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Permanent and temporary monetary policy shocks and the dynamics of exchange rates[☆]

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ABSTRACT

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 several advanced economies from 1971 to 2019 and resorting to a simple structural vector error correction (SVEC) model, 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 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 – results in a depreciation of the USD, in the short as well as over the long run that may contribute to higher (not lower) inflation also in the short run.

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. This is 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, see Miller (2014) for a detailed review of this topic and Froot and Frankel (1989) for early explanations.

☆ The views here are entirely our own and not necessarily those of Banco de Portugal or the Eurosystem.

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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, see Lothian (2016) for a careful documentation. These short-run and long-run dynamics can be reconciled. In the model of Dornbusch (1976), following an expansionary monetary policy shock, the exchange rate depreciates significantly before appreciating to values consistent with UIRP, the so-called overshooting effect.

This paper asks whether different types of nominal disturbances affect exchange rates differently and whether such distinction helps understanding the short-run and long-run relations between nominal rates and exchange rates, in a context where nominal disturbances are typically seen as important drivers of exchange rate movements. We start by assuming that monetary policy shocks (or nominal disturbances) come in two types, just as in Schmitt-Grohé and Uribe (2022): temporary and permanent nominal shocks. This is warranted because nominal interest rates are characterised by long-run variability, likewise a unit-root behaviour. The FED is not constantly raising or cutting policy interest rates unexpectedly and temporarily. It can potentially perturb the long-run level of interest rates (and of interest rate differentials), e.g., if the zero lower bound period persists and likewise if it normalises monetary policy during a protracted period. Are the effects of this policy different from the standard ones? What is the short-run impact of these permanent shocks? These questions are particularly relevant for policy-makers as 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. However, if such shock is a permanent one – one leading to permanently higher interest rates – and its effects over the short run are akin to the long-run effects, then the ensuing USD depreciation comes 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.

The work of Schmitt-Grohé and Uribe (2022) 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. 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 is in line with De Michelis and Iacoviello (2016), who evaluated the impact of an inflation target shock (a permanent shock) on the exchange rate for the Japanese and U.S. economies and found some evidence of domestic currency depreciation after an increase in the inflation target, in a setting where there is no distinction between temporary and permanent monetary policy shocks.

Our work seeks to verify whether the aforementioned diverse results hold by considering a broader set of economies, while employing less stringent (or more innocuous) identifying restrictions and resorting to a much 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; we also impose cointegration and between the nominal interest rate and the inflation rate in the U.S. economy, but unlike Schmitt-Grohé and Uribe (2022), we do not require the coefficients to be one. We can thus account for the well documented slow fall in real interest rates over the past decades, see e.g. Holston et al. (2017), Gordon (2014) or Summers (2014). Importantly, also unlike Schmitt-Grohé and Uribe (2022), we do not impose that the effects of monetary policy shocks on inflation and output are the usual ones. We are as much as possible agnostic on these effects, just as in Valle e Azevedo et al. (2022). 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 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 consider several combinations of possible identifications restrictions, many of which are rather innocuous.

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 turn, 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 where each economy is treated in the same way for identification purposes.

Our results show, with remarkable consistency across countries, setups, identification restrictions and sub-samples, 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 (2022), who find that permanent monetary policy shocks explain the majority of the FEVD of the exchange rate. This divergence can result from their assumption of non-stationary of the rate of change of the exchange rate (for which we do not find statistical support) coupled with the imposed cointegration between this variable and the permanent monetary policy shocks. We provide evidence that these differences go some way in explaining the divergent results.

Another robust feature of our results is that the "neo-Fisher" effect firstly uncovered by Uribe (2022) and corroborated by Valle e Azevedo et al. (2022), i.e., the fact that a permanent increase in policy rates may have a positive impact on inflation even in the short run, survives this opening of the U.S. economy. This has significant consequences for the conduct of monetary policy, in the sense that usual prescriptions in terms of forward guidance or make-up strategies (to raise inflation) or the normalisation of monetary policy (to tame inflation) could be turned upside down, depending on the perturbation such policies induce on the long-run level of nominal rates.

Finally, our results provide new insights on the delayed overshooting puzzle (see, e.g. Eichenbaum and Evans (1995) and the recent discussion in Kim et al. (2017)) as they help reviving Dornbusch's overshooting, again in contrast to Schmitt-Grohé and Uribe (2022), where the effects of the two types of monetary disturbances lack overshooting and even delayed overshooting.

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 Supplemental 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). It is worth mentioning that during our sample period, several European countries were under exchange rate arrangements. The biggest euro area countries except Spain were under the Exchange Rate Mechanism (ERM) since 1979, which limited the variability of exchange rates. This can limit the interest of exploring both DE and FR, given their similarity. However, G.B. is outside the ERM most of the time and CH only adopted a peg to the euro between 2011 and 2015, which together with AU and JP yields a reasonable amount of cross-country variability.

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[2];
- 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, i_t and i_t^*, respectively;
- U.S. core inflation, measured by the year-on-year rate of change of the CPI excluding food and energy - π_t;
- U.S. industrial production index, as a proxy for output at a monthly frequency - y_t.

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. Industrial production is used as a proxy for output, given that it 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. A complete description of the data can be found in Appendix A. In our sample all variables are found to be I(1), with the exception of e_t, which is found to be I(0). Specifically, 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. Hence, it will be assumed that this variable does not have any long-run relation with any other variable in the model. This is a departure from the main specification in Schmitt-Grohé and Uribe (2022), where it is assumed that e_t is non-stationary.

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, in line with the evidence of cointegration of monetary policies in Belke and Cui (2010) or Arouri et al. (2013). Given this evidence, we henceforth assume that nominal interest rates are cointegrated (except for Japan), though not necessarily with a coefficient equal to unity. Finally, as in Valle e Azevedo et al. (2022), we assume cointegration between nominal interest rates and inflation rates for the U.S. and the euro area and the coefficients are not restricted to one, again unlike Schmitt-Grohé and Uribe (2022). Not imposing a unitary cointegration coefficient in the Fisher relation allows for slowly moving long-run real interest rates. A coefficient less than one means that a fall in nominal rates is not fully matched by a fall in inflation, hence real rates are falling. This helps accounting, even if in a rough way, for the well-established fall in real interest rates observed in several advanced economies over the past decades, see, e.g., Gordon (2014), Holston et al. (2017) or Summers (2014). If long-run real

[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.

Table 1
Johansen trace tests for monetary policy cointegration.

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. Schwarz criterion for lag choice.

interest rates were characterised by a non-monotonic behaviour, this specification would not be reasonable. Still, to verify whether this formulation affects our results, we provide in the Supplemental Material File results for models where the Fisher coefficient is restricted to unity. We conclude that, overall, the qualitative impacts of the various shocks are similar to what obtains in the benchmark formulation.

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

$$\Delta X_t = \alpha_0 + \alpha_1 t + \gamma\beta X_{t-1} + \sum_{j=1}^{K} \beta_j \Delta X_{t-j} + u_t \tag{1}$$

where $X_t := (e_t, \pi_t, i_t^*, i_t, y_t)'$. γ and β 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. Notice that we include a trend in the specification. Although in theory this would not be needed and while it has no substantial impact on the estimated model, it improves the behaviour of the bootstrapped confidence bands of the impulse response functions. Focusing on the cointegration part (third term on the right-hand side):

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

where the superscript "F" in β^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) (Chapter 6) shows that matrix Ξ, 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 π_t), the rank of matrix Ξ is two. Given that matrix B is non-singular, the rank of matrix ΞB 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 ΞB 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.[3] This assumption, together with the stationarity of e_t, implies that the first, third and fourth

[3] This permanent shock only has permanent effects on inflation and on 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. We would obtain the same impulse response functions, only with a different labelling of the shocks. 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.

