02 | 0.01 |
| U.S. - DE | $0.10$ | $-0.10$ | $0.01$ | $-0.02^{**}$ | $0.00$ | $-0.02^{**}$ |
| (2 lags) | 0.11 | 0.32 | 0.01 | 0.01 | 0.01 | 0.01 |
| U.S. - FR | $-0.64^{***}$ | $0.14$ | $0.01$ | $-0.05^{***}$ | $-0.01$ | $0.00$ |
| (4 lags) | 0.18 | 0.32 | 0.01 | 0.02 | 0.01 | 0.01 |
| U.S. - AU | $-0.98^{***}$ | $-0.17$ | $0.00$ | $-0.07^{***}$ | $0.00$ | $-0.01$ |
| (2 lags) | 0.36 | 0.18 | 0.01 | 0.02 | 0.01 | 0.01 |
| U.S. - CH | $0.12$ | $0.31$ | $0.00$ | $-0.03^{***}$ | $-0.01$ | $-0.03^{***}$ |
| (2 lags) | 0.24 | 0.36 | 0.01 | 0.01 | 0.01 | 0.01 |

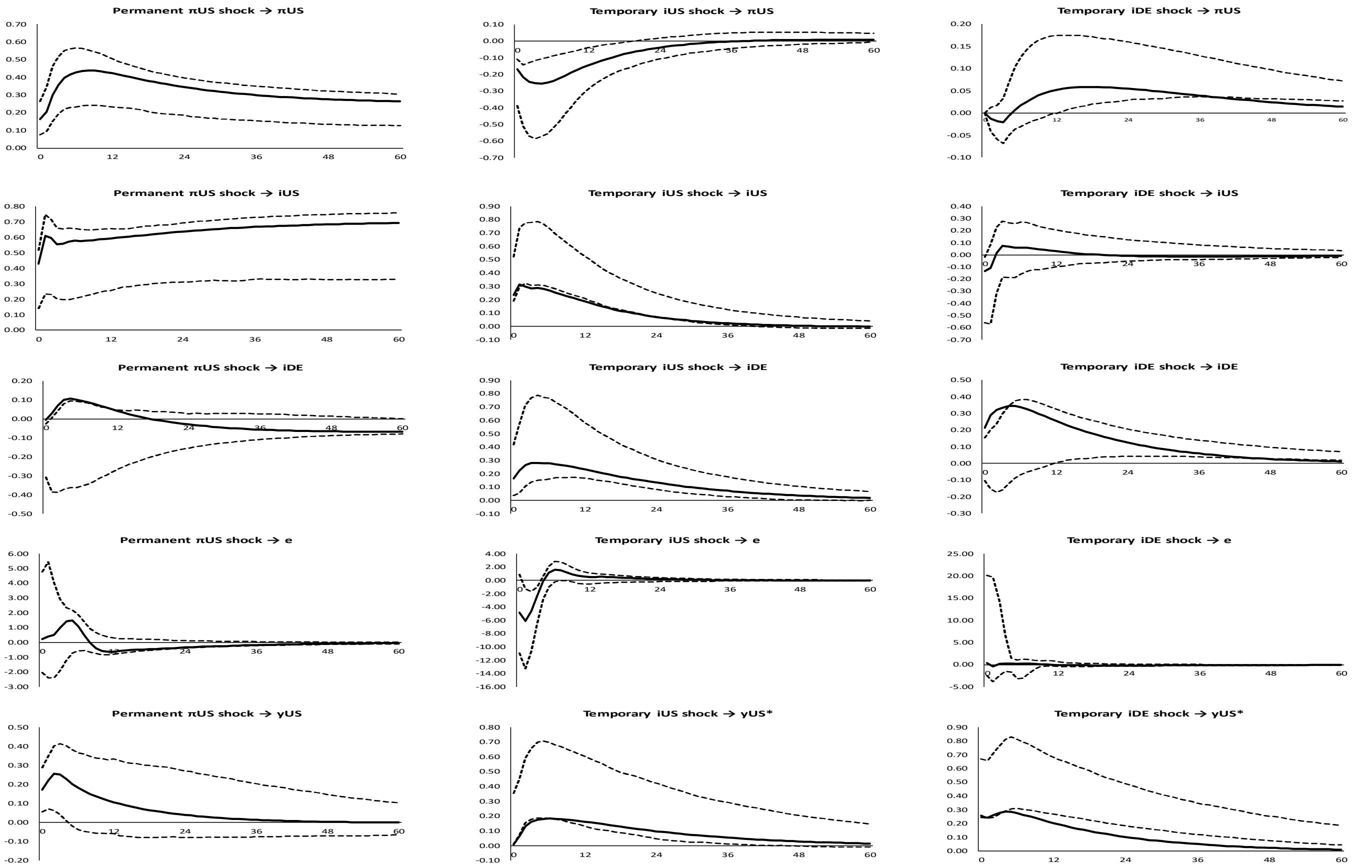
Notes: $\gamma_{var}$ corresponds to the adjustment parameter in the equation for variable "$var$". Superscript "US" was added to variables $\pi$, $i$ and $y$ variables to make clear they are U.S. variables. Standard errors displayed below the coefficient estimates. Simple two step estimator (S2S) employed as implemented in JMulit. $***$ and $**$ denote statistical significance at 1% and 5% levels, respectively. U.S. - G.B. model: 1971M4-1992M1. U.S. - DE model: 1971M1-1999M1. U.S. - FR model: 1971M1 - 1999M1. U.S. - AU model: 1971M4 - 1996M9. U.S. - CH model: 1974M1 - 1999M12. AIC for lag choice.

Table 7. Estimation of Vector Error Correction Model and Cointegration Parameters - Pre-inflation targeting



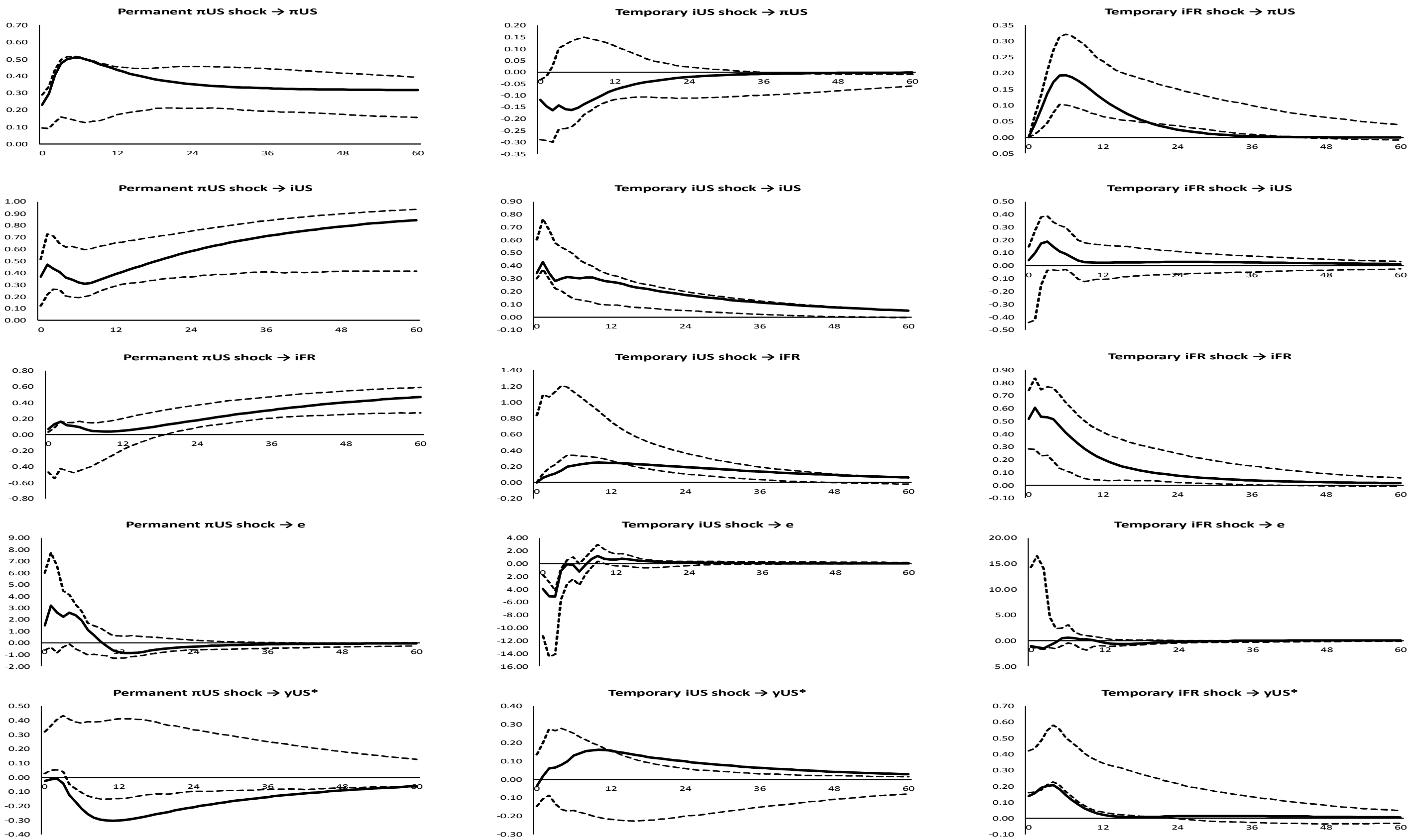

All responses are in percentage points. The first column has the responses to the permanent U.S. monetary shock, the second column the response to the temporary monetary shock and the third column the response to G.B. temporary monetary shock. Sample ranges from 1971M4 to 1992M1. Shocks are identified as described in the text. The structural model is estimated by Maximum Likelihood, using the Scoring Algorithm of Amisano and Giannini (1997) as implemented in JMulTi. Response to a one standard deviation shock. 90 % Hall Bootstrap confidence intervals.

