Effects of monetary policy on the $\$/£$ exchange rate. Is there a 'delayed overshooting puzzle'?

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Abstract

The determination of the $\$/£$ exchange rate is studied in a small symmetric macroeconometric model including UK–US differentials in inflation, output gap, short and long-term interest rates for the four decades since the breakdown of Bretton Woods. The key question addressed is the possible presence of a'delayed overshooting puzzle' in the dynamic reaction of the exchange rate to monetary policy shocks. In contrast to the existing literature, we follow a data-driven modelling approach combining (i) a VAR based cointegration analysis with (ii) a graph-theoretic search for instantaneous causal relations and (iii) an automatic general-to-specific approach for the selection of a congruent parsimonious structural vector equilibrium correction model. We find that the long-run properties of the system are characterized by four cointegration relations and one stochastic trend, which is identified as the long-term interest rate differential and that appears to be driven by long-term inflation expectations as in the Fisher hypothesis. It cointegrates with the inflation differential to a stationary 'real' long-term rate differential and also drives the exchange rate. The short-run dynamics are characterized by a direct link from the short-term to the long-term interest rate differential. Jumps in the exchange rate after short-term interest rate variations are only significant at 10%. Overall, we find strong evidence for delayed overshooting and violations of UIP in response to monetary policy shocks.

Keywords: Exchange Rates; Monetary Policy; Cointegration; Structural VAR; Model Selection.

JEL classification: C22; C32; C50.

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1 Introduction

The determination of the exchange rate has frequently been the focus of the contributions of Giancarlo Gandolfo to macroeconomics. In his empirical work on exchange rates, Gandolfo repeatedly emphasized the importance of a system approach to exchange modelling (see, inter alia, Gandolfo, 1979, Gandolfo, 1981, Gandolfo, Padoan and Paladino, 1990b, and Gandolfo, Padoan and Paladino, 1990a) and delivered with his Italian Continuous Time Model a powerful framework for doing so (see, inter alia, Gandolfo and Padoan, 1990). In the tradition of Gandolfo, we develop in this paper a small economy-wide macroeconometric model for the $\$/£$ exchange rate. Using quarterly data from $1972Q1$ to $2009Q2$ and imposing symmetry, the system consists of five country differences between the UK and the US (indicated by *): the inflation differential $\pi_t^d = \pi_t - \pi_t^*$, the output gap differential, $y_t^d = y_t - y_t^*$, the 3-month interest rate spread $i_t^d = i_t - i_t^*$, the 10-year government bond yield spread $r_t^d = r_t - r_t^*$, and the nominal exchange rate, e_t, itself.

Of particular interest to this paper is the dynamic reaction of the exchange rate to monetary policy shocks in the form of a variation of the short-term interest rate differential. While in the standard overshooting models of Dornbusch (1976) and Frankel (1979), see Rogoff (2002) for a survey, the exchange rate jumps instantaneously in response to an interest rate shock in order to depreciate over time and thereby restoring the uncovered interest rate parity (UIP), there is a growing body of empirical evidence suggesting that exchange rates, rather than adjusting instantly as predicted by the theory, appreciate steadily for several months before finally depreciating. Whether or not such a 'delayed overshooting puzzle' is present in the case of the $\$/£$ exchange rate is the key question this paper seeks to contribute to.

Vector autoregressive (VAR) models have long served as the workhorse for studying the empirical reaction of exchange rates to monetary policy shocks.[1] In the seminal paper of Eichenbaum and Evans (1995), the effects of US monetary policy shocks on five exchange rates were analyzed in a VAR framework with Choleski-type causal ordering. Three different measures of shocks were considered: shocks to the federal funds rate, shocks to the ratio of non-borrowed to total reserves and changes in the Romer and Romer (2004) monetary policy index. For the period from 1974M1 to 1990M5,

[1]There has been some criticism in the literature about the limited information set of a small-scale VAR approach. For example, Mumtaz and Surico (2009) applied a factor augmented VAR with the UK as the domestic country and 17 other industrialized countries as the foreign block. For the period 1974Q1 to 2005Q1, they find no delayed overshooting.

Eichenbaum and Evans (1995) found the considered exchange rates to appreciate for several months after an expansionary US monetary policy shock until reaching a peak from which they then start to decline in value. The detected delay in overshooting was 2 to 3 years, with Japan having the shortest and the UK the longest delay. Pronouncedly shorter delay estimates were produced by Grilli and Roubini (1995, 1996), who discussed the delayed overshooting puzzle within the framework of a 'liquidity model' where, in contrast to sticky price models, goods prices are flexible while asset markets only adjusts gradually.[2]

Following up on the Eichenbaum and Evans (1995) approach, recent contributions including Cushman and Zha (1997), Faust and Rogers (2003), Kim (2005) and Scholl and Uhlig (2008) have all used vector autoregressions with superimposed exclusion, sign or shape identification restrictions usually derived from economic theory to overcome the ad-hoc nature of recursive orderings in a Choleski approach. Commencing from a small open-economy assumption, Cushman and Zha (1997) considered a structural VAR model with imposed block exogeneity, such that the non-domestic block of US variables were not affected by domestic Canadian variables. Allowing the CAD-USD exchange rate to react contemporaneously to (domestic) monetary policy shocks via an information market equation, no puzzles were found for the period 1974 to 1993. Also Kim and Roubini (2000) found no delayed overshooting for non-US G-7 exchange rates from 1974M7 to 1992M12, when identifying the contemporaneous effects with zero restrictions derived from economic theory nonrecursively. These conflicting empirical results were underpinned by Faust and Rogers (2003), who demonstrated the delayed overshooting result can be sensitive to questionable assumptions, such that the peak appreciation could be within one month after the monetary policy shock when allowing for simultaneity. Seeking to avoid 'dubious identifying assumptions', Faust and Rogers (2003) identified the VAR only partly, but used informal restrictions to calculate the impulse responses following the approach in Faust (1998). 7 and 14-variable models of the US-UK and US-German bilateral exchange rate from 1974M1 to 1997M12 showed that monetary policy shocks, while not the main source of exchange rate variability,

[2] Some recent papers have revived the interest in finding an explanation for the delayed-overshooting phenomenon. According to Gourinchas and Tornell (2004), the puzzle is caused by systematic distortion in investors' beliefs about the interest rate process. Suppose investors overestimate the relative importance of transitory interest rate shocks. Confronted with a higher than expected interest rate in the next period, investors revise their beliefs. This 'updating effect' has been suggested as a cause of the forward premium effect and the delayed overshooting puzzle. Kim (2005) proposed that foreign exchange rate interventions of the central bank as driving factors of the delayed overshooting puzzle for the Canadian-US bilateral exchange rate. Exchange rate appreciation on impact might be counteracted by policy interventions in the foreign exchange market.

generate large UIP deviations.

The effects of monetary policy on exchange rates have recently been revisited by Scholl and Uhlig (2008) using an identification procedure with sign restrictions. Analyzing bilateral exchange rate data from 1975M07 to 2002M07 for the US versus Germany, the UK, Japan and the G7, respectively, they found evidence for delayed overshooting. The delay in the response of the $\$/\pounds$ exchange rate was with 17 months the shortest. Even when the possibility of delayed overshooting was excluded by construction, a 'sizeable' positive forward premium remained. It was shown that these deviations from UIP can be exploited by hedging strategies with Sharpe ratios greater than those in equity markets. Combining short and long-run restrictions, i.e., allowing for simultaneity between interest rates and exchange rate, but assuming no long-run effects of monetary policy on exchange rates, Bjørnland (2009) rejected a delayed overshooting puzzle for the real exchange rates of Australia, Canada, New Zealand and Sweden with the US in the period from 1983Q1 to 2004Q4. Finally, using an identification method exploiting breaks in the heteroscedasticity of the structural innovations, Bouakez and Normandin (2010) obtained a delay of about 10 months for US-G7 bilateral exchange rates.[3]

This paper seeks to contribute to the knowledge on the delayed overshooting puzzle by improving on the existing literature in three economically and econometrically important aspects:

(i) Despite the involvement of possibly integrated time series, most of the relevant literature is based on VAR models in levels. By commencing from an unrestricted cointegrated VAR model and developing a parsimonious structural vector equilibrium correction model, which is the adequate $I(0)$ representation of the system, we will be able to carefully study the long-run and short-run properties of the macroeconomic time series under consideration. An econometric model with a well-defined long-run equilibrium imposes important data-coherent constraints on impulse responses functions, which are critical when assessing the effects of macroeconomic stabilization policies.

(ii) The overwhelming part of the existing literature uses unrestricted VAR or just-identified structural VAR models for the analysis of exchange rate responses to monetary policy shocks. Such highly parameterized VAR models require the estimation of a waste number of parameters and

[3] The omitting of multilateral spillover effects was criticized by Binder, Chen and Zhang (2010), who proposed a Global VAR model for the analysis of the effects of US monetary policy shocks. For a sample from 1978 to 2006, no delayed overshooting was found.

suffer from the curse of dimensionality: as the degrees of freedom are being exhausted and estimation uncertainty is inflated with a growing number of variables or lags, so do the impulse responses functions become inconclusive due to a growing width of confidence intervals, which will eventually include the zero line. To avoid this problem we will make use of the breakthrough in automatic general-to-specific model reduction procedures in reducing the complexity of the model while preserving the characteristics of the data.

(iii) Since the original contribution of Eichenbaum and Evans (1995), there has been an intense discussion about the arbitrary assumptions leading to the identification of the direction of instantaneous causality. In Eichenbaum and Evans's orthogonalization of the system, for example, it is assumed that a fed funds shock affects the currency but not the money markets: while the exchange rate can jump immediately in response to the shock, the short-term market interest rate differential is forced to remain unchanged contemporaneously. However, also many of the proposed alternative schemes are based on theoretical ad-hoc assumptions. In this paper, we seek to overcome these limitations by taking advantage of recent advances in graph theory and its application to the search for causality among variables.