columns of ΞB are zeroes. The first row of ΞB 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 ΞB: (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 (zeroes in entries (2,5), (3,5) and (4,5)). 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). Given all this, matrix ΞB 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 will always leave unrestricted the first row of B (such that the effects of the shocks on exchange rates are not restricted in any way). In most specifications we also leave unrestricted the first column of B (such that the effects of exchange rate shocks are not restricted). Restrictions must then be placed in the other rows of the third and fourth columns of B. In our benchmark specification we want to be particularly agnostic; we assume the temporary monetary policy shock of the other economy has no contemporaneous impact on U.S. interest rates and U.S. output (zeroes in entries (4,3) and (5,3) of B). Given this, we need a restriction on the effects of the temporary U.S. monetary policy shock (fourth column of B); we assume it has no contemporaneous impact on U.S. output (a zero in entry (5,4)). This last restriction is less innocuous, even if it is in accordance with the usual view that monetary policy affects economic activity with a lag. The alternative in the fourth column would be to restrict the impact on U.S. inflation, which we leave as a robustness exercise, since restricting the impact on the foreign interest rate is not sufficient for identification. All this implies a matrix B 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. There are certainly many other identification possibilities through matrix B. As an example, to avoid these restrictions on the impact on U.S. output or U.S. inflation we can place zeroes in the first column, e.g. by restricting the contemporaneous response of interest rates to exchange rate shocks. This could be warranted if, despite the close attention central banks pay to exchange rate dynamics, one could argue that interest rate setting does not react contemporaneously (in the same month) and systematically to exchange rate surprises. In the Supplemental Material File we conduct a very thorough analysis considering this and several other identification assumptions.[4]

We should mention that a more rigorous statistical treatment of the data would call for a non-linear specification, on account of the lower bound constraint on nominal interest rates. The setup of Mavroeidis (2021) and the analysis in Rossi (2019) are useful in this regard. Such treatment would entail several additional choices (e.g., treat the lower bound as exogenous or endogenous, how to endogenize the lower bound, assume or not a different behaviour of the economy at the lower bound) and it would significantly complicate the identification of shocks in our setting that employs long-run restrictions. We are somewhat reassured by the sub-sample analysis included in the Supplemental Material File, that focuses on the pre-inflation targeting (and pre-lower bound) period, but we acknowledge there is a selection bias issue, see also Hayashi and Koeda (2019).

[4] We reduce somewhat the possible combinations by discarding unwarranted restrictions: again, we do not restrict in any way the effects of other shocks on exchange rates, since this is our main variable of interest, and we do not restrict the contemporaneous effects of country A interest rate shocks on interest rates of country A. Apart from these, we cover a wide array of combinations that are feasible in terms of achieving the identification of the shocks. We analyse thoroughly and across sub-samples a model that restricts the contemporaneous response of interest rates to exchange rate shocks, allowing us to leave unrestricted the impact response of U.S. output or U.S. inflation to the U.S. temporary monetary shock. Given the doubts this assumption could cast, we cover many other possibilities. We organise them along a few dimensions. We consider two subsets of restrictions. In the first subset, we assume that the temporary foreign monetary policy shock does not affect U.S. output contemporaneously. We also assume the U.S. monetary policy shock does not affect U.S. output on impact. We then consider the following alternatives: the temporary U.S. monetary policy shock does not affect contemporaneously U.S. inflation; the temporary monetary policy shock in the other country does not affect contemporaneously U.S. inflation, which is arguably more innocuous than assuming no contemporaneous effect of the U.S. temporary monetary shock on U.S. inflation. These specifications add to the one where the temporary monetary policy shock in the other country does not affect contemporaneously U.S. interest rates, which is exactly the benchmark specification described earlier. In the second subset of possibilities, we assume that the temporary foreign monetary policy shock does not affect U.S. inflation contemporaneously (instead of assuming that it does not affect U.S. output contemporaneously). We then consider the following alternatives: the temporary U.S. monetary policy shock does not affect contemporaneously U.S. inflation and U.S. output; the U.S. monetary shock does not affect contemporaneously the interest rate of the other country and U.S. output (a zero in the third and fifth rows of the fourth column of B); the U.S. monetary shock does not affect contemporaneously the interest rate of the other country and U.S. inflation (a zero in the second and third rows of the fourth column of B).

Table 2
Estimation of vector error correction model and cointegration parameters - Fisher relation.

	Fisher relation U.S.					
	β^F	γ_e	$\gamma_{\pi^{US}}$	γ_{i^*}	$\gamma_{i^{US}}$	$\gamma_{y^{US}}$
U.S. - G.B.	0.51***	−0.13	−0.02***	−0.01	0.03**	−0.01
(4 lags)	0.15	0.29	0.01	0.02	0.01	0.01
U.S. - DE	0.43***	−0.42	−0.02***	−0.01	0.02**	−0.04***
(2 lags)	0.13	0.37	0.01	0.01	0.01	0.01
U.S. - FR	0.41***	−0.46	−0.02**	−0.02	0.03**	−0.02
(4 lags)	0.14	0.32	0.01	0.01	0.01	0.01
U.S. - AU	0.34***	−0.14	−0.02***	−0.02	0.02*	−0.04***
(2 lags)	0.14	0.32	0.01	0.02	0.01	0.01
U.S. - CH	0.26**	−0.13	−0.02***	−0.02	0.02*	−0.05***
(2 lags)	0.13	0.44	0.01	0.01	0.01	0.01

Notes: γ_{var} corresponds to the adjustment parameter in the equation for variable "var". Superscript "US" was added to variables π, i and y 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 vector error correction model and cointegration parameters - Monetary policy cointegration.

	Monetary policy cointegration					
	β^{MP}	γ_e	$\gamma_{\pi^{US}}$	γ_{i^*}	$\gamma_{i^{US}}$	$\gamma_{y^{US}}$
U.S. - G.B.	0.84***	0.60**	0.00	−0.06***	0.00	−0.02
(4 lags)	0.14	0.25	0.01	0.01	0.01	0.01
U.S. - DE	0.33*	0.07	0.00	−0.02***	−0.01	−0.04***
(2 lags)	0.18	0.29	0.01	0.01	0.01	0.01
U.S. - FR	0.74***	−0.03	0.01	−0.06***	−0.02	−0.04***
(4 lags)	0.11	0.29	0.01	0.01	0.01	0.01
U.S. - AU	1.09***	0.00	0.00	−0.05***	0.00	−0.02**
(2 lags)	0.24	0.19	0.00	0.01	0.01	0.01
U.S. - CH	0.57***	0.50	0.00	−0.04***	−0.02**	−0.03***
(2 lags)	0.19	0.32	0.00	0.01	0.01	0.01

Notes: γ_{var} corresponds to the adjustment parameter in the equation for variable "var". Superscript "US" was added to variables π, i and y 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.

3. Empirical results

3.1. Cointegration

Tables 2 and 3 present the parameter estimates of the VEC model for the cointegration vector parameters (β) and the adjustment parameters (γ) for the models with G.B., DE, FR, AU and CH and the two cointegration relations assumed.[5] The analyses for JP and EA are done separately on account of the different specifications required, as discussed in Section 2. 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 0.26 and 0.51 in all the models, consistent with a slow fall in real interest rates over the past decades and with the results in Valle e Azevedo et al. (2022). Likewise, in all the models, both inflation and nominal interest 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. As for the cointegration between monetary policies, the parameters hover around one, but range from 0.33 to 1.09. Also, the nominal interest rate of the second economy reacts significantly to re-establish the long-run relation, as seen by the significance of γ_{i^*} in Table 3. The reaction of U.S. interest rates to re-establish the long-run relation is usually not significant, except in the model with CH. 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. Overall, these adjustment dynamics suggests a leading behaviour of the U.S. economy *vis-à-vis* the other economies in our sample. Gray (2013) presents empirical and theoretical evidence of this behaviour comparing the U.S. and 12 countries from 1980 to 2008 within a panel regression setting and Belke and Gros (2005) provide some evidence specifically between the FED and the European Central Bank.

[5] All the results in this paper regarding the estimation of the SVEC model are obtained by using JMulTi, see http://jmulti.de.

