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Is there a 'delayed overshooting puzzle'?**

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December 2011

KDPE 1124



Effects of monetary policy on the $\$/\pounds$ exchange rate. Is there a ‘delayed overshooting puzzle’?

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8 December 2011

Abstract

The determination of the $\$/\pounds$ exchange rate is studied in a small symmetric macroeconomic 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.

*Corresponding author, email: hm.krolzig@gmail.com. Paper prepared for the Conference on ‘Advances in International Economics and Economic Dynamics’ in Honor of Giancarlo Gandolfo, Rome 20-21 September 2010. We are grateful to Andrea Bubula, Alan Carruth, Katsuyuki Shibayama, and audiences at seminars at Kent, DIW Macroeconomic Workshop, Berlin, Royal Economics Society 2011 Meeting, Royal Holloway, European Economics and Finance Society Conference, London, ESEM 2011, Oslo, and the Conference on New Developments in Time Series Econometrics, Florence, as well as an unknown referee and the editor Giuseppe de Arcangelis for helpful comments and suggestions. The usual disclaimer applies.

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 $\$/\pounds$ 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 do not tend to jump instantaneously as predicted by the theory, but rather appreciate steadily for several months before finally depreciating. Whether or not such a ‘delayed overshooting puzzle’ is present in the case of the $\$/\pounds$ 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, 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 the ‘liquidity model’ where, in contrast to sticky price models, goods prices are flexible while asset markets only adjust 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

¹There 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.

²Some 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.

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, 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 US-Germany, US-UK, US-Japan and US-G7, they found evidence for delayed overshooting with a delay of around 2 years. The delay in the response of the US-UK 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 which exploits 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 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

³The 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.

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 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 §2 we introduce the data set and provide a brief overview of the UK-US macro history since the breakdown of Bretton Woods in light of the international parity conditions. This will give us valuable insight into the econometric modelling to be discussed in §3, which explains the methodology and reviews the derivation of the model. §4 investigates the effects of a monetary policy shock with focus on the presence of a delayed overshooting puzzle and violations of UIP. Finally §5 concludes.

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

2.1 Time Series

The small macroeconomic 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. Using quarterly time series, rather than monthly ones used in most of the literature, allows us to work with GDP-based measures of inflation and the output gap.

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

Table 1 Time Series Definitions and Source.

Variable	Description	Source	EcoWin code
P_t	UK GDP implicit price deflator (2005=100), SA	OECD	<i>qna : gbr_1489758547q</i>
I_t	UK treasury bills yield, 3 months, GBP	ONS	<i>ons : md_ajnbq</i>
R_t	UK government bond yield benchmarks, bid, 10 years, GBP	Reuters	<i>ew : gbr14020</i>
Y_t	UK output gap of the total economy	OECD	<i>oe : gbr_gapq</i>
P_t^*	US GDP implicit price deflator (2005=100), SA	OECD	<i>qna : usa_463541155q</i>
I_t^*	US treasury bills yield, 3 months, USD	Reuters	<i>ew : usa14430</i>
R_t^*	US government constant maturity yield, 10 years, USD	Fed	<i>ew : usa1402010</i>
Y_t^*	US output gap of the total economy	OECD	<i>oe : usa_gapq</i>
e_t	Spot rates, GBP/USD transformed to USD/GBP,	Reuters	<i>ew : gbr19005</i>

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

US of opposite sign but of 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:

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

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 rates of inflation and output gaps in the UK, the US, and the difference between the two countries. 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. 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.

In contrast to π_t^d and y_t^d , all other series were found to be $I(1)$. Moving to the asset markets, the further