Figure 5: Impulse Response Functions - U.S. - G.B. model - Pre-inflation targeting



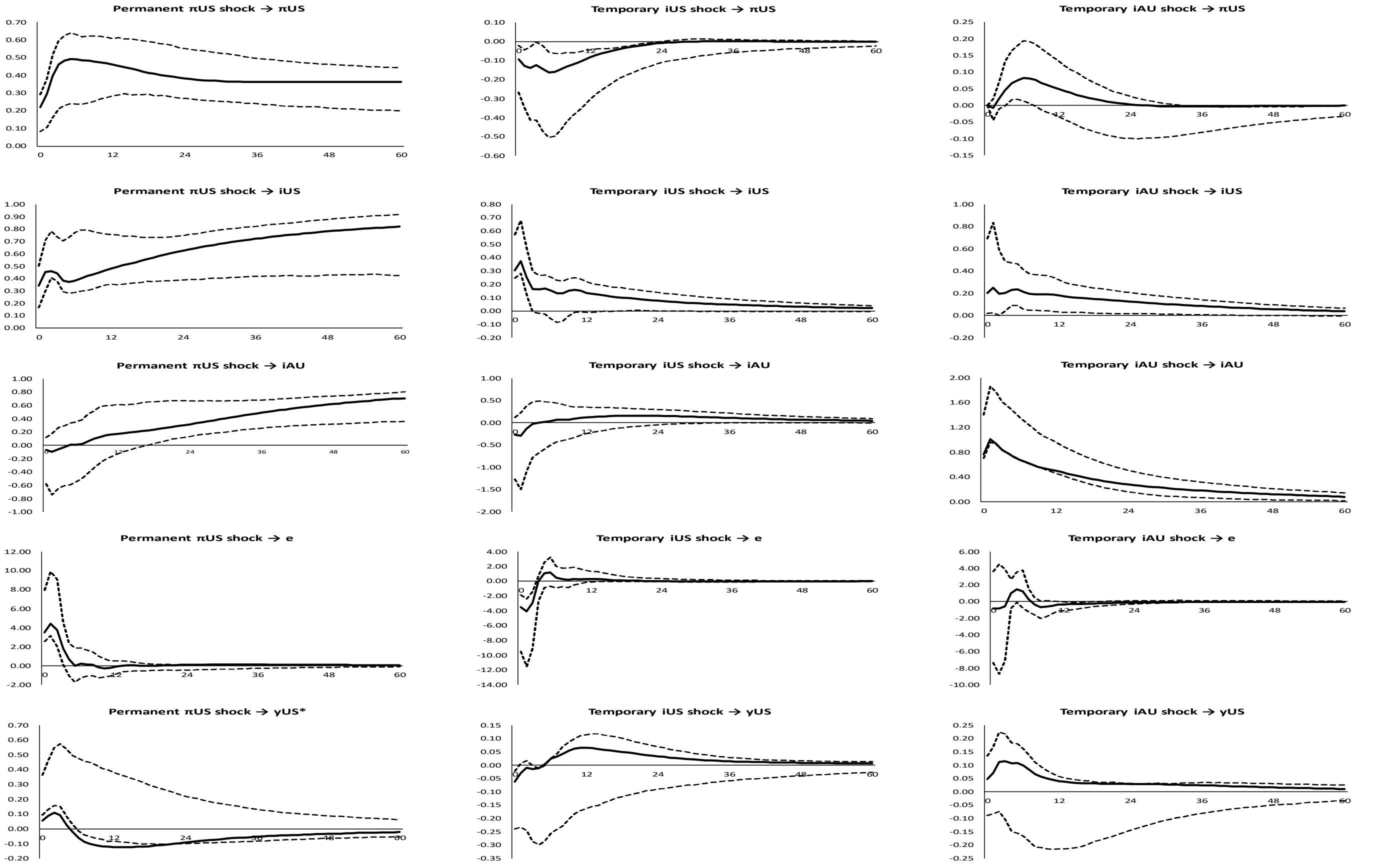

All responses are in percentage points. The first column has the responses to the permanent U.S. monetary shock, the second column the response to the temporary monetary shock and the third column the response to DE temporary monetary shock. Sample ranges from 1971M1 to 1999M1. Shocks are identified as described in the text. The structural model is estimated by Maximum Likelihood, using the Scoring Algorithm of Amisano and Giannini (1997) as implemented in JMulTi. Response to a one standard deviation shock. 90 % Hall Bootstrap confidence intervals. *Some IRFs lay outside the respective confidence intervals, particularly the response on *y*, which can be attributed to imprecise point estimates. This is bound to occur with bootstrapped confidence intervals, but we highlight that our focus is on the impact of the identified shocks on the remaining variables.

Figure 6: Impulse Response Functions - U.S. - DE model - Pre-inflation targeting



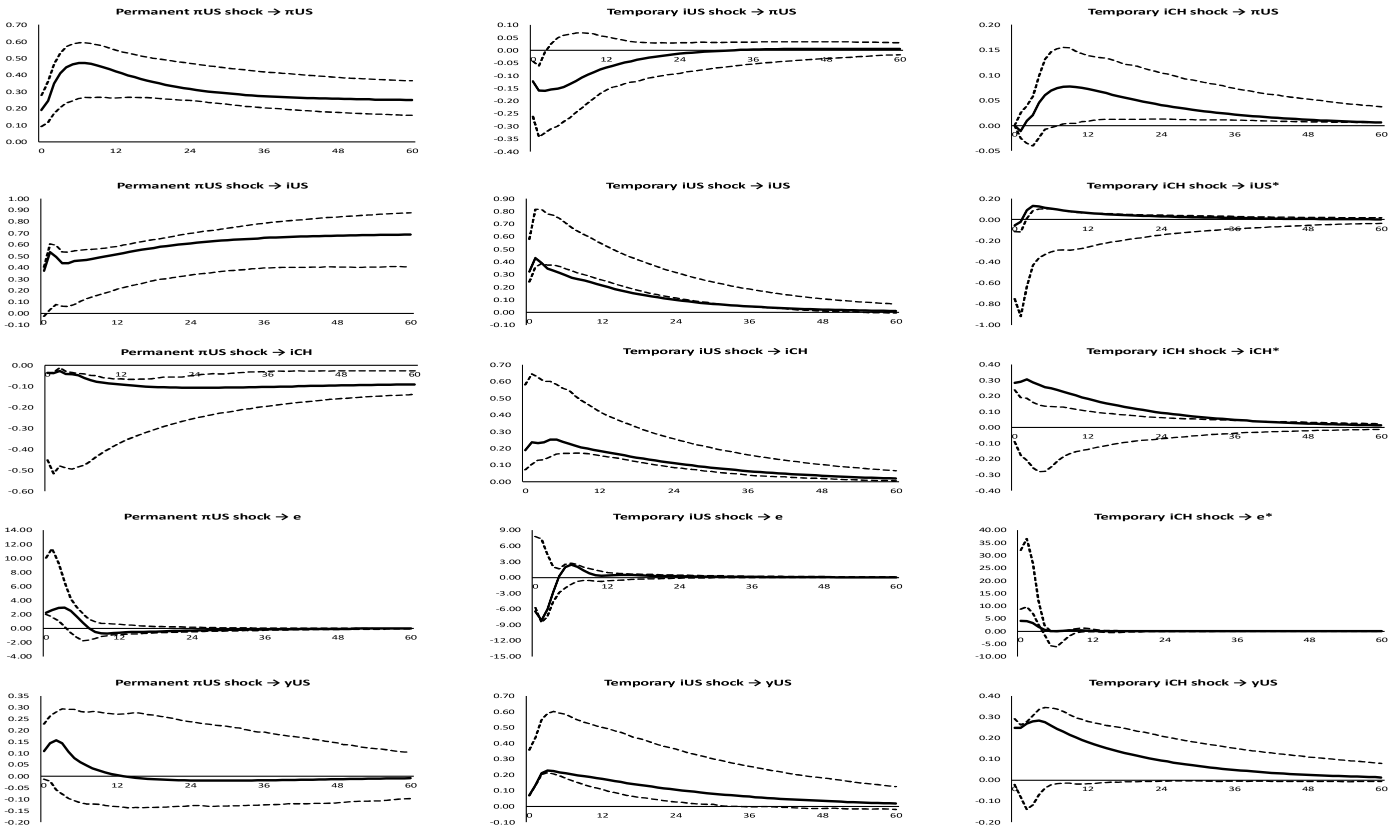

All responses are in percentage points. The first column has the responses to the permanent U.S. monetary shock, the second column the response to the temporary monetary shock and the third column the response to FR temporary monetary shock. Sample ranges from 1971M1 to 1999M1. Shocks are identified as described in the text. The structural model is estimated by Maximum Likelihood, using the Scoring Algorithm of Amisano and Giannini (1997) as implemented in JMulTi. Response to a one standard deviation shock. 90 % Hall Bootstrap confidence intervals. *Some IRFs lay outside the respective confidence intervals, particularly the response of *y*, which can be attributed to imprecise point estimates. This is bound to occur with bootstrapped confidence intervals, but we highlight that our focus is on the impact of the identified shocks on the remaining variables.