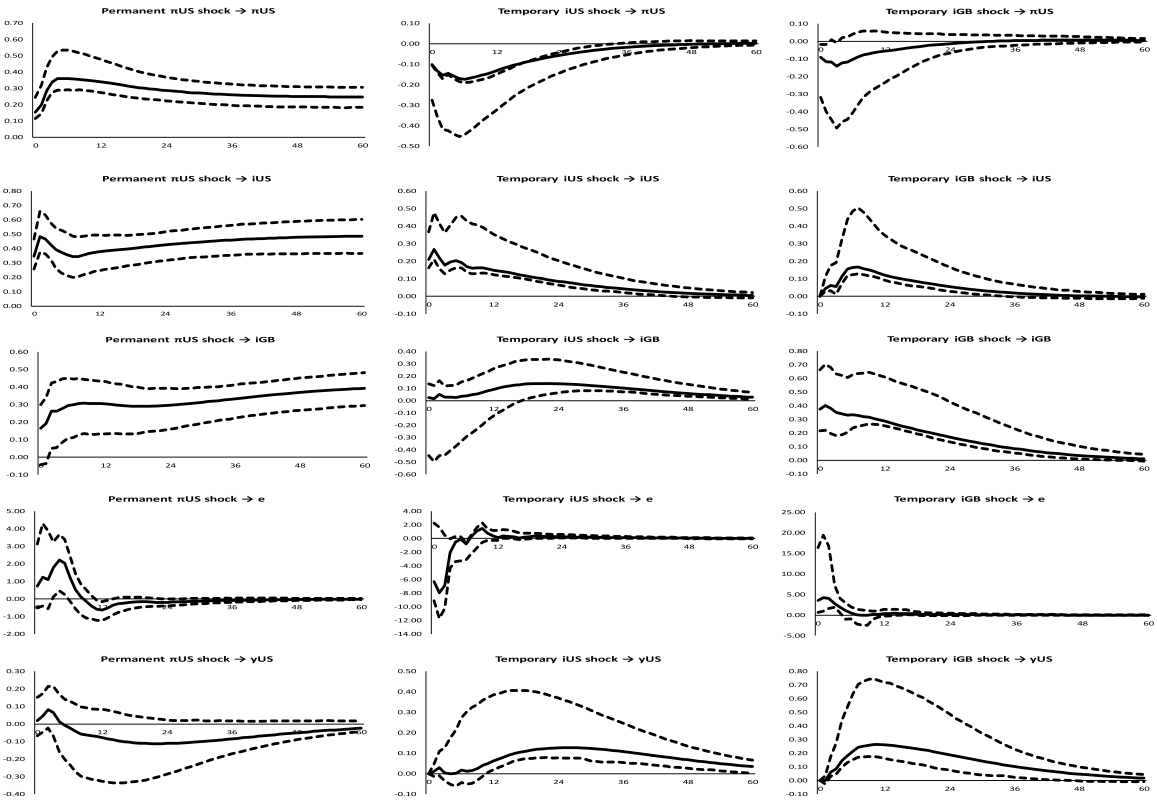

Fig. 1. Impulse response functions - U.S. - G.B. model. All responses are in percentage points. The first column reports the responses to the permanent U.S. monetary shock, the second column the response to the temporary monetary shock and the third column the response to G.B. temporary monetary shock. 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.

3.2. Impulse response functions

As regards the impulse response functions, we analyse here more comprehensively results for G.B.. For the other economies we focus on the effects of the permanent and temporary monetary policy shocks on e_t, the rate of change of the exchange rate. The complete results, figures and detailed comments for the models with DE, FR, AU and CH can be found in the Supplemental Material File. Fig. 1 displays 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. Fig. 2 summarises the results for the other countries, displaying the responses of e_t to the permanent and temporary U.S. monetary policy shocks and to the temporary monetary policy shock in the other economy. We should note that here and in the Supplemental Material File some impulse response functions lay outside the respective confidence intervals, which is bound to occur with bootstrapped confidence intervals and can be attributed to imprecise point estimates. This could warrant bias corrections but we avoided further arbitrary choices in this regard.

Fig. 1 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 Uribe (2022) and Valle e Azevedo et al. (2022): the permanent monetary 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 monetary 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 implied by the cointegration relation between the two nominal interest rates. The temporary U.S. monetary policy shock does not have a significant effect on G.B.'s nominal interest rates, a result also found in the model for AU. In the models for DE, FR and CH, the temporary U.S. monetary policy shock has a positive and significant impact on the interest rate of the other economy.

Focusing on the impacts of G.B.'s temporary monetary policy shock (associated with an increase in the nominal interest rate of that economy), it is found that the impact on U.S. inflation is not statistically significant at conventional levels, while the impact on U.S. interest rates is positive and statically significant (the same is true in the models with other countries, with the exception of AU). Looking at the impacts of the structural shocks on U.S. output, the permanent U.S. monetary policy shock has a positive but not

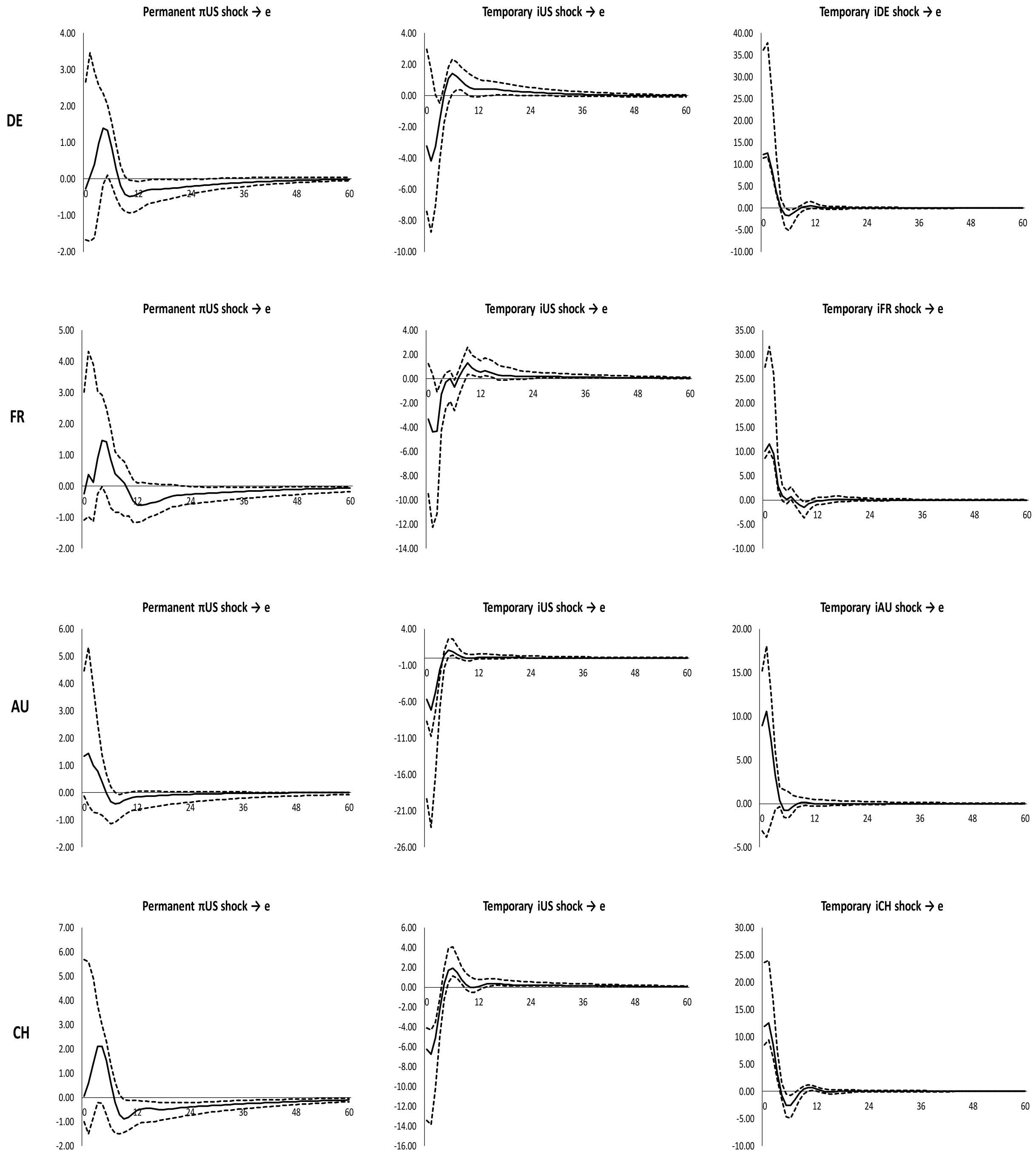

Fig. 2. Impulse response functions to e_t (rate of change of the exchange rate) - Model with other countries. All responses are in percentage points. The first column reports the responses to the permanent U.S. monetary shock, the second column the temporary U.S. monetary policy shock and the third column the response to the temporary monetary shock in the identified economy. 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. 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.

statistically significant impact in the model with G.B. (in the models with FR and CH the impact is negative but still not significant, see the Supplemental Material File), while the U.S. temporary monetary policy shock also seems to have a positive impact in the models for G.B., DE and CH (while for the other countries the impact is most often not significant). Considering the alternative identification schemes (see the Supplemental Material File) this non-standard result is not obtained in many instances. Still, and overall, the sign of the point estimates of the effects on output is not robust and the significance is usually low. This contrasts with the negative effect robustly found in Valle e Azevedo et al. (2022). Recall that a non-positive impact is assumed *a priori* in Uribe (2022).