It is worth noting that in contrast to the overwhelming part of the existing literature, this investigation into the presence of a 'delayed overshooting puzzle' in the response of the $\$/\pounds$ exchange rate to an asymmetric monetary policy shock in the UK and the US will follow the premise to let the data speak. Our data-driven modelling approach will combine (i) the VAR based cointegration analysis of Johansen (1995) and Juselius (2006) with (ii) the graph-theoretic approach of Spirtes, Glymour and Scheines (2001) for the search for instantaneous causal relations (see Demiralp and Hoover, 2003, for its application to econometrics) and (iii) the automatic general-to-specific approach of Krolzig and Hendry (2001) and Krolzig (2003) for the selection of a congruent parsimonious structural vector equilibrium correction model.

The structure of the paper is as follows. In $\S2$ we introduce the data set and show the time series. The $\S3$ discusses the econometric methodology and empirical model construction. $\S4$ investigates the effects of a monetary policy shock with focus on the presence of a delayed overshooting puzzle and violations of UIP. Finally $\S5$ concludes.

2 International parity conditions and the UK-US macro history since the breakdown of Bretton Woods

2.1 Time Series

The small macroeconometric model to be developed for the determination of the $\$/\pounds$ exchange rate will reflect the macro history of the US and UK over the last four decades. More precisely, we are using quarterly data from $1972Q1$ to $2009Q2$, involving a total of 150 quarterly observations. The paper is written from the UK perspective, so we will refer to UK variables as the domestic ones and US variables as foreign ones, marked by a star. Table 1 gives an overview over the macro time series under consideration. Analysing quarterly rather than commonly used monthly time series allows us to work with GDP-based measures of inflation and the output gap.

Insert Table 1 about here

Rather than modelling the system with the nine variables listed in Table 1, which would be quite demanding from the point of cointegration analysis, we reduce the dimension of the model and thus its complexity by imposing symmetry, i.e., we critically assume that the exchange rate depends only on the differences between countries. In other words, a shock in the UK has the same effect as a shock in the US of opposite sign, but of the same size. The system to be analyzed consists of five country differences between the UK and the US (indicated by *): the inflation differential, $\pi_t^d = \pi_t - \pi_t^*$, the 3-month interest rate spread, $i_t^d = i_t - i_t^*$, the 10-year government bond yield spread, $r_t^d = r_t - r_t^*$, the output gap differential, $y_t^d = y_t - y_t^*$, and the nominal exchange rate, e_t, itself. To guarantee the consistency of the parity conditions to be considered in §2.2, variables have been transformed to ensure that each interest rate, bond yield, rate of inflation and currency movement is measured as quarterly log return. Table 2 explains in detail how each variable entering the model has been created:

Insert Table 2 about here

2.2 Discussion

In the following, we discuss the properties of the macro time series as far as they are relevant for the econometric modelling to follow in §3.

2.2.1 Inflation and the output gap

Focussing first on the real economy, Figure 1 plots the inflation rates and output gaps in the UK and the US (left panel), as well their differences (right panel). It can be seen that, except for the most recent years, the UK macro economy is characterized by a far more volatile output gap and a higher rate of inflation.

Insert Figure 1 about here

Augmented Dickey Fuller tests suggest that the output gap differential, $y_t^d = y_t - y_t^*$, and the inflation differential, $\pi_t^d = \pi_t - \pi_t^*$, are stationary. While the stationarity of the output gaps should be ensured by construction, the unit root test results for the inflation differential are rather surprising given the diverging experiences in 1970s. However, the volatility of inflation rates in both countries as well as their differential decline during the Great Moderation, which might have affected the unit root test.

Moving to the asset markets, the further discussion is structured along some of the central international parity conditions (see Gandolfo, 2002, for an excellent overview).

2.2.2 Expectations hypothesis of the term structure

In the expectations model of the term structure, the yield of a zero bond with a maturity of T periods equals the mean of the expected one-period interest rates plus a potential risk premium, ϕ_t:

$$r_t = \frac{1}{T} \sum_{j=0}^{T-1} \mathsf{E}_t i_{t+j} + \phi_t. \tag{1}$$

If the short-term interest rate and the risk premium are stationary processes, it follows from (1) that the spread between i_t and r_t is also stationary, $r_t - i_t \sim I(0)$.

Insert Figure 2 about here

The term spreads for the UK and the US and their difference are plotted in Figure 2. While the term spread appears potentially stationary for the US, this clearly is not the case for the UK term spread. These conjectures were confirmed by ADF tests. We therefore should not expect that short and long-term interest rate differentials cointegrate.

2.2.3 Nominal interest rate parity

Figure 3 looks at the potential cointegration between the nominal short- and long-term interest rates in the UK and the US. Due to the accommodating UK monetary policy in the 1970s, the long-term interest-rate differential shows clear signs of non-stationarity. As the UK short-term interest rates do not fully reflect the inflation problem of that time period, the short-term interest-rate differential conversely is a potential candidate for a cointegration relation, though this was not confirmed by a univariate ADF test.

Insert Figure 3 about here

2.2.4 The Fisher hypothesis and the real interest rate parity

Another important relation for our empirical analysis is the Fisher hypothesis. It states that the nominal interest rate equals the real interest rate ρ_t, invariant to monetary policy, plus inflation expectations,

$$i_t = \rho_t + \mathsf{E}_t \pi_{t+1}, \tag{2}$$

where the real interest rate is determined by the marginal product of capital and thus expected to be stationary with a low variance.

The Fisher relation motivates the real interest rate parity, according to which the ex-ante real interest rates in home and foreign country should equalize in the long run, i.e.:

$$\rho_t - \rho_t^* = (i_t - \mathsf{E}_t \pi_{t+1}) - (i_t^* - \mathsf{E}_t \pi_{t+1}^*) \sim I(0). \tag{3}$$

Theoretically, the calculation of ex-ante real interest rate involves future inflation expectations. Due to the severe measurement problems, we focus here on a naive definition of the real interest rate using the current backward-looking inflation rate.[4] As can be seen in Figure 4, both the short-term and the long-term real interest rates for the UK and the US show a level shift in 1981. Since then a downward trend is present. Overall, the real long-term interest rate differential is more likely to be stationary than

[4] A common alternative measurement approach would involve the use of realized future inflation rates based on the rational expectations hypothesis, which excludes systematic forecast errors of the agents. This procedure is, however, not compatible with the VAR modelling approach used in this paper.

the real short-term differential.

Insert Figure 4 about here

2.2.5 Purchasing power parity

It might have come as a surprise to some readers that we included in our analysis the differences in inflation rates between the UK and US but not the relative price level. In light of the purchasing power parity (PPP) theory, one would have expected that the nominal exchange rate to follow the relative price level of the two countries. Thus, the real exchange rate $s_t = e_t + p_t - p_t^*$, which measures the deviation of the nominal exchange rate from the relative price level, should be mean-reverting, such that the law of one price holds at least in the long term.

However, as can be seen in the upper panels of Figure 5, purchasing power parity clearly does not hold for the $\$/\pounds$ exchange rate over the sample period. The Pound Sterling appreciated in real terms by more than 70% from the end of 1984 to the beginning of 2008. The secular upward trend of the real exchange rate can not be captured by a linear time trend. A possible trend stationarity of the real exchange rate is not supported by the ADF test, which does not reject the null of a unit root at the 5% significance level. In our judgement, the non-stationarity of the real exchange rate can not be explained within the set of macro variables considered here (see also footnote 7). We therefore leave this issue for further investigations.

Insert Figure 5 about here

2.2.6 Uncovered interest parity

A central parity condition is the uncovered interest rate parity (UIP), which requires that the expected return on the domestic asset is, in equilibrium, equal to the expected return, as measured in the home currency, on a foreign asset with otherwise identical characteristics. For a one-period bond, this implies:

$$i_t = i_t^* - \mathsf{E}_t \Delta e_{t+1}. \tag{4}$$

Under rational expectations, there are no systematic forecast errors and equation (4) can be rewritten as:

$$\xi_t = i_t^d + \Delta e_{t+1}, \tag{5}$$

where ξ_t is a martingale difference sequence measuring the excess return of the UK bond. The realized excess returns over the sample period and their cumulation can be seen in the central panels of 5.

The UIP condition in (4) has been formulated for a one-period bond, but can be generalized to bonds with multi-period maturities. Recall that according to the expectations hypothesis of the term structure, we have that the long-term interest rate, or more precisely the yield of a zero bond of maturity of T periods, equalizes the expected average return of one-period bonds over T periods:

$$r_t^d = \frac{1}{T}\sum_{j=0}^{T-1} \mathsf{E}_t i_{t+j}^d. \tag{6}$$

Combining (6) with the forward solution of the UIP relation in (4) for e_t,

$$e_t = \mathsf{E}_t e_{t+T} + \sum_{j=0}^{T-1} \mathsf{E}_t i_{t+j}^d, \tag{7}$$

we get the multi-period form of UIP,

$$e_t = \mathsf{E}_t e_{t+T} + T(r_t - r_t^*), \tag{8}$$

which states that the exchange rate is determined by the long-term exchange rate expectation, $\mathsf{E}_t e_{t+T}$, and long-term interest rate differential, $r_t - r_t^*$.

3 Econometric modelling

In contrast to the existing literature, we follow a data-driven modelling approach that combines the VAR based cointegration analysis of Johansen (1995) and Juselius (2006) with the graph-theoretic approach of Spirtes *et al.* (2001) implemented in TETRAD for the search for instantaneous causal relations and the automatic general-to-specific model selection algorithm implemented in *PcGets* of Krolzig and Hendry (2001) for the selection of a congruent parsimonious structural vector equilibrium correction model.

3.1 Methodology

The General-to-specific approach implemented in this paper follows the modelling approach of Krolzig (2003) and consists of the following four stages (see Hoover, Demiralp and Perez, 2009, for a related approach):

(i) *Specification of the general unrestricted system.*

We commence from a reduced-form vector autoregressive (VAR) model of order p and dimension K, without any equation-specific restrictions, to capture the characteristics of the data:

$$\mathbf{y}_t = \boldsymbol{\nu} + \sum_{j=1}^{p} \mathbf{A}_j \mathbf{y}_{t-j} + \boldsymbol{\varepsilon}_t, \tag{9}$$

where $\boldsymbol{\varepsilon}_t \sim \mathsf{NID}(\mathbf{0}, \boldsymbol{\Sigma})$ is a Gaussian white noise process. This step involves the specification of the deterministic terms, selection of the lag length, p, and misspecification tests to check the validity of the assumptions made.