Figure 7: Impulse Response Functions - U.S. - FR model - Pre-inflation targeting




All responses are in percentage points. The first column has the responses to the permanent U.S. monetary shock, the second column the response to the temporary monetary shock and the third column the response to AU temporary monetary shock. Sample ranges from 1971M4 to 1996M9. Shocks are identified as described in the text. The structural model is estimated by Maximum Likelihood, using the Scoring Algorithm of Amisano and Giannini (1997) as implemented in JMulTi. Response to a one standard deviation shock. 90 % Hall Bootstrap confidence intervals. *Some IRFs lay outside the respective confidence intervals, particularly the response on $y$, which can be attributed to imprecise point estimates. This is bound to occur with bootstrapped confidence intervals, but we highlight that our focus is on the impact of the identified shocks on the remaining variables.

Figure 8: Impulse Response Functions - U.S. - AU model - Pre-inflation targeting




All responses are in percentage points. The first column has the responses to the permanent U.S. monetary shock, the second column the response to the temporary monetary shock and the third column the response to CH temporary monetary shock. Sample ranges from 1974M1 to 1999M12. Shocks are identified as described in the text. The structural model is estimated by Maximum Likelihood, using the Scoring Algorithm of Amisano and Giannini (1997) as implemented in JMulTi. Response to a one standard deviation shock. 90 % Hall Bootstrap confidence intervals. *Some IRFs lay outside the confidence intervals, which can be attributed to imprecise point estimates. This is bound to occur with bootstrapped confidence intervals. It is important to recall that data for interest rate for CH was found to be close to stationary for the full sample, which raises possible misspecification issues in this case.

Figure 9: Impulse Response Functions - U.S. - CH model - Pre-inflation targeting

Next, in Tables 8, 9, 10, 11 and 12, we look at the FEVD of the five variables in the models for the pre-inflation targeting sample.The findings are broadly similar with the ones from the main model: for the model with G.B. the shock that explains the larger fraction continues to be the temporary U.S. monetary policy shock, followed by the shock on $e_t$ itself. For the other models, the shock that explains the majority of the forecast error variance of $e_t$ is the shock on this same variable. The share decreases with the forecast horizon, but after 60 months it continues to explain more than 50% in all models. The second most important shock is the temporary U.S. monetary policy shock, explaining between 17% and 19% after 60 months. A conclusion that stems from all models continues to be that the permanent U.S. monetary policy shock, although explaining the bulk of the forecast error variance of the inflation rate in the U.S. and the two interest rates, particularly at longer horizons, as previously documented in Valle e Azevedo *et al.* (2019), appears to explain only a small fraction of the rate of change of the exchange rate, hinting that the FEVD of the USD is much more driven by the temporary U.S. monetary policy shock than by the permanent one.

| Var. | Shock | Horizon | | | | | | |
|---|---|---|---|---|---|---|---|---|
| | | 1 | 2 | 4 | 12 | 24 | 48 | 60 |
| | $\varepsilon_e$ | 0.41 | 0.37 | 0.34 | 0.31 | 0.31 | 0.31 | 0.31 |
| | $\varepsilon_{\pi_{US}}$ | 0.06 | 0.06 | 0.06 | 0.10 | 0.10 | 0.10 | 0.10 |
| $e - US/GB$ | $\varepsilon_{i_{GB}}$ | 0.05 | 0.04 | 0.04 | 0.10 | 0.10 | 0.10 | 0.10 |
| | $\varepsilon_{i_{US}}$ | 0.46 | 0.49 | 0.51 | 0.44 | 0.43 | 0.43 | 0.43 |
| | $\varepsilon_{y_{US}}$ | 0.02 | 0.03 | 0.05 | 0.05 | 0.05 | 0.05 | 0.05 |
| | $\varepsilon_e$ | 0.40 | 0.40 | 0.33 | 0.22 | 0.13 | 0.07 | 0.06 |
| | $\varepsilon_{\pi_{US}}$ | 0.46 | 0.44 | 0.53 | 0.67 | 0.80 | 0.88 | 0.90 |
| $\pi_{US}$ | $\varepsilon_{i_{GB}}$ | 0.00 | 0.00 | 0.00 | 0.03 | 0.03 | 0.02 | 0.01 |
| | $\varepsilon_{i_{US}}$ | 0.08 | 0.10 | 0.07 | 0.06 | 0.03 | 0.02 | 0.01 |
| | $\varepsilon_{y_{US}}$ | 0.06 | 0.06 | 0.06 | 0.03 | 0.02 | 0.01 | 0.01 |
| | $\varepsilon_e$ | 0.00 | 0.00 | 0.01 | 0.01 | 0.02 | 0.01 | 0.01 |
| | $\varepsilon_{\pi_{US}}$ | 0.14 | 0.17 | 0.28 | 0.55 | 0.72 | 0.84 | 0.87 |
| $i_{GB}$ | $\varepsilon_{i_{GB}}$ | 0.69 | 0.63 | 0.46 | 0.20 | 0.12 | 0.06 | 0.05 |
| | $\varepsilon_{i_{US}}$ | 0.04 | 0.06 | 0.12 | 0.11 | 0.08 | 0.04 | 0.03 |
| | $\varepsilon_{y_{US}}$ | 0.14 | 0.13 | 0.13 | 0.13 | 0.08 | 0.04 | 0.03 |
| | $\varepsilon_e$ | 0.00 | 0.01 | 0.02 | 0.08 | 0.06 | 0.03 | 0.02 |
| | $\varepsilon_{\pi_{US}}$ | 0.52 | 0.60 | 0.65 | 0.72 | 0.82 | 0.91 | 0.93 |
| $i_{US}$ | $\varepsilon_{i_{GB}}$ | 0.06 | 0.04 | 0.03 | 0.02 | 0.01 | 0.01 | 0.01 |
| | $\varepsilon_{i_{US}}$ | 0.24 | 0.23 | 0.23 | 0.15 | 0.09 | 0.05 | 0.04 |
| | $\varepsilon_{y_{US}}$ | 0.17 | 0.13 | 0.07 | 0.03 | 0.02 | 0.01 | 0.01 |
| | $\varepsilon_e$ | 0.02 | 0.01 | 0.02 | 0.13 | 0.12 | 0.07 | 0.06 |
| | $\varepsilon_{\pi_{US}}$ | 0.09 | 0.11 | 0.14 | 0.03 | 0.03 | 0.02 | 0.02 |
| $y_{US}$ | $\varepsilon_{i_{GB}}$ | 0.01 | 0.01 | 0.00 | 0.01 | 0.02 | 0.02 | 0.01 |
| | $\varepsilon_{i_{US}}$ | 0.05 | 0.06 | 0.09 | 0.11 | 0.07 | 0.04 | 0.03 |
| | $\varepsilon_{y_{US}}$ | 0.83 | 0.81 | 0.75 | 0.72 | 0.76 | 0.86 | 0.88 |