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 Fig. 1 and in Fig. 2, 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. 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 (2022) 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 robust result 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 delayed by roughly six months (with the exception of the model with AU) and short-lived (less than one year). 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. While many responses of e_t (and of exchange rates themselves) are not clearly significant, we should underline that qualitatively, and across a host of alternative formulations and identification strategies (see also the Supplemental Material File), the results are quite robust.

3.3. Accumulated impulse response functions

Since we consider the rate of change in exchange rates in all the exercises, it is useful to look at the behaviour of (the log of) exchange rates themselves. In Fig. 3 we report accumulated IRFs, specifically the effects of the permanent and temporary U.S. monetary policy shocks on (the log of) exchange rates. 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 impacts on the exchange rate 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 but with similar dynamics.

Our results seem to provide new insights on the delayed overshooting puzzle, i.e., the fact that a monetary tightening is associated with a rather slow appreciation of the currency, followed by a gradual depreciation, see Eichenbaum and Evans (1995) for an early reference, and the confirmation with alternative identification strategies by, e.g., Scholl and Uhlig (2008). However, other identification strategies (see, e.g., Forni and Gambetti (2010)) and the consideration of the Volcker era (where the delay can be very large and contaminate larger samples, see Kim et al. (2017)) underline the contained generality of the delayed overshooting characterisation. In our setting there seems to be little delay, particularly in what regards the effects of the temporary monetary policy shock. The strong effects of temporary contractionary monetary shocks on the appreciation of the USD tend to be immediate (with peak effects usually around four to six months), followed by a slow depreciation. The effects of the permanent monetary shock are not as strong but have similar dynamics with the opposite sign. This suggests that ignoring the existence of the two types of monetary disturbances may contribute to the puzzle. In this vein, not controlling for permanent monetary shocks would be less problematic (and arguably not lead to puzzles) if one uses standard identification schemes in the post-Volcker – or successful inflation targeting – era, consistent with Kim et al. (2017). Finally, the fact that our setting helps reviving Dornbusch's overshooting does not mean that distinguishing between permanent and temporary shocks is sufficient to achieve this outcome; for a counterexample see Schmitt-Grohé and Uribe (2022), where the effects of the two types of monetary disturbances lack overshooting and even delayed overshooting.

3.4. Forecast error variance decomposition

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 Supplemental Material File). Clearly, the shock that accounts for the majority of the forecast error variance of e_t is the U.S. temporary monetary policy shock, explaining more than 40% over 60 months, followed by the shock on e_t itself. From the evidence of the remaining countries, some comments stand out: First, in contrast with the result for G.B., the shock that explains the majority of the forecast error variance of e_t is the temporary monetary policy shock in the other advanced economy (not the U.S. one) and this result is common across models. Second, the share is quite stable across forecast horizons. Third, the second most important shock in this dimension is the temporary U.S. monetary policy shock, explaining between (roughly) 10% and 30% across countries and horizons. Moreover, an important conclusion, which is robust across all models, including in the model with G.B., 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. (2022) –, appears to explain only a small fraction of the forecast error variance of e_t, thereby hinting that the behaviour of the USD is much more driven by the temporary monetary policy shocks rather than by the permanent monetary policy shock. These conclusions are not in line with the findings in Schmitt-Grohé and Uribe (2022), who report that the permanent monetary policy shocks account for the majority of the forecast error variance of the

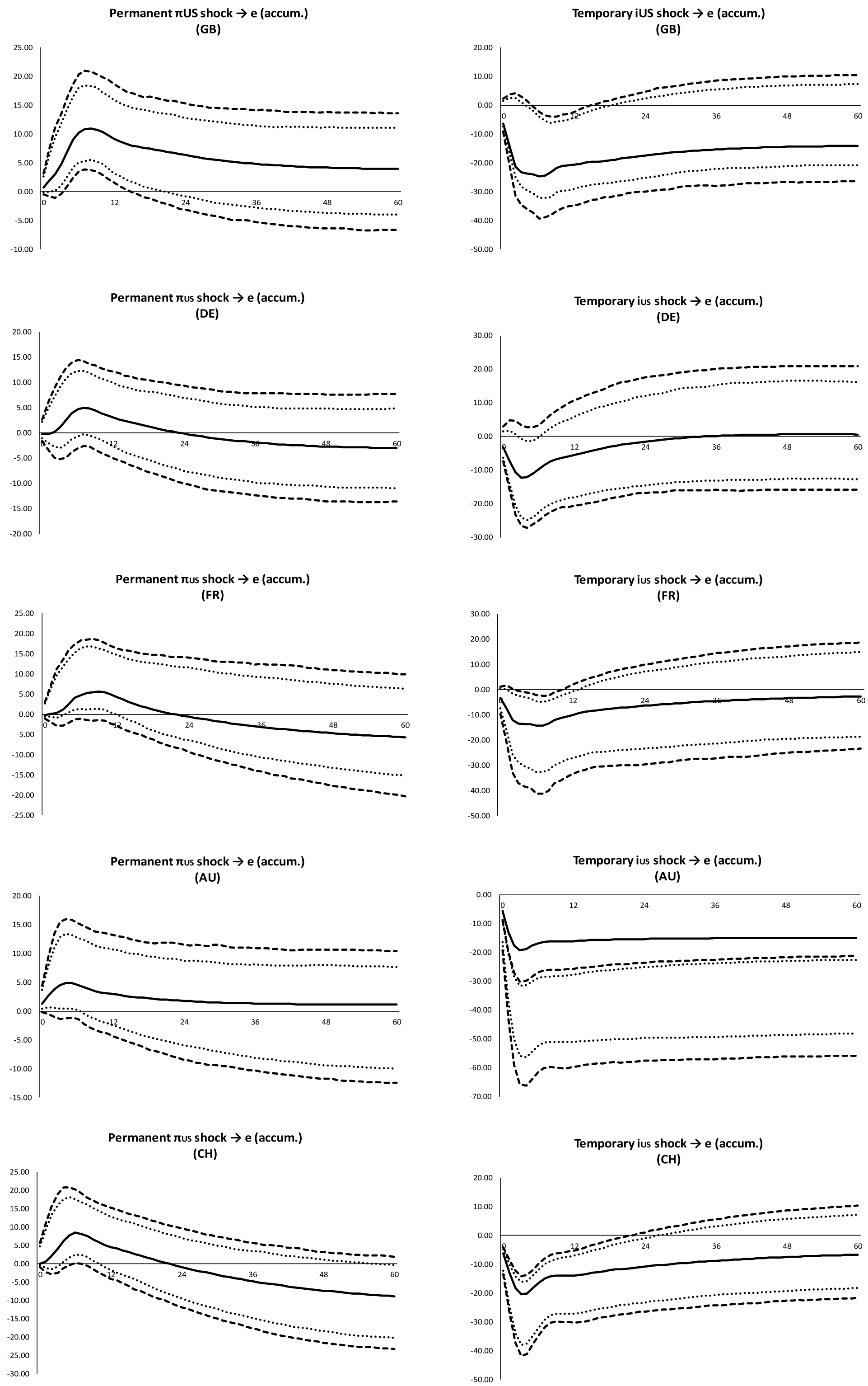

Fig. 3. Accumulated impulse response functions - Effects on e_t. All responses are in percentage points. The first column has the responses to the permanent U.S. monetary shock for the various countries and the second column the response to the temporary monetary shock. Shocks are identified as described 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% and 80% (pointed line) Hall Bootstrap confidence intervals. In one instance for AU, the IRF lays outside the confidence intervals, which can be attributed to imprecise point estimates. This is bound to occur with bootstrapped confidence intervals.