(ii) *Johansen cointegration tests* and *identification of the cointegration vectors.*

The Johansen procedure for determining the cointegration rank, r, is then applied to the VAR in (9) mapped to its vector equilibrium-correction mechanism (VECM) representation:

$$\Delta\mathbf{y}_t = \boldsymbol{\nu} + \boldsymbol{\alpha}\boldsymbol{\beta}'\mathbf{y}_{t-1} + \sum_{j=1}^{p-1} \boldsymbol{\Gamma}_j \Delta\mathbf{y}_{t-j} + \boldsymbol{\varepsilon}_t, \tag{10}$$

where the reduced-rank matrix, $\boldsymbol{\Pi} = \boldsymbol{\alpha}\boldsymbol{\beta}'$ has been decomposed into a $K \times r$ dimensional loading matrix, $\boldsymbol{\alpha}$, and cointegration matrix, $\boldsymbol{\beta}$, containing the information of the long-run structure of the model. For the identification of economically meaningful overidentified cointegration relations, $\boldsymbol{\beta}^r$, we follow the empirical modelling procedure of Juselius (2006).

(iii) *Graph-theoretic search for instantaneous causal relations.*

The determination of the contemporaneous causal links between the variables has been advanced by modern graph-theoretic methods of searching for causal structure based on relations of conditional independence developed by computer scientists (Pearl, 2000) and philosophers (Spirtes *et al.*, 2001). Following Demiralp and Hoover (2003), who introduced this approach to econometrics, we use the PC algorithm implemented in TETRAD 4 (see Spirtes, Scheines, Ramsey and Glymour, 2005 for details). The PC algorithm exploits the information embedded in the residual variance-covariance matrix, $\hat{\mathbf{\Sigma}}$, of the system in (10). A causal structure is represented by a graph with arrows from causes to caused variables. To detect the directed acyclic graph, the algorithm starts by assuming that all variables are linked to each other through an undirected link. In the elimination stage, connections are first removed between variables which are unconditionally uncorrelated. Then connections are eliminated for variables which are uncorrelated conditional on other variables. Having identified the skeleton of the graph, the algorithm seeks to orient the undirected edges by logical reasoning. This involves the analysis of indirect connections by taking into account the whole graph, considering every pair of variables, exploiting already directed edges and the acyclicality condition. If all edges can be oriented, a directed acyclic graph (DAG) results.

Having identified the contemporaneous causal structure of the system, the VECM in (10) can be represented as a *recursive* structural vector equilibrium correction mechanism (SVECM). By suitable ordering of the variables of the system, the DAG can be mapped to a lower-triangular contemporaneous matrix, $\mathbf{B}^r$, with units on the diagonal and non-zero lower-off-diagonal elements representing the causal links found by the PC algorithm. In contrast to a traditional orthogonalisation with the help of a Choleski decomposition of $\hat{\mathbf{\Sigma}}$, this approach results in an overidentified SVECM in the majority of cases. The zero lower-triangular elements of $\mathbf{B}^r$ provide testable overidentifying constraints allowing to verify the validity of the selected contemporaneous structure. Most importantly, as the contemporaneous causal structure captured by $\mathbf{B}^r$ is data determined, it avoids the problems associated with the ad-hoc nature of orthogonalised structural VAR models.

Next we will consider *Gets* reductions of the SVECM to reduce the complexity of the model. However, note that if the PC algorithm finds a link with insufficient information to identify the direction of causality, an undirected edge emerge. In this case, there exists a set of contemporaneous causal structures, $\{\mathbf{B}^{(i)}\}$, that are all consistent with the data evidence. For the selection of $\mathbf{B}^r$ an additional modelling stage is required: After applying the model reduction step in (iv) to each SVECM associated with one of the contemporaneous causal structures found, the dominant econometric model is finally selected in step (v).

(iv) *Single-equation reductions of the recursive SVECM.*

Starting point is the structural VECM with long-run relations $\boldsymbol{\beta}^r$ determined by stage (ii) and contemporaneous structure $\mathbf{B}^r$ given by the corresponding directed acyclic graph found in (iii):

$$\mathbf{B}^r \Delta \mathbf{y}_t = \boldsymbol{\delta} + \tilde{\boldsymbol{\alpha}} \left(\boldsymbol{\beta}^{r\prime} \mathbf{y}_{t-1} \right) + \sum_{j=1}^{p-1} \boldsymbol{\Upsilon}_j \Delta \mathbf{y}_{t-j} + \boldsymbol{\eta}_t, \quad \boldsymbol{\eta}_t \sim \mathsf{NID}(\mathbf{0}, \boldsymbol{\Omega}), \tag{11}$$

where $\boldsymbol{\Omega}$ is a diagonal variance-covariance matrix. A single-equation based *Gets* reduction procedure such as *PcGets* can be applied to the equations in (11) straightforwardly and, as shown in Krolzig (2001), without a loss in efficiency. The parameters of interest are the coefficients collected in the intercept, $\boldsymbol{\delta}$, the adjustment matrix $\tilde{\boldsymbol{\alpha}}$ and the short-run matrices $\boldsymbol{\Gamma}_j$ in the SVECM. The result is a parsimonious structural vector equilibrium correction model denoted PSVECM, which is nested in (11) and defined by the selected $\boldsymbol{\delta}^*$, $\tilde{\boldsymbol{\alpha}}^*$ and $\boldsymbol{\Upsilon}_j^*$ with $j = 1, \ldots, p-1$.

(v) *Selection of the dominant PSVECM.*

If the graph-theoretical search in (iii) produces an acyclic graph with at least one undirected edge, the determination of the direction of instantaneous causal relations has to rely on the information from the PSVECMs resulting from the *Gets* reduction of the SVECMs as defined by the set of contemporaneous causal structures. As the PSVECMs are mutually non-nested and the union is usually unidentified, we propose to select the PSVECM with the greatest penalized likelihood.

3.2 Empirical findings

In the following we seek to develop a congruent and parsimonious statistical model for the macro dynamics involving the inflation differential, $\pi_t^d = \pi_t - \pi_t^*$, the output gap differential, $y_t^d = y_t - y_t^*$, the short-term interest rate differential, $i_t^d = i_t - i_t^*$, the long-term interest rate differential, $r_t^d = r_t - r_t^*$, and the exchange rate e_t.[5] The results of Augmented Dickey Fuller tests indicate that the output gap differential y_t^d and the inflation differential π_t^d are stationary, the other time series were found to be $I(1)$. Thus, the vector process, $\mathbf{y}_t = (\pi_t^d, y_t^d, i_t^d, r_t^d, e_t)'$ is integrated of order one: $\mathbf{y}_t \sim I(1)$.

3.2.1 Cointegrated vector autoregression

As discussed, the first step involves the specification of the deterministic terms, selection of the lag length and misspecification tests to check the validity of the assumptions made. The lag structure analysis of the unrestricted VAR, commencing from a maximum lag length of five with consecutive F-tests for excluded individual and joint lags, indicates a lag order of four. An unrestricted constant is included as the only deterministic term. A linear time trend was found to be statistically insignificant. The residuals are sufficiently well behaved to proceed with this system.

We continue by analyzing the long-run properties of the system. The number of stable long-run relations $\boldsymbol{\beta}'\mathbf{y}_t$, which is equal to the rank of the matrix $\mathbf{\Pi}$ of the VECM in (10), is determined by the Johansen (1995) test for $I(1)$ cointegration. The eigenvalues and trace test results are shown in Table 3. We find that the long-run properties of the system are characterized by four cointegration relations. With dimension $K = 5$ and rank $r = \mathrm{rank}(\mathbf{\Pi}) = 4$ there is one unit root in the system.

Insert Table 3 about here

In Table 4 we test for long-run weak exogeneity of the variables of the system. Under the null hypothesis of a particular zero row in $\boldsymbol{\alpha}$, the corresponding variable is not adjusting towards the long-run equilibrium. The LR test results show that, with a p-value of 0.93, the bond yield differential is the only weakly exogenous variable. Thus, we identified the long-term interest rate differential r_t^d as the unique common stochastic trend in the system.[6]

Insert Table 4 about here

Following the modelling approach by Juselius (2006), the following cointegration vectors could be identified based on their statistical acceptability and economic consistency.[7]

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[5]The discussion paper version of the paper discusses the validity of symmetry assumption by analysing the cointegration relationships of two two-county subsystems and and an impulse response analysis in a 6D system with the UK and the US short-term interest rates rather than their differential.

[6]In the following, we will see that the long-term interest rate differential appears to be driven by long-term inflation expectations as predicted in Fisher hypothesis.

[7]While there exists a stable long-run relation between the *nominal* exchange rate and the *nominal* interest rate differ-

(i) *Stationary output gap differential.*

$$y_t^d = y_t - y_t^* \sim I(0). \tag{12}$$

The first cointegration vector is the difference between the UK and US output gaps. Stationarity is expected here due to the very definition of the output gap.

(ii) *Stationary nominal short-term interest rate differential.*

$$i^d = i_t - i_t^* \sim I(0). \tag{13}$$

This is somewhat surprising that the *long-term* interest rate differential is nonstationary and constitutes the stochastic trend of the system. In other words, while the nominal interest rate parity holds for the money markets, it is violated for the bond markets. The opposite holds for the real interest rate parity:

(iii) *Stationary real long-term interest rate differential.*

$$\rho_t^d = r_t^d - \pi_t^d = (r - \pi)_t - (r^* - \pi^*)_t \sim I(0). \tag{14}$$

The third cointegrating vector reflects the real interest rate parity and is closely related to the Fisher hypothesis, where the real long-term interest rates are calculated naively with the current rather than the expected future inflation. Since r_t^d is nonstationary this must also hold for the inflation differential, which is driven by the same stochastic trend. It is also worth noting that, due to (13) and (14), the UK and US term structures do not cointegrate.