Table 8. Forecast error variance decomposition - U.S. - G.B. model - Pre-inflation targeting

| Var. | Shock | Horizon | | | | | | |
|---|---|---|---|---|---|---|---|---|
| | | 1 | 2 | 4 | 12 | 24 | 48 | 60 |
| $e - US/DE$ | $\varepsilon_e$ | 0.84 | 0.80 | 0.73 | 0.71 | 0.71 | 0.71 | 0.71 |
| | $\varepsilon_{\pi_{US}}$ | 0.01 | 0.02 | 0.04 | 0.06 | 0.06 | 0.06 | 0.06 |
| | $\varepsilon_{i_{DE}}$ | 0.02 | 0.02 | 0.04 | 0.04 | 0.04 | 0.04 | 0.04 |
| | $\varepsilon_{i_{US}}$ | 0.12 | 0.15 | 0.16 | 0.16 | 0.16 | 0.17 | 0.17 |
| | $\varepsilon_{y_{US}}$ | 0.00 | 0.01 | 0.02 | 0.02 | 0.02 | 0.02 | 0.02 |
| $\pi_{US}$ | $\varepsilon_e$ | 0.11 | 0.13 | 0.10 | 0.05 | 0.03 | 0.01 | 0.01 |
| | $\varepsilon_{\pi_{US}}$ | 0.52 | 0.50 | 0.58 | 0.76 | 0.85 | 0.92 | 0.93 |
| | $\varepsilon_{i_{DE}}$ | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 |
| | $\varepsilon_{i_{US}}$ | 0.35 | 0.36 | 0.30 | 0.19 | 0.11 | 0.06 | 0.05 |
| | $\varepsilon_{y_{US}}$ | 0.02 | 0.02 | 0.02 | 0.01 | 0.01 | 0.01 | 0.00 |
| $i_{DE}$ | $\varepsilon_e$ | 0.00 | 0.00 | 0.00 | 0.01 | 0.02 | 0.01 | 0.01 |
| | $\varepsilon_{\pi_{US}}$ | 0.24 | 0.28 | 0.35 | 0.42 | 0.51 | 0.66 | 0.71 |
| | $\varepsilon_{i_{DE}}$ | 0.28 | 0.25 | 0.21 | 0.16 | 0.12 | 0.08 | 0.07 |
| | $\varepsilon_{i_{US}}$ | 0.23 | 0.23 | 0.25 | 0.29 | 0.28 | 0.20 | 0.17 |
| | $\varepsilon_{y_{US}}$ | 0.25 | 0.23 | 0.19 | 0.11 | 0.08 | 0.05 | 0.04 |
| $i_{US}$ | $\varepsilon_e$ | 0.00 | 0.00 | 0.01 | 0.03 | 0.02 | 0.01 | 0.01 |
| | $\varepsilon_{\pi_{US}}$ | 0.52 | 0.53 | 0.54 | 0.62 | 0.74 | 0.85 | 0.88 |
| | $\varepsilon_{i_{DE}}$ | 0.33 | 0.30 | 0.20 | 0.11 | 0.08 | 0.05 | 0.04 |
| | $\varepsilon_{i_{US}}$ | 0.15 | 0.16 | 0.19 | 0.17 | 0.10 | 0.05 | 0.04 |
| | $\varepsilon_{y_{US}}$ | 0.00 | 0.01 | 0.05 | 0.07 | 0.06 | 0.04 | 0.03 |
| $y_{US}$ | $\varepsilon_e$ | 0.00 | 0.00 | 0.00 | 0.01 | 0.01 | 0.01 | 0.00 |
| | $\varepsilon_{\pi_{US}}$ | 0.02 | 0.02 | 0.02 | 0.01 | 0.00 | 0.00 | 0.00 |
| | $\varepsilon_{i_{DE}}$ | 0.15 | 0.10 | 0.06 | 0.04 | 0.02 | 0.01 | 0.01 |
| | $\varepsilon_{i_{US}}$ | 0.02 | 0.05 | 0.12 | 0.15 | 0.11 | 0.06 | 0.04 |
| | $\varepsilon_{y_{US}}$ | 0.81 | 0.82 | 0.80 | 0.80 | 0.86 | 0.93 | 0.94 |

Table 9. Forecast error variance decomposition - U.S. - DE model - Pre-inflation targeting

| Var. | Shock | Horizon | | | | | | |
|---|---|---|---|---|---|---|---|---|
| | | 1 | 2 | 4 | 12 | 24 | 48 | 60 |
| $e - US/FR$ | $\varepsilon_e$ | 0.85 | 0.80 | 0.76 | 0.72 | 0.71 | 0.71 | 0.71 |
| | $\varepsilon_{\pi_{US}}$ | 0.01 | 0.02 | 0.03 | 0.06 | 0.06 | 0.06 | 0.06 |
| | $\varepsilon_{i_{FR}}$ | 0.00 | 0.00 | 0.00 | 0.02 | 0.02 | 0.02 | 0.02 |
| | $\varepsilon_{i_{US}}$ | 0.13 | 0.15 | 0.19 | 0.18 | 0.18 | 0.19 | 0.19 |
| | $\varepsilon_{y_{US}}$ | 0.01 | 0.02 | 0.02 | 0.02 | 0.02 | 0.02 | 0.02 |
| $\pi_{US}$ | $\varepsilon_e$ | 0.09 | 0.08 | 0.05 | 0.02 | 0.01 | 0.01 | 0.00 |
| | $\varepsilon_{\pi_{US}}$ | 0.72 | 0.73 | 0.79 | 0.83 | 0.89 | 0.94 | 0.96 |
| | $\varepsilon_{i_{FR}}$ | 0.00 | 0.02 | 0.04 | 0.10 | 0.07 | 0.04 | 0.03 |
| | $\varepsilon_{i_{US}}$ | 0.17 | 0.15 | 0.08 | 0.03 | 0.01 | 0.01 | 0.01 |
| | $\varepsilon_{y_{US}}$ | 0.02 | 0.02 | 0.01 | 0.01 | 0.01 | 0.01 | 0.01 |
| $i_{FR}$ | $\varepsilon_e$ | 0.00 | 0.00 | 0.00 | 0.01 | 0.03 | 0.02 | 0.02 |
| | $\varepsilon_{\pi_{US}}$ | 0.02 | 0.03 | 0.06 | 0.07 | 0.19 | 0.55 | 0.67 |
| | $\varepsilon_{i_{FR}}$ | 0.91 | 0.87 | 0.80 | 0.61 | 0.40 | 0.19 | 0.14 |
| | $\varepsilon_{i_{US}}$ | 0.05 | 0.08 | 0.12 | 0.30 | 0.36 | 0.23 | 0.17 |
| | $\varepsilon_{y_{US}}$ | 0.02 | 0.02 | 0.01 | 0.01 | 0.01 | 0.01 | 0.01 |
| $i_{US}$ | $\varepsilon_e$ | 0.00 | 0.00 | 0.01 | 0.07 | 0.05 | 0.03 | 0.02 |
| | $\varepsilon_{\pi_{US}}$ | 0.43 | 0.45 | 0.50 | 0.46 | 0.61 | 0.81 | 0.86 |
| | $\varepsilon_{i_{FR}}$ | 0.06 | 0.05 | 0.03 | 0.04 | 0.03 | 0.02 | 0.01 |
| | $\varepsilon_{i_{US}}$ | 0.50 | 0.49 | 0.44 | 0.39 | 0.27 | 0.13 | 0.10 |
| | $\varepsilon_{y_{US}}$ | 0.01 | 0.00 | 0.02 | 0.04 | 0.03 | 0.02 | 0.01 |
| $y_{US}$ | $\varepsilon_e$ | 0.00 | 0.00 | 0.00 | 0.02 | 0.01 | 0.01 | 0.01 |
| | $\varepsilon_{\pi_{US}}$ | 0.06 | 0.07 | 0.06 | 0.03 | 0.03 | 0.02 | 0.02 |
| | $\varepsilon_{i_{FR}}$ | 0.00 | 0.00 | 0.00 | 0.01 | 0.01 | 0.01 | 0.00 |
| | $\varepsilon_{i_{US}}$ | 0.01 | 0.03 | 0.06 | 0.07 | 0.05 | 0.03 | 0.02 |
| | $\varepsilon_{y_{US}}$ | 0.92 | 0.89 | 0.87 | 0.87 | 0.90 | 0.94 | 0.95 |

Table 10. Forecast error variance decomposition - U.S. - FR model - Pre-inflation targeting