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

Var.	Shock	Horizon						
		1	2	4	12	24	48	60
	ε_e	0.36	0.34	0.30	0.32	0.31	0.31	0.31
	$\varepsilon_{\pi_{US}}$	0.01	0.01	0.02	0.05	0.05	0.05	0.05
$e - US/GB$	$\varepsilon_{i_{GB}}$	0.14	0.14	0.17	0.16	0.16	0.16	0.16
	$\varepsilon_{i_{US}}$	0.43	0.46	0.47	0.43	0.43	0.42	0.42
	$\varepsilon_{y_{US}}$	0.06	0.05	0.05	0.04	0.05	0.05	0.05
	ε_e	0.10	0.09	0.08	0.06	0.05	0.04	0.03
	$\varepsilon_{\pi_{US}}$	0.44	0.43	0.53	0.62	0.70	0.79	0.81
π_{US}	$\varepsilon_{i_{GB}}$	0.16	0.15	0.11	0.06	0.04	0.03	0.02
	$\varepsilon_{i_{US}}$	0.19	0.19	0.15	0.14	0.11	0.08	0.07
	$\varepsilon_{y_{US}}$	0.12	0.13	0.13	0.12	0.10	0.07	0.06
	ε_e	0.31	0.33	0.34	0.24	0.15	0.10	0.08
	$\varepsilon_{\pi_{US}}$	0.09	0.10	0.15	0.28	0.37	0.53	0.61
i_{GB}	$\varepsilon_{i_{GB}}$	0.49	0.47	0.43	0.43	0.39	0.37	0.22
	$\varepsilon_{i_{US}}$	0.00	0.00	0.00	0.01	0.04	0.06	0.05
	$\varepsilon_{y_{US}}$	0.11	0.10	0.08	0.04	0.04	0.05	0.04
	ε_e	0.03	0.02	0.02	0.03	0.02	0.01	0.01
	$\varepsilon_{\pi_{US}}$	0.70	0.73	0.76	0.69	0.75	0.86	0.89
i_{US}	$\varepsilon_{i_{GB}}$	0.00	0.00	0.01	0.07	0.06	0.03	0.02
	$\varepsilon_{i_{US}}$	0.26	0.24	0.20	0.17	0.12	0.07	0.05
	$\varepsilon_{y_{US}}$	0.01	0.00	0.01	0.04	0.05	0.03	0.02
	ε_e	0.17	0.15	0.11	0.04	0.02	0.01	0.01
	$\varepsilon_{\pi_{US}}$	0.00	0.01	0.01	0.00	0.01	0.01	0.01
y_{US}	$\varepsilon_{i_{GB}}$	0.00	0.00	0.01	0.05	0.04	0.02	0.02
	$\varepsilon_{i_{US}}$	0.00	0.00	0.00	0.00	0.01	0.01	0.01
	$\varepsilon_{y_{US}}$	0.83	0.84	0.86	0.90	0.93	0.96	0.97

exchange rate. A possible reason could be related to the assumption of non-stationarity of e_t in Schmitt-Grohé and Uribe (2022) (as opposed to treating this variable as stationary), together with the imposed cointegration between this variable and the permanent monetary policy shocks.

We tried to verify this conjecture and estimated models that assume a unit-root in the rate of change of the exchange rate and cointegration of this variable with the interest rates of the two economies. It turns out that with these two elements the share of the forecast error variance of e_t explained by the model increases somewhat (it becomes in general the second most relevant shock in terms of FEVD, after the shock to e_t itself). In the case of the U.S. - G.B. model the contribution of the permanent shock to the FEVD of e_t peaks at about 15% (and wanders around 15%–20% in most countries), but in the case of FR the contribution reaches 86%. In this setting the remaining results are little affected and we certainly confirm the robust result across specifications that permanent nominal shocks account for a large fraction of the FEVD in the case of nominal rates and inflation, exactly as in Schmitt-Grohé and Uribe (2022).

3.5. Robustness analysis

The analysis of the impact of the structural shocks on the variables of the model was repeated with the alternative identifications for the temporary monetary policy shocks referred earlier (for a complete description see the Supplemental Material File). The impulse response analysis shows that the results for the U.S. monetary policy shocks displayed in the baseline identification are very broadly robust to alternative identification strategies. 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. The impacts of the temporary monetary policy shock of the other economy on e_t also typically lead to an appreciation of the respective currency but insignificant and even opposite effects also occur across variations. There is one specification where the effect of the permanent U.S. monetary shock on the exchange rate is still the usual one, but the effect of the U.S. temporary monetary shock is not significant in several instances. It is denoted Alternative 4 in the Supplemental Material File and it restricts the contemporaneous impact of the temporary foreign monetary shock on U.S. inflation and U.S. output as well as the effect of the temporary U.S. monetary shock on U.S. output.

This whole 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 DE, FR and CH. 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; DE and FR from 1971M1 to 1999M1; AU from 1971M4 to 1996M9 and CH from 1974M1 to 1999M12. The results can be found in the Supplemental Material File. In general, the results are quite similar to those obtained in the full sample in terms of cointegration parameters. Still, we should underline that, in the models with DE and CH, the monetary policies cointegration coefficient is not statistically significant, which can be due to the (almost) stationary behaviour of their interest rates even in this sub-sample. The conclusions from the impulse response analysis

Table 5
Estimation of vector error correction model and cointegration parameters - Non-stationary sample.

	Non-stationary sample - Fisher relation U.S.					
	β^F	γ_e	$\gamma_{\pi^{US}}$	γ_{i^*}	$\gamma_{i^{US}}$	$\gamma_{y^{US}}$
U.S. - JP	0.53***	0.25	−0.03***	−0.01	0.04**	0.00
(4 lags)	0.14	0.50	0.01	0.01	0.02	0.01

Notes: γ_{var} corresponds to the adjustment parameter in the equation for variable "*var*". Superscript "US" was added to variables π, 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.

broadly confirm the results for the whole sample. In particular, the impacts of the monetary shocks 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 (and the foreign temporary monetary shock leads to a depreciation of the USD), while the permanent U.S. shock results in a USD depreciation (even if in the models for DE and CH the effect is not significant). The impact of the temporary U.S. monetary policy shock is in general large, contrasting with the milder impact of the permanent monetary policy shock.

4. No cointegration of 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 to our benchmark model. Also, the prolonged experience of JP in a low inflation and low interest rate environment makes it a relevant case to analyse. 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 is on the cointegration relations. Given the lack of statistical evidence on cointegration between interest rates in JP 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 as a non-stationary variable with no cointegration relations with the other variables. Given that the long-run restrictions depend on the cointegration relations, this change also implies a modification in those restrictions. The long-run identification matrix now has 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 ΞB that now allows Japan's monetary policy shock to have permanent effects only on the nominal interest rate of JP. This is a necessary assumption given the non-stationary behaviour of this variable and it implies that the monetary policy shock in JP will be considered a permanent shock. Regarding the short-run restrictions matrix, given that only two temporary shocks must be distinguished now, 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 temporary U.S. monetary policy shock has no immediate impact on U.S. output, in line with the identification scheme of the baseline model in Section 3.[6]

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. is statistically significant and both inflation and U.S. interest rates adjust to the long-run relation.

[6] Again, in the Supplemental Material File we consider several alternative identification assumptions in the case of Japan, including in the non-stationary sample. We consider versions where we restrict the contemporaneous response of U.S. interest rates to the exchange rate shock and another where we restrict the contemporaneous response of U.S. inflation to the U.S. temporary monetary shock. The main results described below are robust to these different specifications.