(iv) *Nominal long-term interest-rate differential based exchange rate determination.*

ential, no such relationship can be found for a system consisting of four real variables, the real short-term interest rate differential, $i_t^d - \pi_t^d$, the real long-term rate differential, $r_t^d - \pi_t^d$, the output gap differential, y_t^d, and the real exchange rate in form of deviations from PPP, $e_t + p_t^d$, as well as the inflation differential as the only nominal variable. The Johansen trace test indicates a cointegration rank of three rather than four in Table 3, where the cointegration vectors coincide with relations (12) to (14) passing the LR test at 36%. At a marginal significance level of 31%, the real exchange is one of two stochastic trends. The real exchange rate can also not be recovered as a cointegration relation by adding a linear trend to the cointegration space.

The last cointegration vector is a UIP inspired exchange rate determination relation:

$$e_t - 26.4(r - r^*)_t \sim I(0). \tag{15}$$

This cointegration vector should be interpreted in light of the multi-period form of UIP. For zero bonds with a maturity of 10 years, respectively $T = 40$ quarters, the formula in (8) results in:

$$e_t = \mathsf{E}_t e_{t+40} + 40 r_t^d. \tag{16}$$

While, for the type of government bonds analyzed here, the relation above only holds approximately, the estimated multiplier of 26.4 with a 2σ interval of $[11.76, 41.06]$ is consistent with the theory. Furthermore, with sample averages of 9.0 and 7.4 of the yield of 10-year government bonds of the UK and the US, the average duration is only 6.8 and 7.2 years, respectively. Thus, actually, the point estimate of 26.4 is very close to the predicted values of 27.2 and 28.7. According to (15) and (16), the long-term equilibrium movement in the foreign exchange rate can be traced back to the non-stationary long-term interest rate differential, exhibiting long swings, and stable long-term exchange rate expectations.

The system estimation results for the four cointegration vectors and their interaction with the variables of the system are shown in Table 5.

Insert Table 5 about here

The three over-identifying restrictions on the cointegration space are accepted by the LR test with a statistic of $\chi^2(3) = 3.80$ and a p-value of 0.28. The only unrestricted $\boldsymbol{\beta}$-coefficient is precisely estimated. In contrast, only few $\boldsymbol{\alpha}$-coefficients are statistically different from zero. Altogether we find that the long-term interest rate differential, r_t^d, is of central importance to the system. It constitutes the common stochastic trend, it cointegrates with the inflation differential π_t^d to a stationary 'real' long-term rate differential, and it also drives the exchange rate $e_t = 26.4 r_t^d$, which is consistent with UIP and stable long-term exchange rate expectations $\mathsf{E}_t e_{t+40}$. The output gap y_t^d and the short-term rate differential i_t^d are both self error correcting and weakly exogenous to the other cointegration relations. The time series of the four cointegrating relations can be seen in the upper right panel of Figure 3 as well as the lower right panels of Figures 1, 4 and 5.

3.2.2 Identifying instantaneous causality

The graph-theoretic identification of the contemporaneous causation in the VECM(3) with the identified cointegration vectors in (12) - (15) is based on the residual correlation matrix reported in Table 6. The only statistically significant contemporaneous correlation of shocks is between the short and long-term interest rates, $\rho_{ir} = 0.52$. Thus, in the very short term, the term structure is the strongest link between the macroeconomic variables. As the dominant force in transmitting and absorbing macroeconomic shocks, it will play an important role in the transmission of monetary shocks to the exchange rate. In contrast there is no instantaneous causality between the exchange rate and the other variables of the system. This indicates that the exchange rate does not jump in response to an interest rate shock, implying a delayed overshooting.

Insert Table 6 about here

For further investigation, the correlation matrix in Table 6 is subjected to a graph-theoretical search for instantaneous causal relations. At the 5% and 1% significance levels, the PC algorithm consistently finds only a single undirected edge linking the short and long-term interest rate differentials, coinciding with the only significant residual correlation. Thus, both directions of the instantaneous causality need to be considered at the next modelling stage.

Following the same ordering of the variables as throughout the paper, the two candidate designs for the contemporaneous matrix are:

$$\mathbf{B}^{(1)} = \begin{pmatrix} 1 & . & . & . & . \\ 0 & 1 & . & . & . \\ 0 & 0 & 1 & . & . \\ 0 & 0 & b_{ri} & 1 & . \\ 0 & 0 & 0 & 0 & 1 \end{pmatrix} \quad \text{and} \quad \mathbf{B}^{(2)} = \begin{pmatrix} 1 & . & . & . & . \\ 0 & 1 & . & . & . \\ 0 & 0 & 1 & b_{ir} & . \\ 0 & 0 & . & 1 & . \\ 0 & 0 & 0 & 0 & 1 \end{pmatrix}, \tag{17}$$

where zeros indicate over-identifying restrictions. Note that, while the ordering of the variables is not unique, our choice does not affect the further analysis. The first design, $B^{(1)}$, represents the case where the short-term interest rate differential, i_t^d, instantaneously causes the bond yield differential, r_t^d. In the second design, $B^{(2)}$, the causality is inverted to r_t^d causing i_t^d.

In order to determine the direction of the instantaneous causality, we followed the procedure laid out in $\S 3.1$. Since the PSVECMs emerging from the *Gets* model reduction for the two competing designs of the contemporaneous causality matrix are mutually non-nested, the contemporaneous causality structure associated with the PSVECM with the highest penalized likelihood is being selected. According to both the Akaike and Schwarz information criterion, $\mathbf{B}^{(1)}$ is the dominant design. Hence, it is the short-term interest rate differential, i_t^d, that instantaneously causes the bond yield differential, r_t^d. The following discussion is based on the selected $\mathbf{B}^{(1)}$ matrix.[8]

3.2.3 The parsimonious structural vector equilibrium correction model

Having specified the SVECM in (11) with the cointegration relations found in $\S 3.2.1$ and the contemporaneous relations detected by the PC causal search algorithm in $\S 3.2.2$, the model reduction is performed with an automatic general-to-specific model selection procedure with regard to the parameters of the short-run dynamics, $\mathbf{\Gamma}_1, \ldots, \mathbf{\Gamma}_3$, and the long-run equilibrium adjustment, $\boldsymbol{\alpha}$, while ensuring that the rank of the long-run matrix $\mathbf{\Pi}$ is unchanged by the constraints on $\boldsymbol{\alpha}$. As shown in Krolzig (2003), when commencing from a structural VECM with known causal order and diagonal variance-covariance matrix, all possible reductions of the SVECM can be efficiently estimated by OLS and model selection procedures can operate equation-by-equation without a loss in efficiency. The conservative strategy of *PcGets* used here approximates in large samples the Schwarz information criteria (for more about mapping information criteria to significance levels see Campos, Hendry and Krolzig, 2003). The properties of automatic *Gets* selection are discussed in more detail in Hendry and Krolzig (2005).

The final parsimonious model selected by *PcGets* and estimated with OLS is as follows: All coefficients are significant with a t-value of at least 2. The adjusted R^2 of the reduced single equations are close to 30% for the short-term interest rate, output gap and exchange rate equations and approximately 65% for the inflation rate and the bond rate equations. Major outliers are corrected by including impulse dummies.

We start with the law-of-motion of the inflation differential, which falls into the class of augmented Phillips curves with the output gap as measure of macroeconomic activity and the long-term interest

[8] The results for the PSVECM with $\mathbf{B}^{(2)}$ can be requested from the authors.

rate as measure of inflation expectations:

$$
\begin{aligned}
\widehat{\Delta\pi_t^d} \;=\;& \underset{(0.098)}{0.746}\ (r_{t-1}^d - \pi_{t-1}^d) + \underset{(0.039)}{0.17}\ y_{t-1}^d - \underset{(0.085)}{0.264}\ \Delta\pi_{t-1}^d - \underset{(0.064)}{0.186}\ \Delta\pi_{t-2}^d \\
& - \underset{(0.32)}{0.795}\ \Delta r_{t-1}^d - \underset{(0.008)}{0.044}\ \text{I1973Q2}_t + \underset{(0.008)}{0.0424}\ \text{I1979Q3}_t, \\
& \hat{\sigma} = 0.00772,\ \bar{R}^2 = 0.65.
\end{aligned} \tag{18}
$$

The speed of adjustment of the inflation differential toward the cointegrating real interest rate differential is with 75% per quarter very high. This suggests that the long-term interest differential is a good proxy of differences in inflation expectations in the UK and the US, $\pi_t^{e,d} = \pi_t^e - \pi^{*e}_t$, as predicted by Irving Fisher, such that $r_t^d - \pi_t^d \approx \pi_t^{e,d} - \pi_t^d$. This should, however, give rise to a price puzzle (see Demiralp, Hoover and Perez, 2009, for a detailed graph-theoretical analysis of the US price puzzle). The Phillips-curve interpretation is supported by the positive feedback from the output gap to inflation. The short-run dynamics are characterized by a self-stabilizing feedback in the quarterly changes in the inflation differential as well as changes in the bond rate. Note that short-term interest rates are not affecting inflation directly. Thus, for monetary policy to be effective in controlling inflation, short-term interest rate changes need to affect the long-term rates or the output gap.