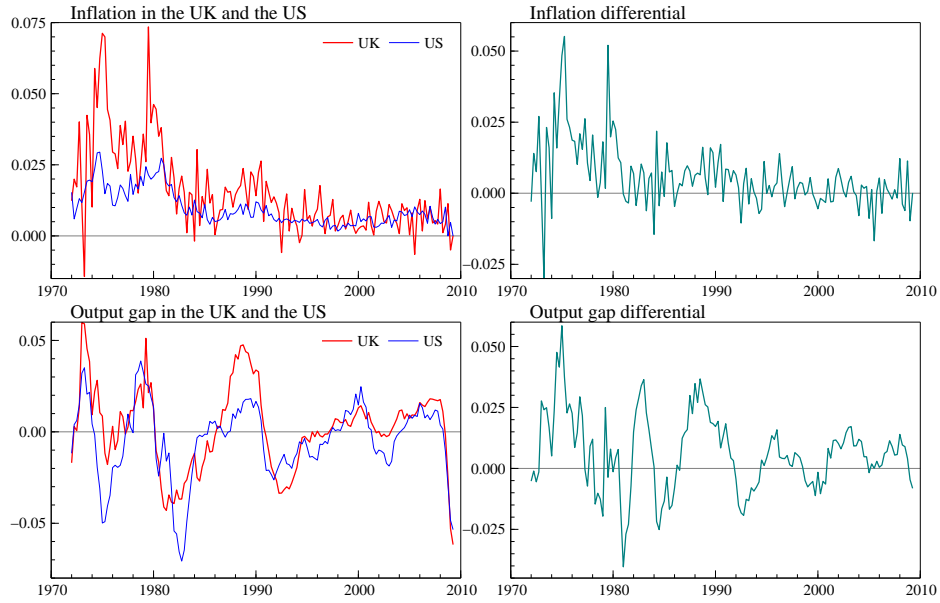


Figure 1 Inflation rates and output gaps.

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

2.2.2 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 follows 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 Figure 2, purchasing power parity clearly does not hold for the $\$/\mathcal{L}$ 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.

2.2.3 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 average of the expected one-period interest rates plus a potential risk premium, ϕ_t :

$$r_t = \frac{1}{T} \sum_{j=0}^{T-1} E_t i_{t+j} + \phi_t. \quad (1)$$

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, $r_t - i_t \sim I(0)$.

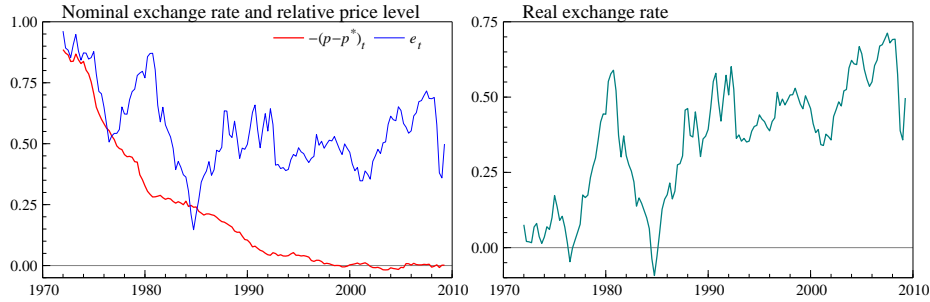


Figure 2 Nominal exchange rate, relative prices and the real exchange rate.

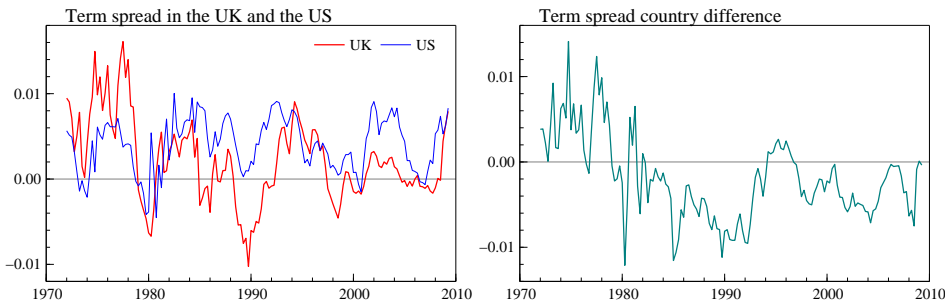


Figure 3 Term structure in the UK and the US and their differences.

Figure 3 plots the term spread for the UK and the US and their difference. 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.

2.2.4 Nominal interest rate parity

Figure 4 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.