| Var. | Shock | Horizon | | | | | | |
|---|---|---|---|---|---|---|---|---|
| | | 1 | 2 | 4 | 12 | 24 | 48 | 60 |
| | $\varepsilon_e$ | 0.65 | 0.65 | 0.66 | 0.65 | 0.65 | 0.65 | 0.65 |
| | $\varepsilon_{\pi_{US}}$ | 0.16 | 0.17 | 0.18 | 0.17 | 0.17 | 0.17 | 0.17 |
| $e - US/AU$ | $\varepsilon_{i_{AU}}$ | 0.01 | 0.01 | 0.01 | 0.03 | 0.03 | 0.03 | 0.03 |
| | $\varepsilon_{i_{US}}$ | 0.16 | 0.16 | 0.14 | 0.14 | 0.14 | 0.14 | 0.14 |
| | $\varepsilon_{y_{US}}$ | 0.02 | 0.02 | 0.01 | 0.02 | 0.02 | 0.02 | 0.02 |
| | $\varepsilon_e$ | 0.18 | 0.14 | 0.13 | 0.09 | 0.06 | 0.04 | 0.03 |
| | $\varepsilon_{\pi_{US}}$ | 0.56 | 0.58 | 0.63 | 0.70 | 0.80 | 0.87 | 0.89 |
| $\pi_{US}$ | $\varepsilon_{i_{AU}}$ | 0.00 | 0.00 | 0.00 | 0.01 | 0.01 | 0.01 | 0.00 |
| | $\varepsilon_{i_{US}}$ | 0.10 | 0.11 | 0.07 | 0.06 | 0.04 | 0.03 | 0.02 |
| | $\varepsilon_{y_{US}}$ | 0.16 | 0.18 | 0.16 | 0.14 | 0.10 | 0.06 | 0.05 |
| | $\varepsilon_e$ | 0.00 | 0.00 | 0.01 | 0.08 | 0.16 | 0.15 | 0.12 |
| | $\varepsilon_{\pi_{US}}$ | 0.01 | 0.01 | 0.00 | 0.01 | 0.07 | 0.32 | 0.45 |
| $i_{AU}$ | $\varepsilon_{i_{AU}}$ | 0.85 | 0.88 | 0.92 | 0.86 | 0.70 | 0.45 | 0.36 |
| | $\varepsilon_{i_{US}}$ | 0.10 | 0.08 | 0.05 | 0.03 | 0.04 | 0.04 | 0.03 |
| | $\varepsilon_{y_{US}}$ | 0.04 | 0.03 | 0.02 | 0.02 | 0.03 | 0.04 | 0.04 |
| | $\varepsilon_e$ | 0.00 | 0.00 | 0.00 | 0.14 | 0.11 | 0.05 | 0.04 |
| | $\varepsilon_{\pi_{US}}$ | 0.47 | 0.49 | 0.56 | 0.53 | 0.66 | 0.84 | 0.88 |
| $i_{US}$ | $\varepsilon_{i_{AU}}$ | 0.16 | 0.16 | 0.14 | 0.13 | 0.09 | 0.05 | 0.03 |
| | $\varepsilon_{i_{US}}$ | 0.37 | 0.35 | 0.25 | 0.13 | 0.07 | 0.03 | 0.02 |
| | $\varepsilon_{y_{US}}$ | 0.00 | 0.01 | 0.05 | 0.08 | 0.06 | 0.03 | 0.02 |
| | $\varepsilon_e$ | 0.10 | 0.07 | 0.04 | 0.02 | 0.02 | 0.01 | 0.01 |
| | $\varepsilon_{\pi_{US}}$ | 0.03 | 0.04 | 0.05 | 0.03 | 0.03 | 0.02 | 0.02 |
| $y_{US}$ | $\varepsilon_{i_{AU}}$ | 0.02 | 0.03 | 0.05 | 0.02 | 0.01 | 0.01 | 0.01 |
| | $\varepsilon_{i_{US}}$ | 0.04 | 0.02 | 0.01 | 0.01 | 0.01 | 0.00 | 0.00 |
| | $\varepsilon_{y_{US}}$ | 0.82 | 0.83 | 0.85 | 0.92 | 0.94 | 0.96 | 0.97 |

Table 11. Forecast error variance decomposition - U.S. - AU model - Pre-inflation targeting

| Var. | Shock | Horizon | | | | | | |
|---|---|---|---|---|---|---|---|---|
| | | 1 | 2 | 4 | 12 | 24 | 48 | 60 |
| $e-US/CH$ | $\varepsilon_e$ | 0.77 | 0.77 | 0.73 | 0.70 | 0.69 | 0.69 | 0.69 |
| | $\varepsilon_{\pi_{US}}$ | 0.00 | 0.00 | 0.02 | 0.05 | 0.05 | 0.05 | 0.05 |
| | $\varepsilon_{i_{CH}}$ | 0.00 | 0.01 | 0.02 | 0.02 | 0.03 | 0.03 | 0.03 |
| | $\varepsilon_{i_{US}}$ | 0.12 | 0.16 | 0.19 | 0.18 | 0.18 | 0.18 | 0.18 |
| | $\varepsilon_{y_{US}}$ | 0.10 | 0.06 | 0.05 | 0.05 | 0.06 | 0.06 | 0.06 |
| $\pi_{US}$ | $\varepsilon_e$ | 0.07 | 0.06 | 0.04 | 0.01 | 0.01 | 0.00 | 0.00 |
| | $\varepsilon_{\pi_{US}}$ | 0.75 | 0.75 | 0.83 | 0.92 | 0.95 | 0.97 | 0.98 |
| | $\varepsilon_{i_{CH}}$ | 0.00 | 0.00 | 0.00 | 0.01 | 0.01 | 0.00 | 0.00 |
| | $\varepsilon_{i_{US}}$ | 0.15 | 0.15 | 0.09 | 0.04 | 0.02 | 0.01 | 0.01 |
| | $\varepsilon_{y_{US}}$ | 0.03 | 0.03 | 0.04 | 0.02 | 0.02 | 0.01 | 0.01 |
| $i_{CH}$ | $\varepsilon_e$ | 0.00 | 0.00 | 0.01 | 0.02 | 0.01 | 0.01 | 0.01 |
| | $\varepsilon_{\pi_{US}}$ | 0.21 | 0.21 | 0.22 | 0.21 | 0.22 | 0.35 | 0.43 |
| | $\varepsilon_{i_{CH}}$ | 0.36 | 0.34 | 0.33 | 0.29 | 0.24 | 0.18 | 0.15 |
| | $\varepsilon_{i_{US}}$ | 0.04 | 0.06 | 0.09 | 0.21 | 0.32 | 0.31 | 0.27 |
| | $\varepsilon_{y_{US}}$ | 0.38 | 0.39 | 0.34 | 0.27 | 0.21 | 0.15 | 0.13 |
| $i_{US}$ | $\varepsilon_e$ | 0.00 | 0.00 | 0.00 | 0.03 | 0.02 | 0.01 | 0.01 |
| | $\varepsilon_{\pi_{US}}$ | 0.51 | 0.50 | 0.45 | 0.45 | 0.61 | 0.81 | 0.85 |
| | $\varepsilon_{i_{CH}}$ | 0.04 | 0.04 | 0.03 | 0.02 | 0.02 | 0.01 | 0.01 |
| | $\varepsilon_{i_{US}}$ | 0.36 | 0.42 | 0.49 | 0.46 | 0.30 | 0.14 | 0.11 |
| | $\varepsilon_{y_{US}}$ | 0.08 | 0.04 | 0.03 | 0.04 | 0.04 | 0.02 | 0.02 |
| $y_{US}$ | $\varepsilon_e$ | 0.04 | 0.04 | 0.02 | 0.01 | 0.00 | 0.00 | 0.00 |
| | $\varepsilon_{\pi_{US}}$ | 0.01 | 0.01 | 0.01 | 0.06 | 0.06 | 0.04 | 0.03 |
| | $\varepsilon_{i_{CH}}$ | 0.20 | 0.19 | 0.14 | 0.07 | 0.03 | 0.02 | 0.01 |
| | $\varepsilon_{i_{US}}$ | 0.13 | 0.19 | 0.30 | 0.26 | 0.19 | 0.10 | 0.08 |
| | $\varepsilon_{y_{US}}$ | 0.62 | 0.58 | 0.53 | 0.61 | 0.71 | 0.84 | 0.87 |

Table 12. Forecast error variance decomposition - U.S. - CH model Pre-inflation targeting

### 4. Forecast Error Variance Decomposition - Japan and euro area

Here we discuss the forecast error variance decomposition of the models for Japan and the euro area. For Japan in the non-stationary sample (1974M1 to 1994M12), as seen in Table 13 and in line with the main model most of the forecast error variance of the rate of change of the exchange rate continues to be explained by its own shock, followed by the contribution of the temporary monetary policy shock in the U.S. economy, while the permanent monetary policy shock in the U.S. explains a small share. As expected, the permanent U.S. monetary policy shock is the main driver of the forecast error variance of both inflation and interest rates in the U.S. economy. Unlike in the baseline model, the forecast error variance of interest rates in Japan is now mostly explained by Japan's monetary policy shock. Recall that cointegration between interest rates is not imposed in this version of the model.

In the stationary sample (1995M1 to 2017M6), see Table 14, the forecast error variance of $e_t$ is now mostly explained by the temporary monetary policy shock in the U.S. economy, even after three years, followed by the temporary monetary policy shock in Japan. The shock on $e_t$ now only accounts for a residual share. Nevertheless, the permanent U.S. monetary policy shock continues to explain a small amount of the forecasts error variance of $e_t$, as concluded in the other models.