The output-gap equation is error correcting. There is a strong and surprisingly positive response to interest rate changes, which is consistent with a forward-looking monetary policy rule. This output gap puzzle, in combination with the Phillips-curve term in the inflation equation, amplifies the price puzzle phenomenon. Also a weak short-run Mundell effect can be found:

$$
\begin{aligned}
\widehat{\Delta y_t^d} \;=\;& -\underset{(0.039)}{0.156}\ y_{t-1}^d + \underset{(0.051)}{0.184}\ \Delta\pi_{t-3}^d + \underset{(0.19)}{0.805}\ \Delta i_{t-1}^d + \underset{(0.008)}{0.0279}\ \text{I1974Q3}_t \\
& + \underset{(0.0057)}{0.0372}\ \Delta\text{I1979Q2}_t, \quad \hat{\sigma} = 0.00793,\ \bar{R}^2 = 0.37.
\end{aligned} \tag{19}
$$

The short-term interest rate differential is dominated by a stable self-feedback mechanism:

$$
\begin{aligned}
\widehat{\Delta i_t^d} \;=\;& -\underset{(0.047)}{0.169}\ i_{t-1}^d - \underset{(0.018)}{0.0528}\ \Delta\pi_{t-2}^d + \underset{(0.11)}{0.339}\ \Delta r_{t-2}^d - \underset{(0.003)}{0.0106}\ \text{I1977Q1}_t \\
& - \underset{(0.0021)}{0.0086}\ \Delta_4\text{I1980Q4}_t + \underset{(0.00037)}{0.00101}\ , \quad \hat{\sigma} = 0.0029,\ \bar{R}^2 = 0.28.
\end{aligned} \tag{20}
$$

It also reacts positively to changes in both the real bond rate, $\Delta(r^d - \pi^d)$, and the inflation rate:

$$\widehat{\Delta i_t^d} = -0.169 i_{t-1}^d + 0.339(\Delta r_{t-2}^d - \Delta\pi_{t-2}^d) + 0.286\Delta\pi_{t-2}^d + \text{det.terms}.$$

Returning to the Fisher interpretation, we can think here of the central banks' aggressive stance against realized past inflation as well as expected future inflation:

$$\widehat{\Delta i_t^d} = -0.169 i_{t-1}^d + 0.339\Delta\rho_{t-2} + 0.339(\Delta\pi_{t-2}^{e,d} - \Delta\pi_{t-2}^d) + 0.286\Delta\pi_{t-2}^d + \text{det.terms}.$$

However, the equation should not be interpreted as a backward-looking Taylor rule.[9]

As discussed earlier, the long-term interest rate differential represents the common stochastic trend of the system and is as such weakly exogenous for the cointegration relations. In the short run, it adjusts in response to contemporaneous and previous changes of the interest rate differentials. Being a unit root process it has the highest $\bar{R}^2$:

$$\begin{aligned}\widehat{\Delta r_t^d} \;=\; & \underset{(0.031)}{0.299}\ \Delta i_t^d + \underset{(0.037)}{0.118}\ \Delta i_{t-1}^d - \underset{(0.061)}{0.294}\ \Delta r_{t-1}^d + \underset{(0.00094)}{0.00776}\ \Delta \text{I1974Q4}_t \\ & - \underset{(0.0013)}{0.00562}\ \text{I1979Q1}_t - \underset{(0.0013)}{0.00432}\ \text{I1980Q1}_t, \quad \hat{\sigma} = 0.0012,\ \bar{R}^2 = 0.63.\end{aligned} \tag{21}$$

In the four equations so far, the exchange rate is not involved. The nominal exchange rate is in our model not pushing the rest of the system, but purely adjusting to a vast amount of information crossing markets:

$$\begin{aligned}\widehat{\Delta e_t} \;=\; & -\underset{(0.026)}{0.0916}\ (e_{t-1} - 26.4 r_{t-1}^d) + \underset{(0.38)}{1.15}\ \Delta\pi_{t-1}^d + \underset{(0.44)}{1.38}\ \Delta\pi_{t-2}^d + \underset{(0.38)}{1.1}\ \Delta\pi_{t-3}^d \\ & - \underset{(1.9)}{5.82}\ \Delta r_{t-3}^d + \underset{(0.079)}{0.299}\ \Delta e_{t-1} - \underset{(0.08)}{0.218}\ \Delta e_{t-2} + \underset{(0.082)}{0.238}\ \Delta e_{t-3} + \underset{(0.013)}{0.0394}\ , \\ & \hat{\sigma} = 0.04679,\ \bar{R}^2 = 0.23.\end{aligned} \tag{22}$$

The exchange rate equation is driven by the fourth cointegration vector, which stabilizes the exchange rate along the common stochastic trend given by the long-term interest rate differential. With 10%

[9] We will return to this issue in §4.

adjustment of the exchange rate per quarter, there is more predictability than allowed under UIP. The high significance of the error correction terms shows the large loss of information, that would occur when the VAR would be specified in differences.

3.2.4 Testing for the validity and congruency of the model

The congruency of the model is investigated in Table 7. For the highly reduced model, there are no signs of dynamic misspecification. Some issues of non-normality and heteroscedasticity remain. A re-estimation of the equations with GARCH errors largely replicated the results presented here. As the efficiency of the single-equation reduction procedure applied here depends on the lack of correlation among the error terms of the model, we also verified that all contemporaneous effects had been captured by the causal search algorithm.

Insert Table 7 about here

4 The effects of a monetary policy shock

In this section, we consider the dynamic responses to an asymmetric monetary policy shock in form of an unpredicted one percentage-point increase of the nominal short-term interest rate differential[10], $i_t^d = i_t - i_t^*$. We start with the impulse response analysis of the model just derived, before evaluating the robustness of the results in three alternative scenarios.

4.1 An impulse response analysis

Figure 6 displays the responses of the system variables, i.e., the inflation differential π^d, the output gap differential, y^d, the 3-month interest rate spread i^d, the 10-year government bond yield spread, r^d, and the nominal exchange rate, e, with regard to an one-percentage point increase in the quarterly 3-month treasury bill return differential. Three sets of impulse response functions are plotted offering insights into the dynamics of the unrestricted cointegrated VAR (top panel), the exactly identified SVECM, and the selected PSVECM (bottom panel) presented in §3.2.3. For the just-identified SVECM, the causal

[10] Note that due to the construction of the interest rate variable, a unit shock in i_t^d corresponds to an approximately 400 basis point increase in the 3M treasury bill interest rate differential.

order $\pi_t^d \to y_t^d \to i_t^d \to r_t^d \to e_t$ was imposed to ensure consistency with the TETRAD results in §3.2.2 and making the SVECM nest the PSVECM.

Insert Figure 6 about here

The most striking feature when comparing the three sets of impulse response functions is the remarkable difference in the width of the 95% confidence intervals between the selected PSVECM and the unrestricted systems. The confidence intervals have been computed using the bootstrap procedure of Hall (1992) with 2000 replications. For unrestricted systems, hardly anything substantial can be said about the responses to a monetary policy shock as only very few elements of the impulse response functions are statistically different from zero. As shown in Krolzig (2003), we can expect the estimated responses of the PSVECM to be much closer to economic reality (in the MSE sense).

Common features among the three models are, firstly, the short life span of the increases in i^d, the short-term interest rate spread is fading out quickly within ten to twenty quarters after the initial shock, and, secondly, the presence of a pronounced price puzzle: In the short-term (up to 2-3 years), both the output gap and the inflation rate differential increase after a monetary tightening. This, from a theoretical but not empirical perspective, surprising result could to some extent be due to the backward-looking nature of the VAR approach. Suppose that monetary policy can be described by a forward-looking Taylor rule, then it is conceivable that, under rational expectations, the impulse response function measures this inverse causality from state of (expected) future inflation and excess demand to current policy rates. [11]

In the PSVECM, the long-term bond yield differential jumps in response to a monetary policy shock, with the response of the former being about a third of the size of the initial shock, $\nabla r_t^d \approx 0.3\nabla i_t^d$, followed by a slow but steady decline to the origin. The ad-hoc reaction of the nominal bond spread without an immediate reaction of the exchange rate leads to a disequilibrium with an abnormally high nominal bond yield spread for the first six periods and a relatively undervalued currency. In turn this drives a steady appreciation for several quarters before the exchange rate is finally depreciating and returns to its original level after ten years.

The delayed overshooting puzzle merits further analysis as its presence is forced by the lack of a contemporaneous link between the short-term interest rate differential and exchange rate in the

[11] In §4.4, we show that the possibly resulting identification issues do not affect the delayed overshooting finding.

PSVECM, as determined by the specification of the direct, $b_{ei} = \partial e_t / \partial i_t^d$, and indirect interest rate channel of monetary policy, $b_{er} = \partial e_t / \partial r_t^d > 0$. As shown in §3.2.2, a shock in i_t^d leads to an immediate jump in the bond yield differential, $b_{ri} = \partial r_t^d / \partial i_t^d > 0$, but for the other variables, including the exchange rate, no statistical evidence was found in support of an impact effect.[12] In the case of the lack of this link, a slow response of the exchange rate follows, implying a delayed overshooting and the violation of UIP.

4.2 Allowing the exchange rate to jump

In the following we examine the robustness of our results for the $\$/\pounds$ exchange rate by adding a contemporaneous link between short-term interest rate differential and exchange rate to the PSVECM. The key question is whether the delayed overshooting puzzle will emerge *even if* we allow the exchange rate to jump in response to shocks to the short-term interest rates – despite the lack of any statistical evidence for this behaviour. Imposing the sign restriction $b_{ei} \neq 0$ forces the selection of i_t^d:

$$\begin{aligned} \widehat{\Delta e_t} \;=\; & -\underset{(0.027)}{0.0803}\,(e_{t-1} - 26.4 r_{t-1}^d) + \underset{(1.23)}{2.11}\,\Delta i_t^d + \underset{(0.081)}{0.258}\,\Delta e_{t-1} \\ & -\underset{(0.082)}{0.182}\,\Delta e_{t-2} + \underset{(0.013)}{0.0339}\,, \quad \hat{\sigma} = 0.050,\; \bar{R}^2 = 0.12. \end{aligned} \tag{23}$$

The new exchange rate equation has changed greatly when compared to the baseline model in (22) with a pronounced drop in $\bar{R}^2$. Note that $\hat{b}_{ei} = 2.11$ with a t-value of 1.715 remained statistically insignificant at 5%.