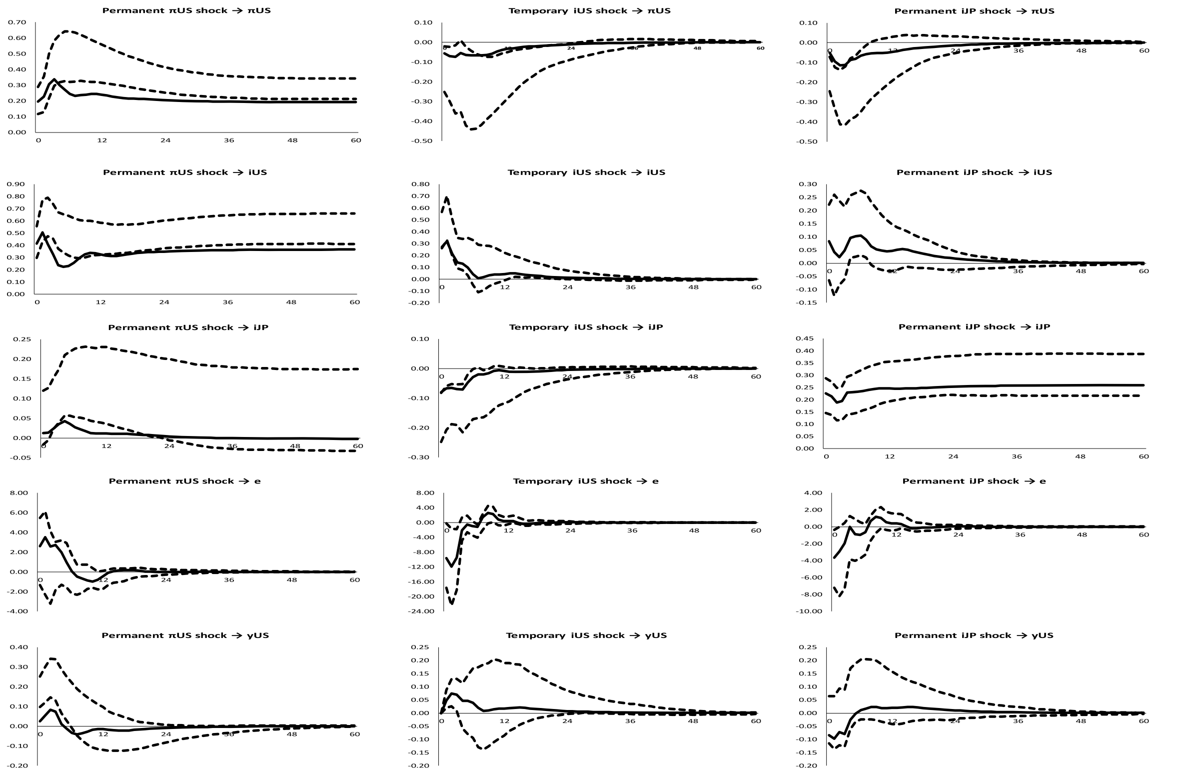

Fig. 4. Impulse response functions - U.S. - JP model - Non-stationary sample. All responses are in percentage points. The first column has the responses to the permanent U.S. monetary shock, the second column the response to the temporary monetary shock and the third column the response to JP temporary 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, which can be attributed to imprecise point estimates. This is bound to occur with bootstrapped confidence intervals.

Fig. 4 presents the estimated impulse response functions, which in this case are often imprecisely estimated. Still, and again, the permanent U.S. monetary policy shock increases permanently inflation and interest rates in the U.S., while the temporary monetary policy shock, 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. 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 a negative effect on U.S. inflation. As regards the effects on e_t (the rate of change of the exchange rate), the permanent U.S. monetary policy shock has a positive impact but it is not significant. The permanent monetary policy shock in Japan leads to an appreciation of the USD, consistent with the findings so far (that 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 on the differentiated impacts of temporary and permanent monetary policy shocks on exchange rates are robust to a setting not characterised by monetary policy cointegration.

Regarding the FEVD, the results are similar to those in the model with G.B., as documented in the Supplemental Material File: most of the forecast error variance of the rate of change of the exchange rate is explained by the U.S. temporary monetary policy shock, whereas the shock to e_t stands as the second most important. The permanent monetary policy shock in the U.S. continues to explain only a small share of the forecast error variance of e_t. 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 for most countries, the forecast error variance of interest rates in JP is now mostly explained by its own monetary policy shock. Recall that cointegration between interest rates is not imposed in this version of the model.

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 Japan's data. As mentioned before, in this period Japan's interest rate is well described as stationary, hence it is not cointegrated with the U.S. interest rate and it is not impacted or impacts any other variable permanently. To accommodate

Table 6
Estimation of vector error correction model and cointegration parameters - Stationary sample.

	Stationary - Fisher relation U.S.					
	β^F	γ_e	$\gamma_{\pi^{US}}$	γ_{i^*}	$\gamma_{i^{US}}$	$\gamma_{y^{US}}$
U.S. - JP	0.24***	−3.64	−0.09***	0.00	0.00	−0.16
(4 lags)	0.06	2.27	0.02	0.01	0.03	0.11

Notes: γ_{var} corresponds to the adjustment parameter in the equation for variable "*var*". Superscript "US" was added to variables π, 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.

these features, Japan's interest rate will be treated just like the rate of change of the exchange rate e_t. The model has thus three cointegration relations: the Fisher relation in the U.S., the cointegration relation to account for the stationary of e_t and another one for Japan's interest rate. These changes imply a different format 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 to the original model for JP, 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 assume, as in the original model, that the temporary Japanese monetary policy shock has no contemporaneous impact on U.S. interest rates and output, while the temporary U.S. monetary policy shock has no impact on U.S. output (i.e. matrix B is the same as in Section 3). Table 6 displays the estimated cointegration and loading coefficients. In this model, the cointegration coefficient for the Fisher relation is somewhat smaller compared to the previous sub-sample. Again, the only variable that adjusts to the long-run relation is U.S. inflation, in contrast to what occurs in the models with other countries.

Fig. 5 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 a negative impact on U.S. inflation and it increases Japan's interest rate (but the effect is not significant). As for the temporary Japanese monetary policy shock, associated with a temporary increase in Japan's interest rate (and also an increase in the U.S. interest rate), it has a negative impact on U.S. inflation. Finally, the impacts on e_t are exactly the ones found in the benchmark model for most economies. The permanent contractionary U.S. monetary policy shock leads to a depreciation of the USD, while the temporary shock results in the appreciation of the U.S. currency. The temporary Japanese monetary policy shock also leads to a depreciation of the USD. These findings continue to suggest that the conclusions of the original model remain valid in a more recent sub-sample with a stationary interest rate. Moreover, these results are robust to various configurations in terms of short-run identification assumptions. The Supplemental Material File also contains results for the additional identification restrictions considered.

This version of the model follows closely the results of the FEVD identified in the models with DE, FR, AU and CH (see the Supplemental Material File). The forecast error variance of e_t is mostly explained by the Japanese temporary monetary policy shock, even after three years, followed by the U.S. temporary monetary policy shock. The permanent U.S. monetary policy shock continues to explain a small fraction of the forecast error variance of e_t, as concluded with other models.

5. Extended symmetric model – Euro area

Our 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. We can extend the model to account for the distinction between temporary and permanent monetary policy shocks also in the second economy. Here, we expand the baseline model, using data for the U.S. and the euro area from 1971M1 to 2019M6, and using DE as a proxy to extend the sample before 1999, which could raise concerns considering the diverse exchange rate arrangements since 1971. Since 1979 the biggest euro area countries except Spain are under the Exchange Rate Mechanism (ERM), which limited the variability of exchange rates, even though several "realignments" did take place. Spain joined the ERM in 1989 while a systematic depreciation of the Italian lira was observed, with a temporary exit from the ERM between the end of 1992 and the end of 1996. One possibility would be to cut the sample in 1999. It turns out that this leads to serious numerical troubles in the estimation of the model (and inability to estimate it under standard settings). We tried to extend the sample backwards from 1999 until we were confident that the algorithms converged, but unfortunately this required going back until the mid-1980s, with results very similar to those reported using the full sample.