Regarding the FEVD for the model with the euro area, see Table 15, the forecast error variance of $e_t$ is strongly accounted for by the temporary monetary policy shocks (combined account for 73% after three years), with the shock in itself explaining 18% after three years. The permanent monetary policy shocks continue to explain only a residual amount (6% after three years), in line with the results of the baseline model. The permanent monetary policy shocks in both economies continues to be the main driver of the FEVD of the respective interest rate and inflation at longer horizons.

| Var. | Shock | Horizon | | | | | | |
|---|---|---|---|---|---|---|---|---|
| | | 1 | 2 | 4 | 12 | 24 | 48 | 60 |
| $e - US/JP$ | $\varepsilon_e$ | 0.84 | 0.83 | 0.84 | 0.84 | 0.84 | 0.84 | 0.84 |
| | $\varepsilon_{\pi_{US}}$ | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 |
| | $\varepsilon_{i_{JP}}$ | 0.07 | 0.05 | 0.05 | 0.05 | 0.05 | 0.05 | 0.05 |
| | $\varepsilon_{i_{US}}$ | 0.09 | 0.11 | 0.11 | 0.11 | 0.11 | 0.11 | 0.11 |
| | $\varepsilon_{y_{US}}$ | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 |
| $\pi_{US}$ | $\varepsilon_e$ | 0.03 | 0.02 | 0.02 | 0.02 | 0.02 | 0.01 | 0.01 |
| | $\varepsilon_{\pi_{US}}$ | 0.54 | 0.53 | 0.62 | 0.74 | 0.81 | 0.86 | 0.88 |
| | $\varepsilon_{i_{JP}}$ | 0.00 | 0.01 | 0.01 | 0.01 | 0.00 | 0.00 | 0.00 |
| | $\varepsilon_{i_{US}}$ | 0.19 | 0.18 | 0.12 | 0.07 | 0.06 | 0.04 | 0.04 |
| | $\varepsilon_{y_{US}}$ | 0.24 | 0.26 | 0.23 | 0.15 | 0.11 | 0.08 | 0.07 |
| $i_{JP}$ | $\varepsilon_e$ | 0.01 | 0.00 | 0.00 | 0.01 | 0.01 | 0.00 | 0.00 |
| | $\varepsilon_{\pi_{US}}$ | 0.00 | 0.00 | 0.01 | 0.05 | 0.04 | 0.02 | 0.02 |
| | $\varepsilon_{i_{JP}}$ | 0.94 | 0.94 | 0.89 | 0.85 | 0.90 | 0.94 | 0.96 |
| | $\varepsilon_{i_{US}}$ | 0.04 | 0.03 | 0.05 | 0.04 | 0.02 | 0.01 | 0.01 |
| | $\varepsilon_{y_{US}}$ | 0.01 | 0.02 | 0.04 | 0.05 | 0.04 | 0.02 | 0.02 |
| $i_{US}$ | $\varepsilon_e$ | 0.00 | 0.00 | 0.00 | 0.02 | 0.01 | 0.00 | 0.00 |
| | $\varepsilon_{\pi_{US}}$ | 0.66 | 0.69 | 0.71 | 0.75 | 0.85 | 0.94 | 0.95 |
| | $\varepsilon_{i_{JP}}$ | 0.01 | 0.01 | 0.01 | 0.01 | 0.00 | 0.00 | 0.00 |
| | $\varepsilon_{i_{US}}$ | 0.32 | 0.29 | 0.21 | 0.12 | 0.07 | 0.03 | 0.02 |
| | $\varepsilon_{y_{US}}$ | 0.00 | 0.01 | 0.06 | 0.10 | 0.07 | 0.03 | 0.02 |
| $y_{US}$ | $\varepsilon_e$ | 0.04 | 0.05 | 0.03 | 0.01 | 0.00 | 0.00 | 0.00 |
| | $\varepsilon_{\pi_{US}}$ | 0.11 | 0.13 | 0.12 | 0.04 | 0.02 | 0.01 | 0.01 |
| | $\varepsilon_{i_{JP}}$ | 0.00 | 0.01 | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 |
| | $\varepsilon_{i_{US}}$ | 0.13 | 0.08 | 0.03 | 0.01 | 0.00 | 0.00 | 0.00 |
| | $\varepsilon_{y_{US}}$ | 0.71 | 0.73 | 0.81 | 0.95 | 0.98 | 0.99 | 0.99 |

Table 13. Forecast error variance decomposition - U.S. - JP model - Non-stationary sample 1974M1 to 1994M12.

| Var. | Shock | Horizon | | | | | | |
|---|---|---|---|---|---|---|---|---|
| | | 1 | 2 | 4 | 12 | 24 | 48 | 60 |
| $e-US/JP$ | $\varepsilon_e$ | 0.06 | 0.04 | 0.06 | 0.11 | 0.11 | 0.11 | 0.11 |
| | $\varepsilon_{\pi_{US}}$ | 0.04 | 0.03 | 0.03 | 0.03 | 0.03 | 0.03 | 0.03 |
| | $\varepsilon_{i_{JP}}$ | 0.10 | 0.15 | 0.16 | 0.15 | 0.15 | 0.15 | 0.15 |
| | $\varepsilon_{i_{US}}$ | 0.80 | 0.78 | 0.75 | 0.69 | 0.69 | 0.69 | 0.69 |
| | $\varepsilon_{y_{US}}$ | 0.00 | 0.00 | 0.01 | 0.02 | 0.02 | 0.02 | 0.02 |
| $\pi_{US}$ | $\varepsilon_e$ | 0.85 | 0.80 | 0.78 | 0.77 | 0.66 | 0.46 | 0.38 |
| | $\varepsilon_{\pi_{US}}$ | 0.01 | 0.02 | 0.02 | 0.07 | 0.18 | 0.42 | 0.51 |
| | $\varepsilon_{i_{JP}}$ | 0.00 | 0.00 | 0.00 | 0.00 | 0.02 | 0.03 | 0.02 |
| | $\varepsilon_{i_{US}}$ | 0.10 | 0.13 | 0.14 | 0.12 | 0.10 | 0.07 | 0.06 |
| | $\varepsilon_{y_{US}}$ | 0.04 | 0.05 | 0.05 | 0.04 | 0.04 | 0.03 | 0.02 |
| $i_{JP}$ | $\varepsilon_e$ | 0.00 | 0.00 | 0.00 | 0.01 | 0.01 | 0.01 | 0.01 |
| | $\varepsilon_{\pi_{US}}$ | 0.53 | 0.42 | 0.36 | 0.28 | 0.27 | 0.26 | 0.26 |
| | $\varepsilon_{i_{JP}}$ | 0.20 | 0.27 | 0.34 | 0.54 | 0.58 | 0.58 | 0.58 |
| | $\varepsilon_{i_{US}}$ | 0.14 | 0.13 | 0.13 | 0.09 | 0.08 | 0.07 | 0.07 |
| | $\varepsilon_{y_{US}}$ | 0.12 | 0.18 | 0.17 | 0.09 | 0.07 | 0.07 | 0.07 |
| $i_{US}$ | $\varepsilon_e$ | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 |
| | $\varepsilon_{\pi_{US}}$ | 0.35 | 0.37 | 0.38 | 0.53 | 0.72 | 0.87 | 0.90 |
| | $\varepsilon_{i_{JP}}$ | 0.52 | 0.52 | 0.54 | 0.43 | 0.26 | 0.12 | 0.09 |
| | $\varepsilon_{i_{US}}$ | 0.00 | 0.00 | 0.00 | 0.01 | 0.01 | 0.00 | 0.00 |
| | $\varepsilon_{y_{US}}$ | 0.13 | 0.11 | 0.07 | 0.02 | 0.01 | 0.01 | 0.00 |
| $y_{US}$ | $\varepsilon_e$ | 0.03 | 0.03 | 0.02 | 0.01 | 0.01 | 0.00 | 0.00 |
| | $\varepsilon_{\pi_{US}}$ | 0.16 | 0.16 | 0.13 | 0.09 | 0.06 | 0.03 | 0.02 |
| | $\varepsilon_{i_{JP}}$ | 0.14 | 0.16 | 0.21 | 0.19 | 0.12 | 0.06 | 0.05 |
| | $\varepsilon_{i_{US}}$ | 0.02 | 0.01 | 0.01 | 0.00 | 0.00 | 0.00 | 0.00 |
| | $\varepsilon_{y_{US}}$ | 0.66 | 0.64 | 0.63 | 0.70 | 0.81 | 0.91 | 0.93 |