The critical question is now whether or not our delayed overshooting result remains robust when the possibility of an immediate jump of the exchange rate is taken explicitly into account. The upper panels of Figure 7 seek to shed some light on this matter by comparing the emerging impulse response function of e_{t+h} with regard to a unit shock in i_t^d with our earlier results for the just-identified SVECM and the baseline model. The panel on the left displays the rather persistent response of the exchange rate in the SVECM. Due to the underlying estimation uncertainty, the result is fairly inconclusive. Only for the four quarters following the shock is the response significantly different from nil. In contrast, the two other figures highlight the advantage of the selected PSVECM. The second panel corresponds to

[12] Indeed, TETRAD consistently finds no evidence for such a link even at the 10% significance level, though the efficacy of the PC algorithm could have been affected by a poor signal-to-noise ratio (see Demiralp and Hoover, 2003).

the preferred statistical model, in which only the bond yield jumps in response to the monetary policy shock. As already seen in §4.1, there is clear evidence for the presence of an 'overshooting puzzle' with the exchange rate peaking after eight quarters. The panel on the right depicts the alternative scenario where the exchange rate is allowed to jump instantaneously in response to the monetary policy shock. The impact effect is with an estimated value of 2.11 is statistically close to the peak response of 3.52 in the baseline scenario as the latter lies within the former's 95% confidence band, thereby eliminating some excess return potentials, which will be analyzed in more detail in §4.3. However, due to the poorly estimated impact parameter, the width of the confidence band for the initial response is also consistent with the lagged response of the exchange rate in the baseline model.

Insert Figure 7 about here

4.3 Dissecting the delayed overshooting puzzle

Having established solid evidence for the delayed overshooting hypothesis, we finally examine its implications for the size and dynamic profile of the violations of UIP during the transmission process of the monetary policy shock. We again compare the responses of the exchange rate for the exactly identified recursive SVAR with the baseline model, i.e., the best statistical representation of the macro dynamics in the sample period, and the alternative model where e_t is allowed to jump.

Delayed overshooting generates excess returns violating UIP. This can be seen in the top panels of Figure 7 from the deviation of the response of the exchange rate, ∇e_h, from the line entitled UIP representing the equilibrium response of the exchange rate consistent with the uncovered interest parity hypothesis:

$$\nabla e_h^{\mathsf{UIP}} = \nabla e_T + \sum_{s=h}^{T-1} \nabla i_s^d \tag{24}$$

where in the plots above $T = 150$ was used.[13] When comparing the blue UIP line (as defined above) with the green confidence bands, it can be seen that, even for the alternative model, the initial jump is insufficient and the final depreciation process is too sluggish to meet the UIP target.

Approaching the same issue from a different angle, the panels below measure the deviations from

[13] As the exchange rate converges much earlier, the results are insensitive to the choice of T.

UIP with the ex-ante one-period excess return series:

$$\nabla \xi_h = \nabla i_h^d + \Delta \nabla e_{h+1}. \tag{25}$$

The plots reveal, consistently among models, excess returns for UK treasury bonds after a tightening of Bank of England policy. In the bottom panels, the cumulated excess returns,

$$\nabla \xi_{0,h} = \sum_{j=0}^{h} \nabla \xi_j = \sum_{j=0}^{h} \nabla i_j^d + \Delta_h \nabla e_{h+1}, \tag{26}$$

are plotted. Over the first two years excess returns of up to 7.5 percent are observed constituting major violations of the UIP hypothesis. Across models, there is strong evidence for the presence of a delayed overshooting puzzle for the $\$/\pounds$ exchange rate.

4.4 Assessing the robustness of the delayed overshooting puzzle

During the model discussion and impulse response analysis, we observed that the PSVECM exhibits all signs of a price puzzle in combination with an output gap puzzle, which raises serious questions regarding the identification of the monetary policy shock. In light of the possible presence of an identification problem, an assessment of the robustness of the results of the impulse response analysis is warranted.[14] To dissect the contribution of the price and output puzzles for the emergence of the delayed overshooting puzzle, we ran a thought experiment in which the links from the real sector of the economy, $(\pi_t^d, y_t^d)'$, to the financial markets, $(i_t^d, r_t^d, e_t)'$, were cut off. As shown by the 'no interaction' impulse responses in the upper panels of Figure 8, the exchange rate reaction is qualitatively unchanged when the inflation and output gap differentials are fixed, demonstrating that the price puzzle can not invalidate the delayed overshooting results of the paper. In a second experiment two theory-driven model interventions were implemented: A sign restriction on the (real) interest rate terms in the output gap equation and a substitution of the bond yield differential in the inflation equation by an adaptively updated inflation expectation differential. The impulse response analysis confirms that this alternative identification scheme, deduced from economic theory but born out of the spirit of the PSVECM, not only resolves the price and the output puzzles (lower panels) but also reproduces our delayed

[14] A detailed description can be found in the Kent Discussion Paper version of this paper.

overshooting results qualitatively and, when the exchange rate is allowed to jump, quantitatively (upper panels).

5 Conclusion

In our investigation into the presence of a 'delayed overshooting puzzle' in the response of the $\$/\pounds$ exchange rate to an asymmetric monetary policy shock in the UK and the US, we emphasized the need to let the data speak. To facilitate a congruent representation of the macro dynamics in the sample period, we proposed a data-driven modelling approach combining a VAR based cointegration analysis with a graph-theoretic search for instantaneous causal relations and an automatic general-to-specific approach for the selection of a parsimonious structural vector equilibrium correction model. We can now conclude by summarizing the main findings of our econometric analysis:

(i) *Long-run properties.* We found four cointegration relations and one stochastic trend, which could be identified as the long-term interest rate differential, r_t^d, and appeared to be driven by long-term inflation expectations as in the Fisher hypothesis. r_t^d cointegrated with the inflation differential π_t^d to a stationary 'real' long-term rate differential. It was also found to drive the exchange rate, $e_t = 26.4 r_t^d$, which is consistent with UIP and stationary long-term exchange rate expectations, $\mathsf{E}_t e_{t+40}$. The output gap, y_t^d, and the short-term rate differential, i_t^d, are error-correcting and weakly exogenous to the other cointegration relations.

(ii) *The short-run dynamics.* The bond yield differential, r_t^d, jumping in the case of shocks in the short-rate differential, i_t^d, was the only statistically significant simultaneity in the model. Jumps in the exchange rate after monetary shocks are only significant at 10%. With a systemic certainty of 95%, we can be sure that the jump does not have the size needed for UIP to hold.

(iii) *Model reduction.* The need for parsimony was confirmed by the problem of an inconclusive impulse response analysis in the case of the unrestricted (S)VAR caused by the inherent estimation uncertainty due to the large number of parameters. The general-to-specific model selection procedures employed in this paper overcame those limitations.

(iv) *Monetary policy shock.* Consistently, we found strong evidence for delayed overshooting and violations of UIP.

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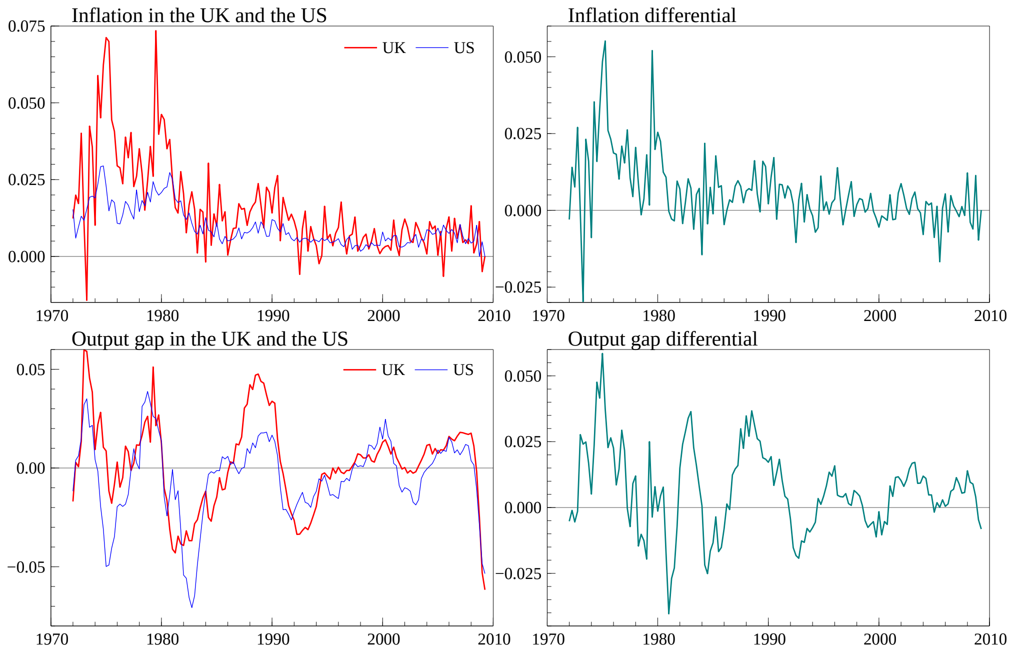

Figure 1 Inflation rates and output gaps.

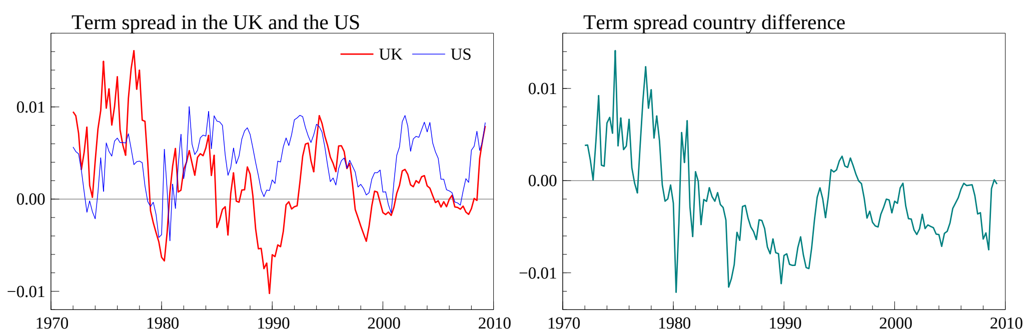

Figure 2 Term structure in the UK and the US and their differences.