We employ a fully symmetric SVEC model with seven variables: the five variables in the original model plus inflation and industrial production in the euro area. In the original model the two cointegration relations are the Fisher relation in the U.S. and the cointegration of monetary policies (interest rates) of the U.S. and the other economy. Since we include the euro area

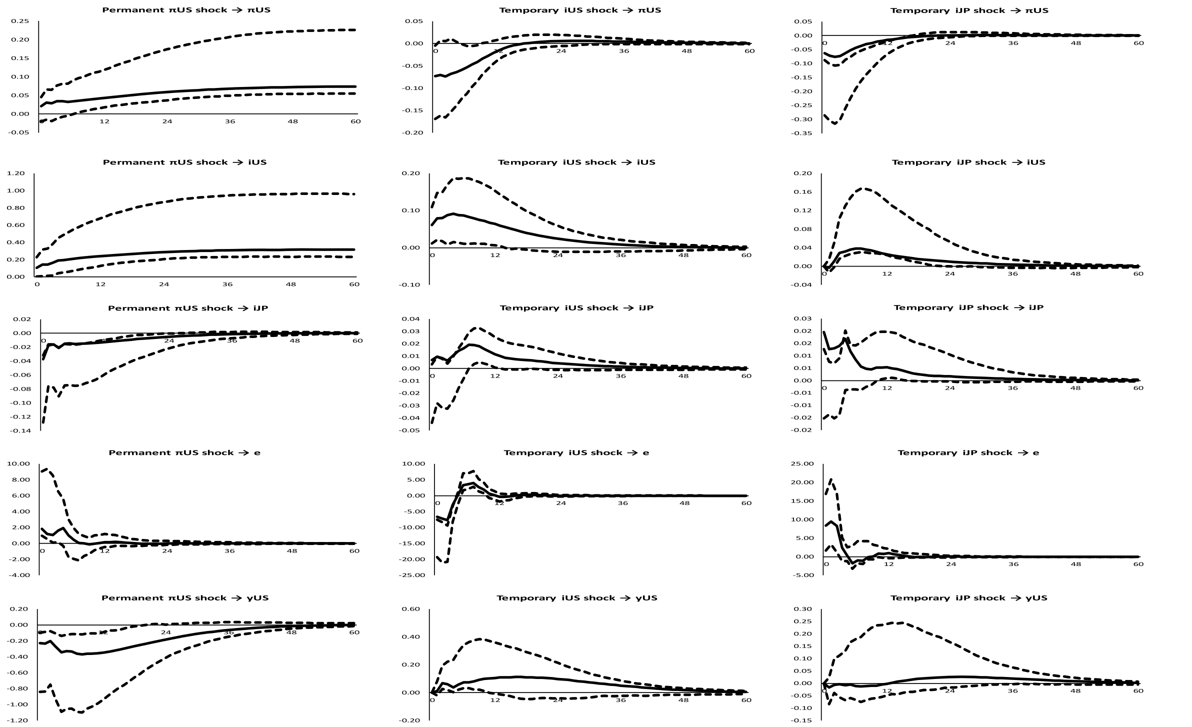

Fig. 5. Impulse response functions - U.S. - JP model - stationary sample. All responses are in percentage points. The first column has the responses to the permanent U.S. monetary shock, the second column the response to the temporary monetary shock and the third column the response to JP temporary monetary 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.

inflation rate we can also consider the Fisher relation in the euro area as another cointegration relation. This feature is important to distinguish between temporary and permanent monetary policy shocks in the euro area. Additionally, the cointegration between monetary policies is not assumed *a priori* in order to simplify the identification and to allow for the distinction between permanent and temporary monetary policy shocks in the two economies. A version of the model with monetary policy cointegration has been considered and the main results do not change substantially.

To sum up, 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 benchmark model). 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 impose a symmetric structure to the imposed restrictions, while keeping features of the identification of the original model, using both short and long-run restrictions. The long-run identification matrix ΞB is as follows:

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

This specification means 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 ΞB). 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. Next, just as in the benchmark model, the structural shocks to output are assumed to have only long-run impacts on the output of the respective economy (asterisks in columns 6 and 7 of matrix ΞB).

Table 7
Estimation of vector error correction model and cointegration parameters - Fisher relation U.S.

	Fisher relation U.S.							
	$\beta^{F_{US}}$	γ_e	$\gamma_{\pi^{US}}$	$\gamma_{\pi^{EA}}$	$\gamma_{i^{US}}$	$\gamma_{i^{EA}}$	$\gamma_{y^{US}}$	$\gamma_{y^{EA}}$
U.S. - EA	0.79***	−0.56	−0.02**	0.03***	0.03**	−0.02**	−0.05***	−0.01
(2 lags)	0.09	0.43	0.01	0.01	0.01	0.01	0.02	0.04

Notes: γ_{var} corresponds to the adjustment parameter in the equation for variable "var". Superscript "US" or "EA" was added to variables π, 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 - Fisher relation EA.

	Fisher relation EA							
	$\beta^{F_{EA}}$	γ_e	$\gamma_{\pi^{US}}$	$\gamma_{\pi^{EA}}$	$\gamma_{i^{US}}$	$\gamma_{i^{EA}}$	$\gamma_{y^{US}}$	$\gamma_{y^{EA}}$
U.S. - EA	0.49***	−0.92	−0.01	−0.09***	−0.02	0.02	0.05	0.01
(2 lags)	0.05	0.71	0.01	0.01	0.02	0.02	0.03	0.07

Notes: γ_{var} corresponds to the adjustment parameter in the equation for variable "var". Superscript "US" or "EA" was added to variables π, 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.

The short-run identification B matrix has the following representation:

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

It is thus assumed that the permanent monetary policy shocks do not affect contemporaneously the interest rate of the other economy (the zeroes in columns 2 and 3 of matrix B). Overall, the consideration of the two economies makes more innocuous the identification as restrictions can be placed on the effects of shocks in the other economy. In this vein, to complete the identification we simply assume that the temporary monetary policy shocks in both economies have no contemporaneous impact on the interest rates and inflation of the other economy (zeroes in the fourth and fifth columns of matrix B). The Supplemental Material File contains results for an alternative identification strategy, where we leave unrestricted the effects of interest rates on the other economy's interest rates and we restrict instead the contemporaneous response of interest rates in the U.S. and the euro area to exchange rate shocks. The main results are very in much in line with those provided here and we underline the exceptions below. 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.

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 below one. In the euro area, this parameter stands slightly lower at 0.49. The adjustment parameters in the case of the U.S. suggest 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 significantly. Other estimates for the adjustment parameters are also statistically significant at the conventional levels, particularly 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.

Fig. 6 displays the IRFs from this 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 euro area permanent and temporary monetary policy shocks. Permanent monetary policy shocks have the already familiar effects on each region's inflation and interest rates: the permanent monetary policy shock has a positive impact on both interest rates and inflation, also in the short run. The permanent U.S. monetary policy shock also has a positive impact on the nominal variables in the euro area (suggesting some correlation of monetary policies), while the permanent euro area shock has an immediate positive impact on U.S. interest rate and a negative impact on U.S. inflation. These results are qualitatively robust to the alternative identification provided in the Supplemental Material File. The temporary U.S. monetary policy shock, associated with an increase of U.S. interest rates, has now a non-significant impact on inflation (negative with the alternative identification), while the temporary euro area monetary policy shock does have a negative impact on euro area inflation. Regarding the impacts of the temporary shocks on the euro area economy, there are several non-significant impacts or significant impacts that are not robust across specifications (see the Supplemental Material File). As for the impact of the shocks on e_t, the impacts of the U.S. monetary policy shocks (both permanent and temporary) are similar to those obtained thus far: the temporary U.S. monetary policy shock leads to an appreciation of the USD, while the permanent U.S. monetary policy shock has the