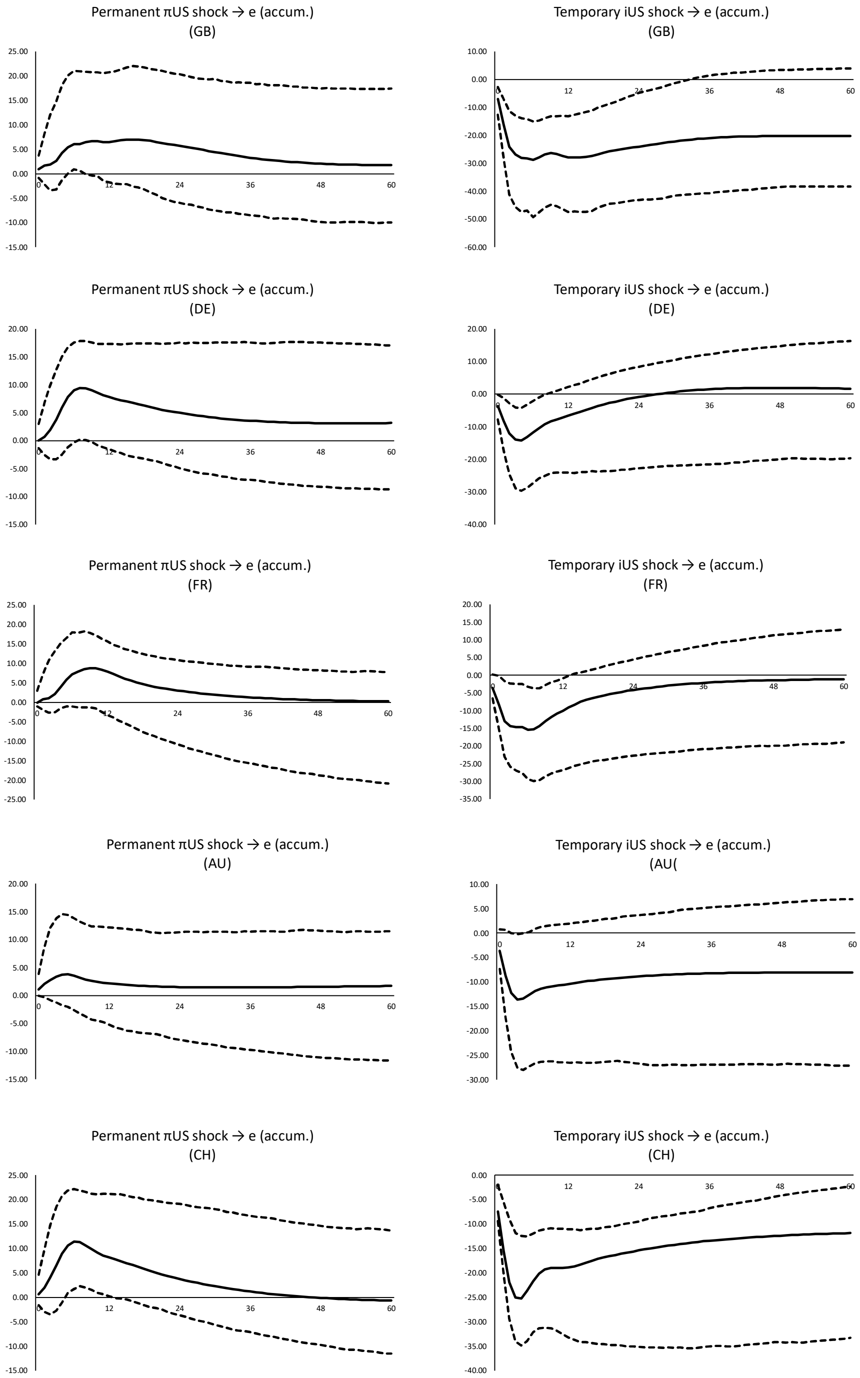
Table 14. Forecast error variance decomposition - U.S. - JP model - Stationary sample 1995M1 to 2017M6

| Var. | Shock | Horizon | | | | | | |
|---|---|---|---|---|---|---|---|---|
| | | 1 | 2 | 4 | 12 | 24 | 48 | 60 |
| | $\varepsilon_e$ | 0.22 | 0.21 | 0.19 | 0.18 | 0.18 | 0.18 | 0.18 |
| | $\varepsilon_{\pi_{US}}$ | 0.00 | 0.00 | 0.01 | 0.03 | 0.03 | 0.03 | 0.03 |
| | $\varepsilon_{\pi_{EA}}$ | 0.00 | 0.00 | 0.00 | 0.02 | 0.03 | 0.03 | 0.03 |
| $e - US/EA$ | $\varepsilon_{i_{US}}$ | 0.11 | 0.12 | 0.15 | 0.15 | 0.15 | 0.16 | 0.16 |
| | $\varepsilon_{i_{EA}}$ | 0.64 | 0.63 | 0.62 | 0.59 | 0.57 | 0.57 | 0.57 |
| | $\varepsilon_{y_{US}}$ | 0.01 | 0.02 | 0.02 | 0.02 | 0.02 | 0.02 | 0.02 |
| | $\varepsilon_{y_{EA}}$ | 0.02 | 0.02 | 0.02 | 0.02 | 0.02 | 0.02 | 0.02 |
| | $\varepsilon_e$ | 0.17 | 0.18 | 0.14 | 0.08 | 0.05 | 0.03 | 0.02 |
| | $\varepsilon_{\pi_{US}}$ | 0.22 | 0.22 | 0.33 | 0.52 | 0.67 | 0.81 | 0.84 |
| | $\varepsilon_{\pi_{EA}}$ | 0.18 | 0.19 | 0.19 | 0.13 | 0.08 | 0.05 | 0.04 |
| $\pi_{US}$ | $\varepsilon_{i_{US}}$ | 0.27 | 0.24 | 0.19 | 0.16 | 0.12 | 0.07 | 0.06 |
| | $\varepsilon_{i_{EA}}$ | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 |
| | $\varepsilon_{y_{US}}$ | 0.16 | 0.16 | 0.15 | 0.11 | 0.08 | 0.05 | 0.04 |
| | $\varepsilon_{y_{EA}}$ | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 |
| | $\varepsilon_e$ | 0.59 | 0.58 | 0.62 | 0.54 | 0.33 | 0.14 | 0.11 |
| | $\varepsilon_{\pi_{US}}$ | 0.01 | 0.00 | 0.01 | 0.03 | 0.05 | 0.06 | 0.06 |
| | $\varepsilon_{\pi_{EA}}$ | 0.09 | 0.14 | 0.17 | 0.30 | 0.48 | 0.71 | 0.77 |
| $\pi_{EA}$ | $\varepsilon_{i_{US}}$ | 0.00 | 0.01 | 0.01 | 0.05 | 0.08 | 0.05 | 0.04 |
| | $\varepsilon_{i_{EA}}$ | 0.29 | 0.25 | 0.18 | 0.07 | 0.03 | 0.01 | 0.01 |
| | $\varepsilon_{y_{US}}$ | 0.02 | 0.01 | 0.01 | 0.01 | 0.03 | 0.02 | 0.02 |
| | $\varepsilon_{y_{EA}}$ | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 |
| | $\varepsilon_e$ | 0.00 | 0.00 | 0.01 | 0.01 | 0.01 | 0.01 | 0.01 |
| | $\varepsilon_{\pi_{US}}$ | 0.84 | 0.86 | 0.88 | 0.90 | 0.93 | 0.96 | 0.97 |
| | $\varepsilon_{\pi_{EA}}$ | 0.00 | 0.00 | 0.01 | 0.02 | 0.02 | 0.01 | 0.01 |
| $i_{US}$ | $\varepsilon_{i_{US}}$ | 0.16 | 0.14 | 0.09 | 0.05 | 0.03 | 0.02 | 0.01 |
| | $\varepsilon_{i_{EA}}$ | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 |
| | $\varepsilon_{y_{US}}$ | 0.00 | 0.00 | 0.01 | 0.01 | 0.01 | 0.01 | 0.01 |
| | $\varepsilon_{y_{EA}}$ | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 |
| | $\varepsilon_e$ | 0.00 | 0.00 | 0.01 | 0.00 | 0.00 | 0.00 | 0.00 |
| | $\varepsilon_{\pi_{US}}$ | 0.00 | 0.00 | 0.02 | 0.05 | 0.06 | 0.06 | 0.06 |
| | $\varepsilon_{\pi_{EA}}$ | 0.78 | 0.77 | 0.79 | 0.83 | 0.88 | 0.91 | 0.92 |
| $i_{EA}$ | $\varepsilon_{i_{US}}$ | 0.09 | 0.08 | 0.07 | 0.05 | 0.03 | 0.01 | 0.01 |
| | $\varepsilon_{i_{EA}}$ | 0.02 | 0.01 | 0.01 | 0.00 | 0.00 | 0.00 | 0.00 |
| | $\varepsilon_{y_{US}}$ | 0.11 | 0.12 | 0.10 | 0.06 | 0.03 | 0.01 | 0.01 |
| | $\varepsilon_{y_{EA}}$ | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 |
| | $\varepsilon_e$ | 0.00 | 0.00 | 0.00 | 0.01 | 0.01 | 0.00 | 0.00 |
| | $\varepsilon_{\pi_{US}}$ | 0.11 | 0.14 | 0.18 | 0.17 | 0.14 | 0.12 | 0.12 |
| | $\varepsilon_{\pi_{EA}}$ | 0.02 | 0.02 | 0.04 | 0.09 | 0.10 | 0.09 | 0.09 |
| $y_{US}$ | $\varepsilon_{i_{US}}$ | 0.18 | 0.16 | 0.12 | 0.06 | 0.03 | 0.01 | 0.01 |
| | $\varepsilon_{i_{EA}}$ | 0.02 | 0.02 | 0.02 | 0.01 | 0.00 | 0.00 | 0.00 |
| | $\varepsilon_{y_{US}}$ | 0.66 | 0.64 | 0.63 | 0.67 | 0.72 | 0.77 | 0.79 |
| | $\varepsilon_{y_{EA}}$ | 0.01 | 0.01 | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 |
| | $\varepsilon_e$ | 0.01 | 0.01 | 0.01 | 0.00 | 0.00 | 0.00 | 0.00 |
| | $\varepsilon_{\pi_{US}}$ | 0.00 | 0.00 | 0.00 | 0.01 | 0.01 | 0.01 | 0.01 |
| | $\varepsilon_{\pi_{EA}}$ | 0.02 | 0.02 | 0.04 | 0.07 | 0.11 | 0.13 | 0.14 |
| $y_{EA}$ | $\varepsilon_{i_{US}}$ | 0.00 | 0.01 | 0.02 | 0.03 | 0.02 | 0.01 | 0.01 |
| | $\varepsilon_{i_{EA}}$ | 0.01 | 0.01 | 0.01 | 0.00 | 0.00 | 0.00 | 0.00 |
| | $\varepsilon_{y_{US}}$ | 0.09 | 0.08 | 0.06 | 0.04 | 0.02 | 0.01 | 0.01 |
| | $\varepsilon_{y_{US}}$ | 0.88 | 0.87 | 0.86 | 0.85 | 0.84 | 0.83 | 0.83 |