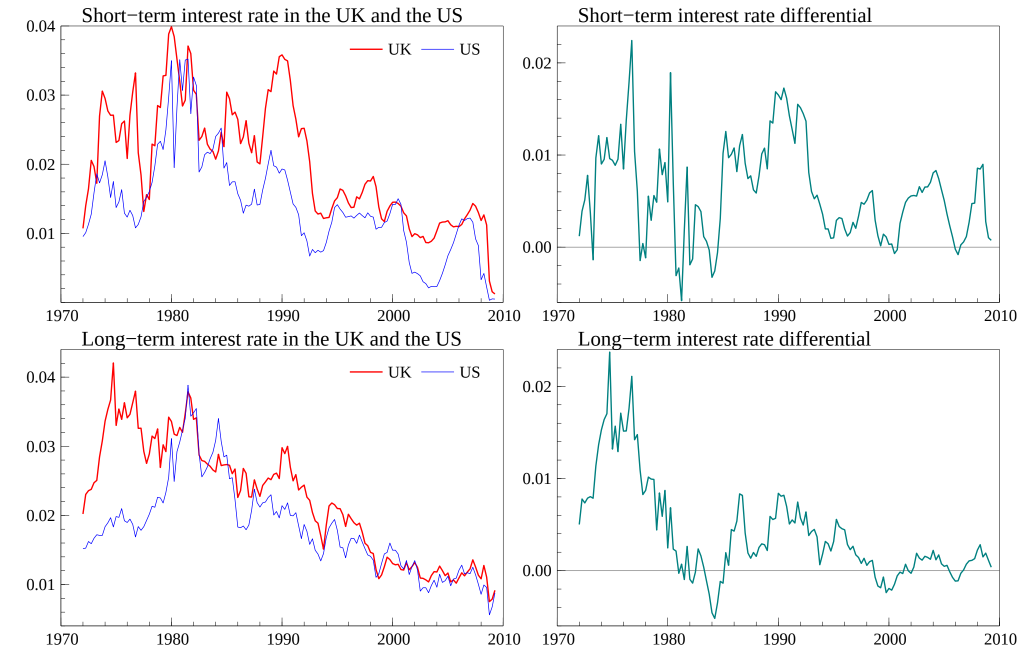

Figure 3 Nominal short- and long-term interest rates and country differences.

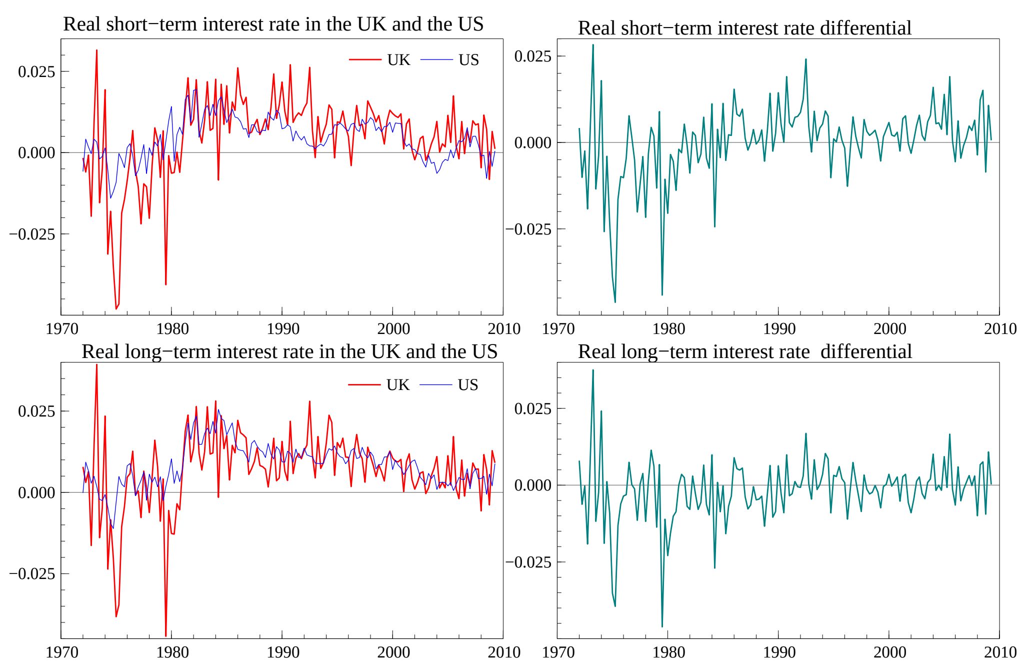

Figure 4 Naive real short- and long-term interest rates and country differences.

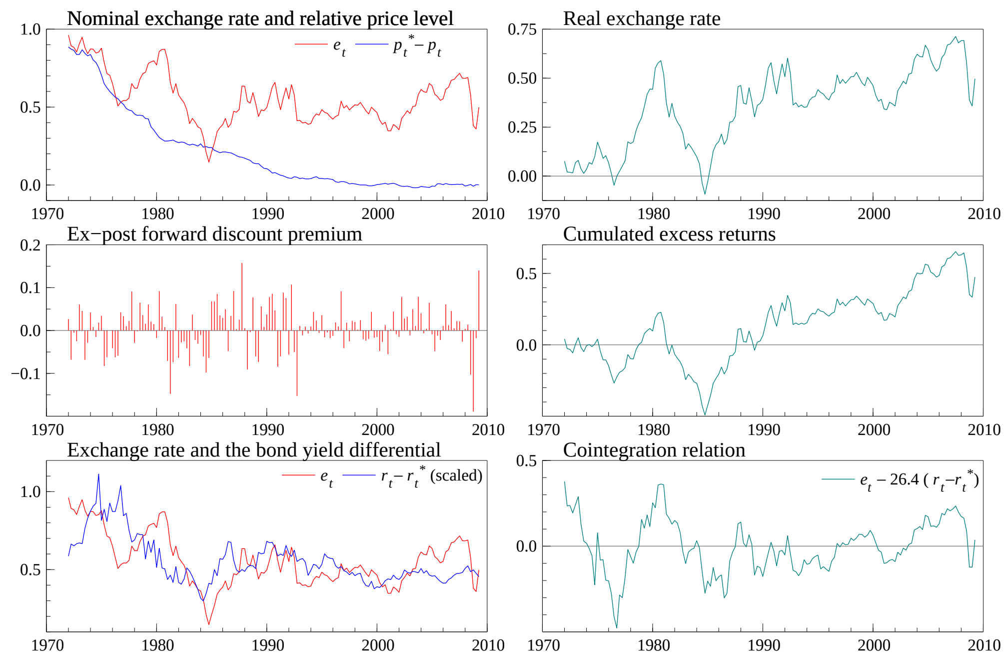

Figure 5 PPP, UIP and the $\$/\pounds$ exchange rate determination. Upper panels depicts the nominal exchange rate, relative prices and the real exchange rate; the central ones show the deviations from UIP: Ex-post excess returns and their cumulation; the bottom panels plot the long-run co-movement of the nominal exchange rate and the bond yield differential.

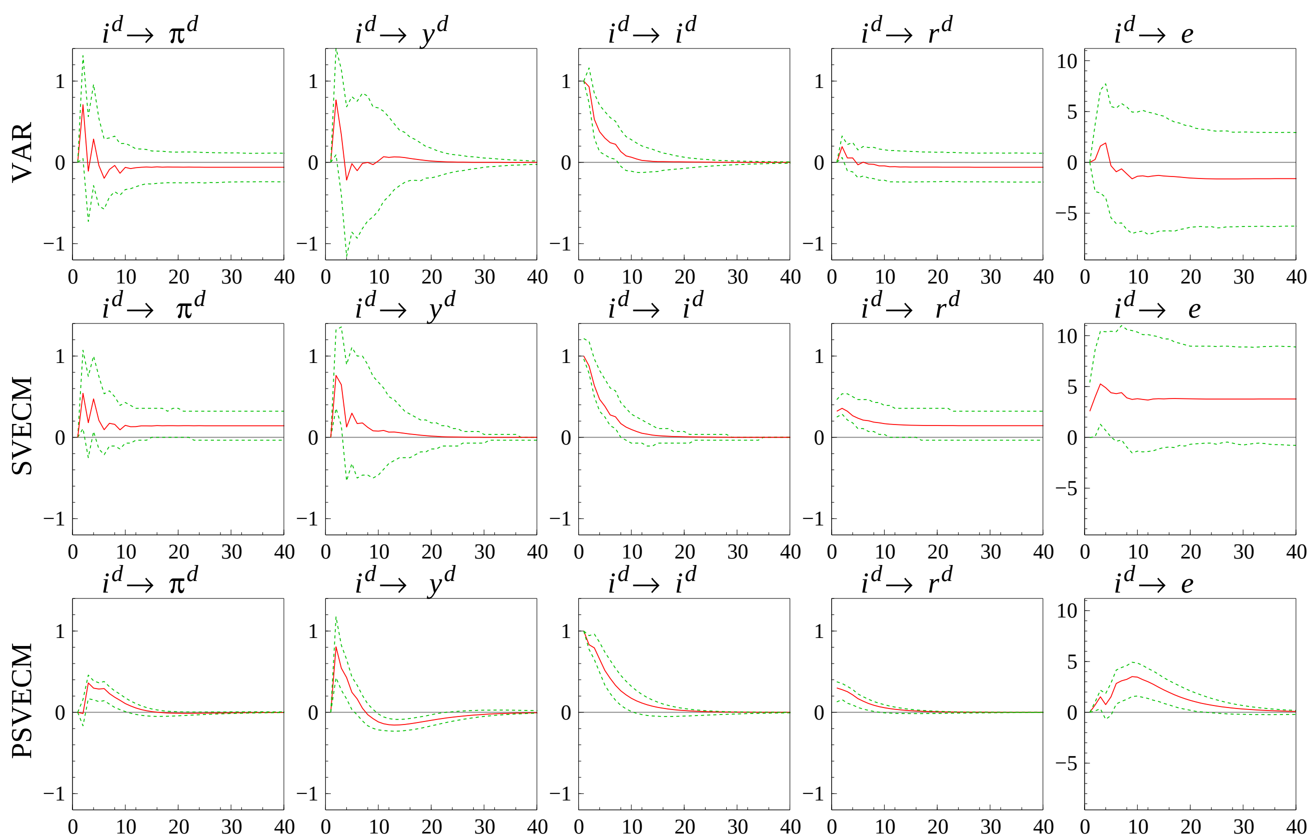

Figure 6 Responses to an asymmetric monetary policy shock in the cointegrated reduced-form VAR, the just-identified and parsimonious SVECM.