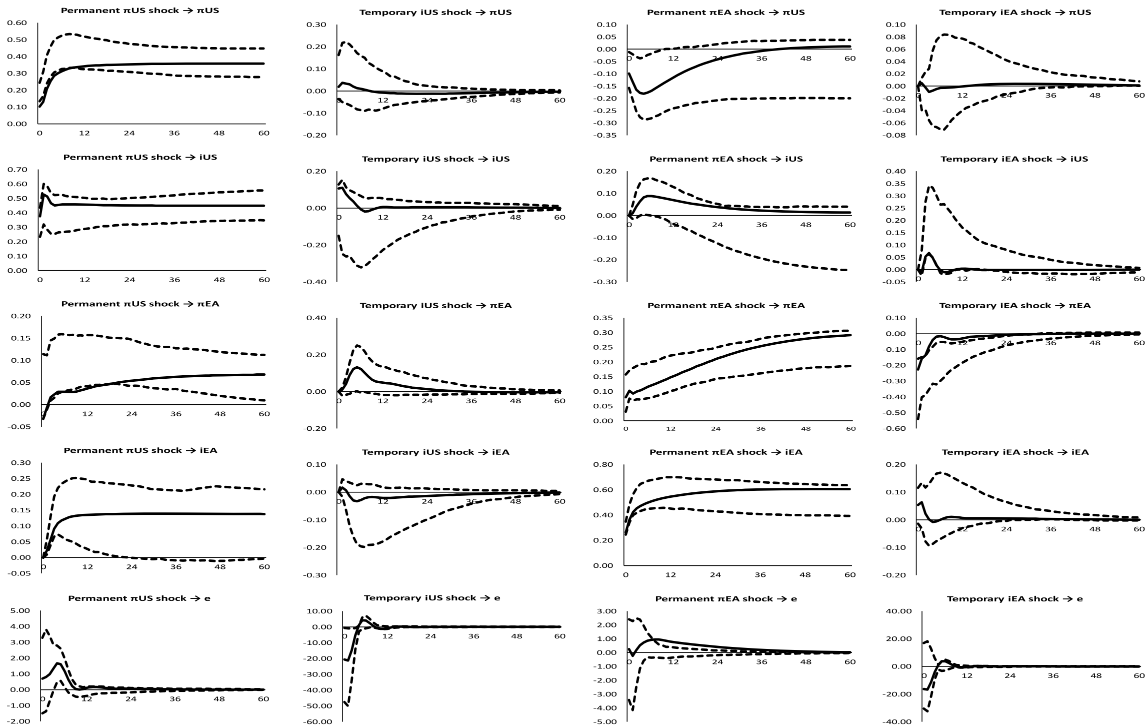

Fig. 6. Impulse response functions - U.S. - EA model. All responses are in percentage points. The first column has the responses to the permanent U.S. monetary shock, the second column the response to the temporary U.S. monetary shock, the third column the response to the EA permanent shock and the fourth 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.

opposite effect, resulting in an depreciation of the USD. The impacts of the euro area monetary policy shocks on the USD are not statistically significant (and in the alternative specification, despite non-significant results the point estimates indicate a negative impact of the euro area temporary monetary policy shock on e_t, i.e., an appreciation of the USD which is by now counter-intuitive).

Looking at the FEVD (see the Supplemental Material File), the forecast error variance of e_t is strongly explained by the temporary monetary policy shocks (together they account for 84% of that variance after three years), while the permanent monetary policy shocks explains only a residual amount (4% after three years), in line with the results for other countries. 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.

6. Concluding remarks

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. These results fit well in 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 – and with theoretical arguments and empirical literature that finds exchange rate movements consistent with interest rare parity conditions. 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. This suggests that monetary policy ought to consider the way it perturbs the long-run level of nominal rates. Otherwise, it risks obtaining opposite effects to those intended.

Our results also provide new insights on the delayed overshooting puzzle. We find strong and immediate effects of temporary contractionary monetary shocks on the appreciation of the USD, which helps reviving Dornbusch's overshooting. This suggests that ignoring the existence of permanent monetary disturbances (characterised by somewhat delayed and opposite effects), especially when considering samples with long-run variability in nominal rates and inflation, may contribute to the puzzle, consistent with Kim et al. (2017).

Table 9
Description and sources of the variables used 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	π_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.

We confirm in a flexible setting the results of Schmitt-Grohé and Uribe (2022), who first explored the distinction between permanent and temporary nominal shocks in explaining exchange rate dynamics. For identification we take advantage of two cointegration relations that characterise the data and we do not need to impose the standard effects of temporary monetary policy shocks. We perform various robustness checks and analyse several departures from this standard setting, dictated by the properties of the data (e.g., the varying long-run variability in Japan's data depending on the sample) or by the desirability of a more symmetric treatment of the economies (e.g. when we consider the U.S. and the euro area). The main results are remarkably stable across different specifications and identifying assumptions.

Declaration of competing interest

The authors declare that they have no known competing financial interests or personal relationships that could have appeared to influence the work reported in this paper.

Data availability

https://data.mendeley.com/datasets/6yfvs3dswb/1.

Appendix A. 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. 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.

Appendix B. Supplementary data

Supplementary material related to this article can be found online at https://doi.org/10.1016/j.jinteco.2023.103871.

References

Amisano, G., Giannini, C., 1997. Topics in Structural VAR Econometrics, second ed. Springer, Berlin.
Arouri, M., Jawadi, F., Nguyen, N., 2013. What can we tell about monetary policy synchronization and interdependence over the 2007–2009 global financial crisis? J. Macroecon. 36, 175–187.
Belke, A., Cui, Y., 2010. US–euro area monetary policy interdependence: New evidence from taylor rule-based VECMs. World Econ. 33 (5), 778–797.
Belke, A., Gros, D., 2005. Asymmetries in transatlantic monetary policymaking:Does the ECB follow the fed? J. Common Mark. Stud. 43 (5), 921–946.

Bjornland, H.C., 2008. Monetary policy and exchange rate interactions in a small open economy. Scand. J. Econ. 110 (1), 197–221.
De Michelis, A., Iacoviello, M., 2016. Raising an inflation target: The Japanese experience with Abenomics. Eur. Econ. Rev. 88, 67–87.
Dornbusch, R., 1976. Expectations and exchage rate dynamics. J. Polit. Econ. 84 (6), 1161–1176.
Eichenbaum, M., Evans, C., 1995. Some empirical evidence on the effects of shocks to monetary policy on exchange rates. Q. J. Econ. 110 (4), 975–1009.
Fama, E.F., 1984. Forward and spot exchange rates. J. Monetary Econ. 14 (3), 319–338.
Forni, M., Gambetti, L., 2010. The dynamic effects of monetary policy: A structural factor model approach. J. Monetary Econ. 57 (2), 203–216.
Froot, K., Frankel, J., 1989. Forward discount bias: Is it an exchange risk premium? Q. J. Econ. 104 (1), 139–161.
Gordon, R., 2014. The Demise of US Economic Growth: Restatement, Rebuttal and Reflections. NBER Working Paper No. 19895.
Gray, C., 2013. Responding to a monetary superpower: Investigating the behavioral spillovers of U.S. monetary policy. Atl. Econ. J. 41 (2), 173–184.
Hayashi, F., Koeda, J., 2019. Exiting from quantitative easing. Quant. Econ. 10, 1069–1107.
Holston, K., Laubach, T., Williams, J.C., 2017. Measuring the natural rate of interest: International trends and determinants. J. Int. Econ. (108), 59–75.
Kim, S., Moon, S., Velasco, C., 2017. Delayed overshooting: Is it an '80s puzzle? J. Polit. Econ. 125 (5), 1570–1598.
Kim, S., Roubini, N., 2000. Exchange rate anomalies in the industrial countries: A solution with a structural VAR approach. J. Monetary Econ. 45 (3), 561–586.
Lothian, J.R., 2016. Uncovered interest parity: The long and the short of it. J. Empir. Financ. 36 (1), 1–7.
Lütkepohl, H., 2006. New Introduction To Multiple Time Series Analysis. Springer, Berlin and Heidelberg.
Mavroeidis, S., 2021. Identification at the zero lower bound. Econometrica 89 (6), 2855–2885.
Miller, N.C., 2014. Exchange Rate Economics: The Uncovered Interest Parity Puzzle and Other Anomalies. Edward Elgar Publishing.
Rossi, B., 2019. Identifying and Estimating the Effects of Unconventional Monetary Policy: How to Do It and What Have We Learned?. Discussion Paper DP14064, CEPR.
Schmitt-Grohé, S., Uribe, M., 2022. The effects of permanent monetary shocks on exchange rates and uncovered interest rate differentials. J. Int. Econ. (135).
Scholl, A., Uhlig, H., 2008. New evidence on the puzzles: Results from agnostic identification on monetary policy and exchange rates. J. Int. Econ. 76 (1), 1–13.
Summers, L., 2014. Secular stagnation, hysteresis and the zero lower bound. Bus. Econom. (49), 65–73.
Uribe, M., 2022. The neo-Fisher effect: Econometric evidence from empirical and optimizing models. Am. Econ. J.: Macroecon. 14 (3), 133–162.
Valle e Azevedo, J., Ritto, J., Teles, P., 2022. The neutrality of nominal rates: How long is the long run? Internat. Econom. Rev. 63 (4), 1745–1777.
Zettelmeyer, J., 2004. The impact of monetary policy on the exchange rate: evidence from three small open economies. J. Monetary Econ. 51 (3), 635–652.