Table 15. Forecast error variance decomposition - U.S. - EA model

## 5. Accumulated Impulse Response Functions

Since we consider the variation in exchange rates in all the exercises, it is useful to look at the behaviour of exchange rates themselves. A proxy for this can be the accumulated impulse response functions. In Figure 10 we report accumulated IRFs of the five models for the permanent and temporary U.S. monetary policy shocks on the rate of change of the exchange rate: a permanent U.S. monetary policy shock results in a depreciation of the USD, while temporary U.S. monetary policy shocks have the opposite impact. As expected, the accumulated impacts from temporary monetary policy shocks are stronger and significant for a longer period of time, while the accumulated effects of permanent monetary policy shocks are smaller and short-lived. In fact, for the models with FR and AU, where IRFs for the permanent monetary policy shock on $e_t$ are statistically significant for a very short period of time, the accumulated IRFs that are not significant.

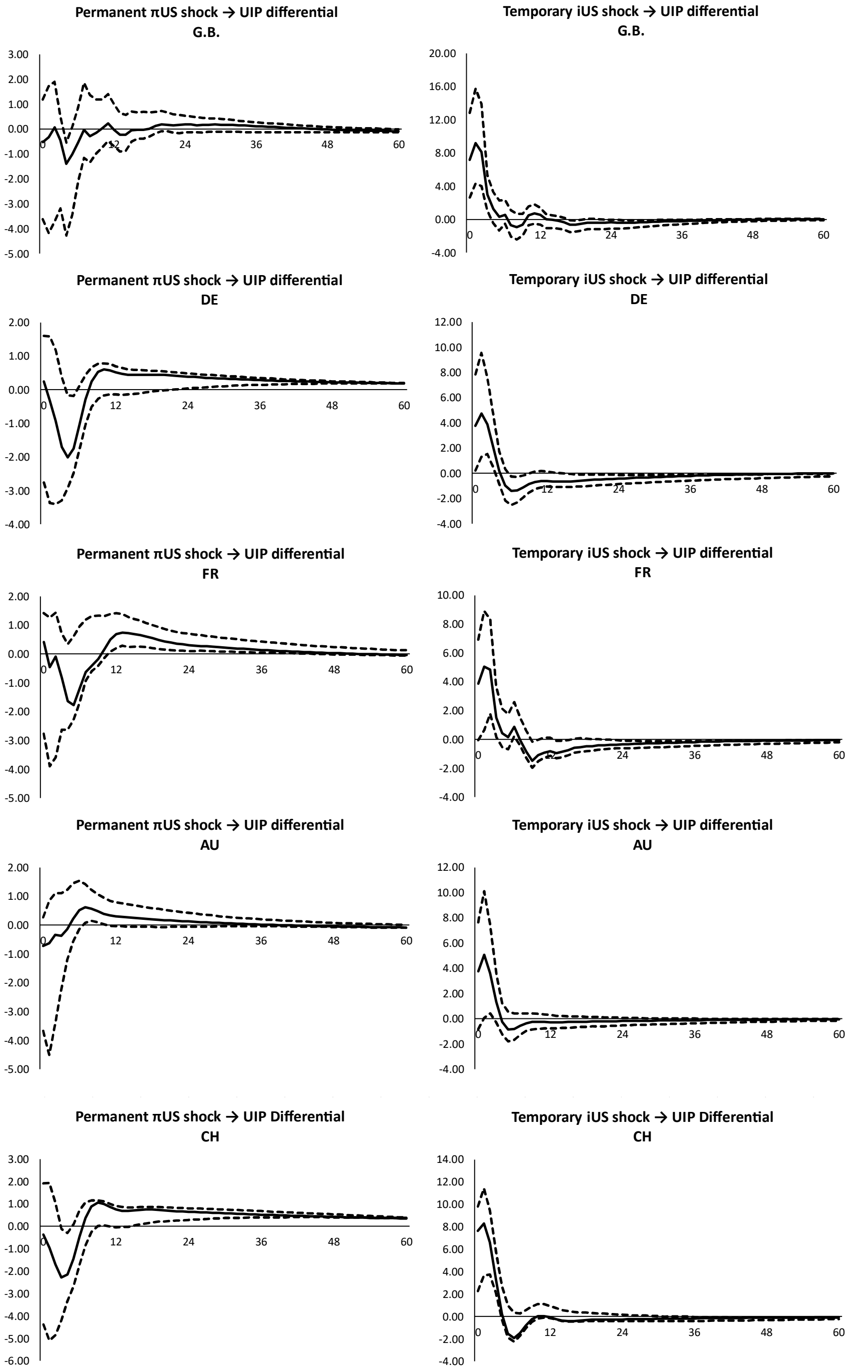



All responses are in percentage points. The first column has the responses to the permanent U.S. monetary shock for the various countries and the second column the response to the temporary monetary shock. Shocks are identified as described in the text. The structural model is estimated by Maximum Likelihood, using the Scoring Algorithm of Amisano and Giannini (1997) as implemented in JMulTi. Response to a one standard deviation shock. Full Sample. 90 % Hall Bootstrap confidence intervals.

Figure 10: Accumulated Impulse Response Functions - Effects on USD exchange rate

## 6. Impulse Response - Deviations from Uncovered Interest Parity

The analysis of the temporary and permanent monetary policy shocks can also be performed by looking at the deviations from UIRP. In the spirit of Schmitt-Grohé and Uribe (2021), in Figure 11 we present the impact on these deviations. The deviations are computed as the impact of the identified shocks on the U.S. interest rate, subtracted from the impact on the interest rate of the second economy and the impact on $e_t$. This implies that positive (negative) values correspond to deviations in UIRP parity in favour of the U.S. (foreign) economy. The results support their main findings: while the temporary monetary policy shock generates a deviation in favour of the USD and against the predictions of UIRP, the permanent monetary policy shock creates a deviation in favour of the foreign currency (less so in the models with FR and specially AU, given that these models, as mentioned before, display small USD depreciations following the permanent monetary policy shock in the U.S. economy).




All responses are in percentage points. The first column has the responses to the permanent U.S. monetary shock for the various countries and the second column the response to the temporary monetary shock. Shocks are identified as described in the text. The structural model is estimated by Maximum Likelihood, using the Scoring Algorithm of Amisano and Giannini (1997) as implemented in JMulTi. Response to a one standard deviation shock. Full Sample. 90 % Hall Bootstrap confidence intervals.

Figure 11: Deviations from uncovered interest rate parity

**References**

Amisano, G. and C. Giannini (1997). *Topics in Structural VAR Econometrics*. Springer, 2nd edn., Berlin.

Arouri, M., F. Jawadi, and N. Nguyen (2013). "What can we tell about monetary policy synchronization and interdependence over the 2007-2009 global financial crisis?" *Journal of Macroeconomics*, 36, 175–187.

Belke, A. and Y. Cui (2010). "US–Euro Area Monetary Policy Interdependence: New Evidence from Taylor Rule-based VECMs." *The World Economy*, 33(5), 778–797.

Lütkepohl, H. (2006). *New Introduction to Multiple Time Series Analysis*. Springer, Berlin and Heidelberg.

Schmitt-Grohé, S. and M. Uribe (2021). "Exchange Rates and Uncovered Interest Differentials: The Role of Permanent Monetary Shocks." *Journal of International Economics*, (forthcoming).

Uribe, M. (2021). "The Neo-Fisher Effect: Econometric Evidence from Empirical and Optimizing Models." *American Economic Journal: Macroeconomics*, (forthcoming).

Valle e Azevedo, J., J. Ritto, and P. Teles (2019). "The Neutrality of Nominal Rates: How Long is the Long Run?" *Banco de Portugal Working Papers*, (11).

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