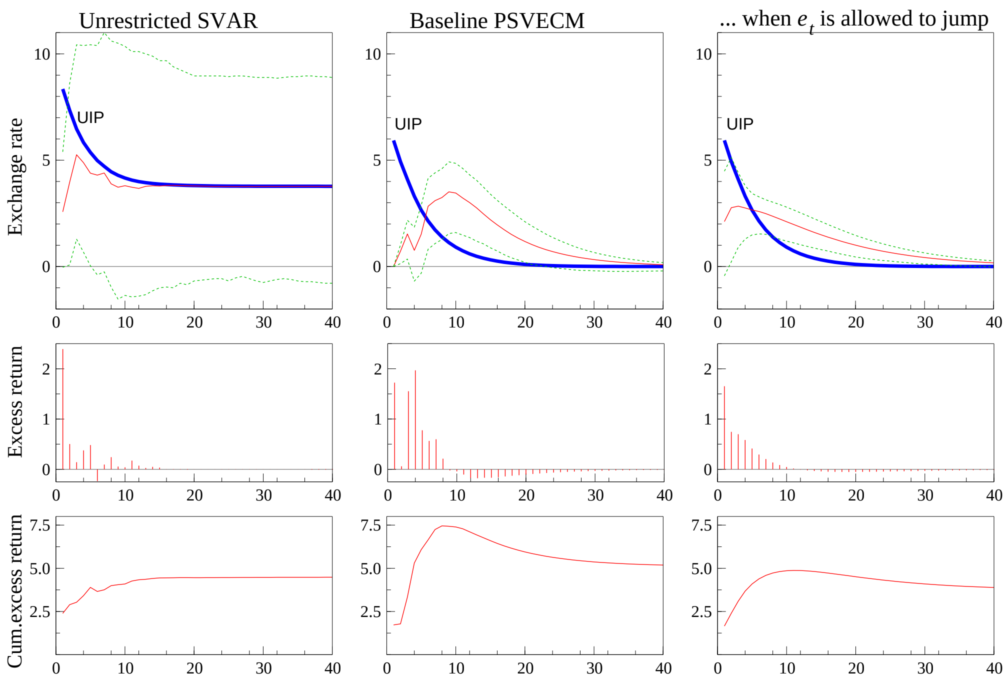

Figure 7 Delayed overshooting in response to an asymmetric monetary policy shock in the CVAR, the baseline PSVECM and alternative PSVECM: Responses of exchange rate, ∇e_{t+h}, the excess return, $\nabla \xi_h$ and cumulative excess return, $\nabla \xi_{0,h}$, to a one percentage point increase in in the 3-month interest rate differential, i_t^d (with 95% confidence intervals) .

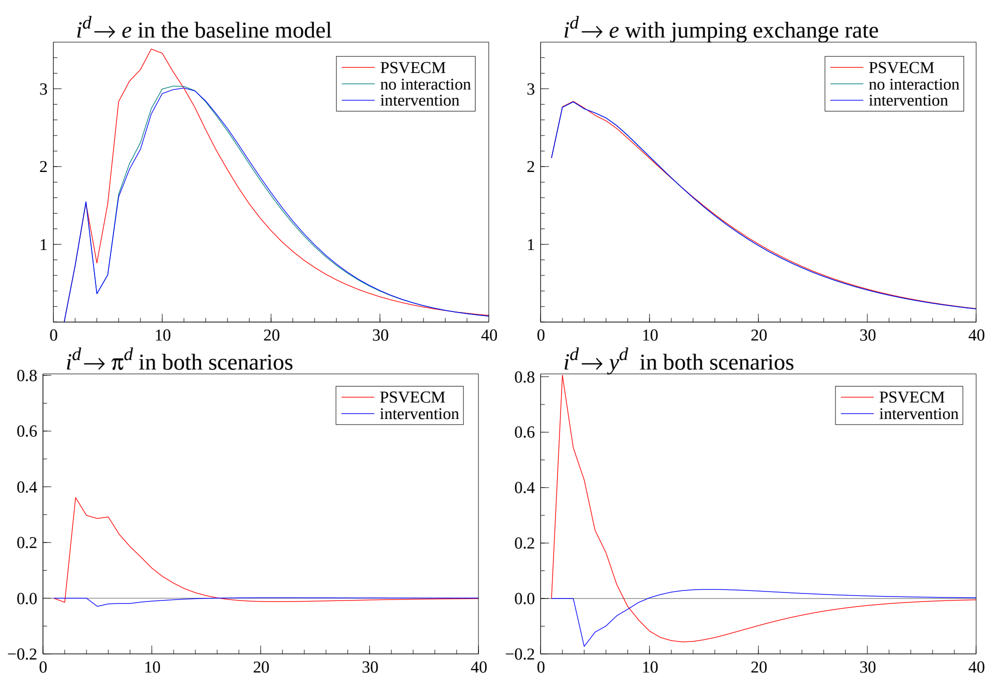

Figure 8 Impulse responses of the inflation differential, the output gap differential and of the exchange rate to a monetary policy shock in the baseline model, the model without feedback from the real sector of the economy, and the intervention model with a sign-restricted IS curve and an adaptive inflation expectation mechanism.

Table 1 Time Series Definitions and Source.

Variable	Description	Source	EcoWin code
P_t	UK GDP implicit price deflator (2005=100), SA	OECD	$qna : gbr\_1489758547q$
I_t	UK treasury bills yield, 3 months, GBP	ONS	$ons : md\_ajnbq$
R_t	UK government bond yield benchmarks, bid, 10 years, GBP	Reuters	$ew : gbr14020$
Y_t	UK output gap of the total economy	OECD	$oe : gbr\_gapq$
P_t^*	US GDP implicit price deflator (2005=100), SA	OECD	$qna : usa\_463541155q$
I_t^*	US treasury bills yield, 3 months, USD	Reuters	$ew : usa14430$
R_t^*	US government constant maturity yield, 10 years, USD	Fed	$ew : usa1402010$
Y_t^*	US output gap of the total economy	OECD	$oe : usa\_gapq$
e_t	Spot rates, GBP/USD transformed to USD/GBP,	Reuters	$ew : gbr19005$

Variables without a superscript are of the domestic country (UK), * indicates the foreign country (US) and ^{d} indicates a country difference. All financial variables are end-of-period series.

Table 2 Model variables.

Variable	Description
$\pi_t = \Delta \log P_t$	rate of inflation
$y_t = \log(1 + Y_t/100)$	output gap
$i_t = \log(1 + I_t/400)$	short-term interest rate
$r_t = \log(1 + R_t/400)$	long-term interest rate
$e_t = \log E_t$	exchange rate

Table 3 Johansen likelihood ratio trace test of $H_0 : \mathrm{rank} \leq \mathrm{r}$.

r	eigenvalue	trace test	prob
0	0.251	95.595 **	[0.000]
1	0.118	52.287 *	[0.017]
2	0.107	33.390 *	[0.018]
3	0.087	16.369 *	[0.035]
4	0.018	2.725	[0.099]

** significant at 1% level, * significant at 5% level.

Table 4 Testing for weak exogeneity.

	π_t^d	y_t^d	i_t^d	r_t^d	e_t
$\chi^2(4)$	28.0	19.7	12.9	0.84	15.4
	[0.00]	[0.00]	[0.01]	[0.93]	[0.00]

Table 5 Cointegration vectors and loadings, t-values in brackets.

	Cointegration vectors				Loadings			
	β_1	β_2	β_3	β_4	α_1	α_2	α_3	α_4
π_t^d	0	0	-1	0	0.103	-0.051	0.733^{**}	0.0002
					(1.60)	(-0.25)	(4.89)	(-0.03)
y_t^d	1	0	0	0	-0.246^{**}	0.106	0.098	0.0011
					(-3.86)	(0.52)	(0.66)	(0.19)
i_t^d	0	1	0	0	0.0218	-0.249^{**}	-0.043	0.0014
					(0.84)	(-3.64)	(-0.86)	(-0.70)
r_t^d	0	0	1	-26.4	0.003	-0.019	0.010	0.0010
				(3.6)	(-0.23)	(-0.44)	(0.32)	(0.85)
e_t	0	0	0	1	-0.383	-0.288	0.159	-0.0983^{**}
					(-1.16)	(-0.27)	(0.20)	(-3.25)

** significant at 1% level, * significant at 5% level.

Table 6 Contemporaneous correlation.

	π_t^d	y_t^d	i_t^d	r_t^d	e_t
π_t^d	1				
y_t^d	-0.11	1			
i_t^d	0.15	0.18	1		
r_t^d	-0.05	0.15	0.52^{**}	1	
e_t	-0.15	0.04	0.14	0.04	1

** significant at 1% level, * significant at 5% level.

Table 7 Misspecification tests of the parsimonious SVECM.

Test	π_t^d	y_t^d	i_t^d	r_t^d	e_t
AR 1-5	1.372	1.099	0.813	1.193	0.495
	[0.239]	[0.364]	[0.543]	[0.316]	[0.779]
Normality	9.546^{**}	6.917^{*}	21.482^{**}	3.841	1.462
	[0.009]	[0.032]	[0.000]	[0.146]	[0.481]
ARCH 1-4	5.826^{**}	1.223	2.604^{*}	2.675^{*}	2.680^{*}
	[0.000]	[0.304]	[0.039]	[0.035]	[0.034]
Hetero	3.709^{**}	1.867	2.699^{**}	1.113	1.029
	[0.000]	[0.062]	[0.006]	[0.357]	[0.432]
RESET	6.588^{*}	1.610	0.668	0.923	4.257^{*}
	[0.011]	[0.207]	[0.414]	[0.338]	[0.041]

** significant at 1% level, * significant at 5% level.