2.2.5 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,

$$i_t = \rho_t + E_t \pi_{t+1}, \quad (2)$$

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.:

$$\rho_t - \rho_t^* = (i_t - E_t \pi_{t+1}) - (i_t^* - E_t \pi_{t+1}^*) \sim I(0). \quad (3)$$

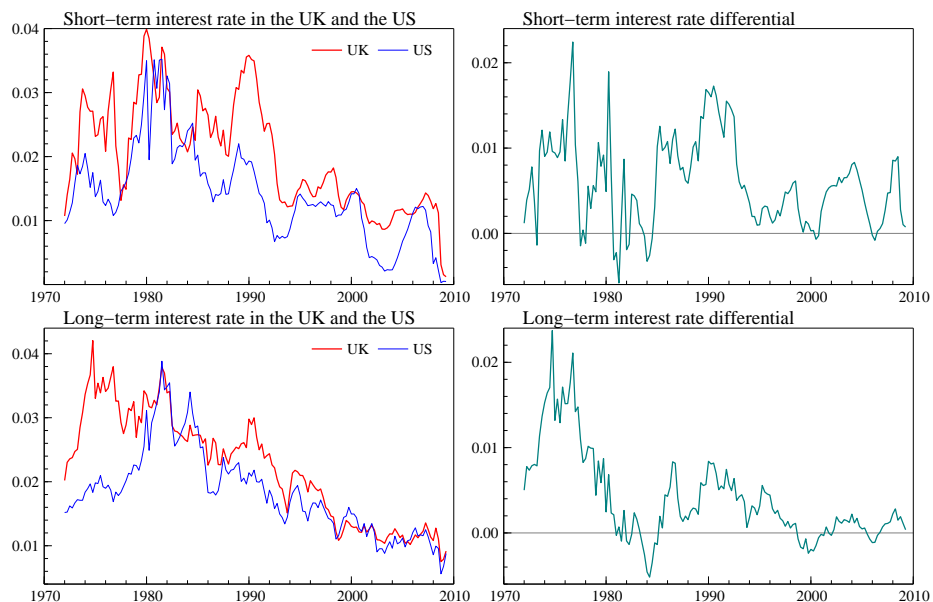


Figure 4 Nominal short- and long-term interest rates and country differences.

Theoretically, the calculation of ex-ante real interest rate involves future inflation expectations. As those are empirically difficult to measure, we focus here on a naive definition of the real interest rate using the current backward-looking inflation.⁴ These are plotted in Figure 5. 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 the real short-term differential.

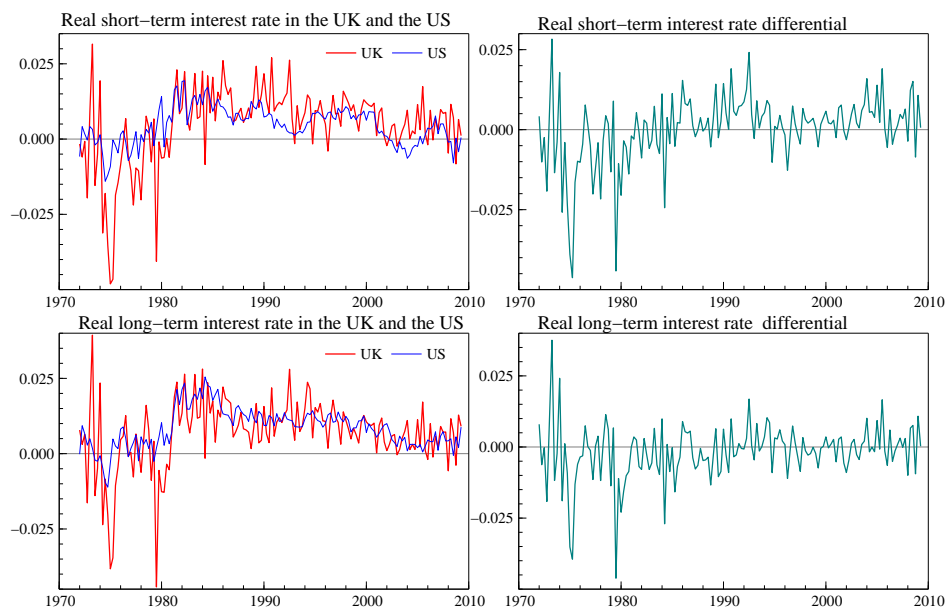


Figure 5 Naive real short- and long-term interest rates and country differences.

⁴A 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.

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 expected return, measured in the home currency, on a foreign asset with otherwise identical characteristics. For a one-period bond, this implies:

$$i_t = i_t^* - E_t \Delta e_{t+1}. \quad (4)$$

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

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

where ξ_t is a martingale difference sequence and measures the excess return of the UK bond. The realized excess returns over the sample period and their cumulation can be seen in Figure 6.

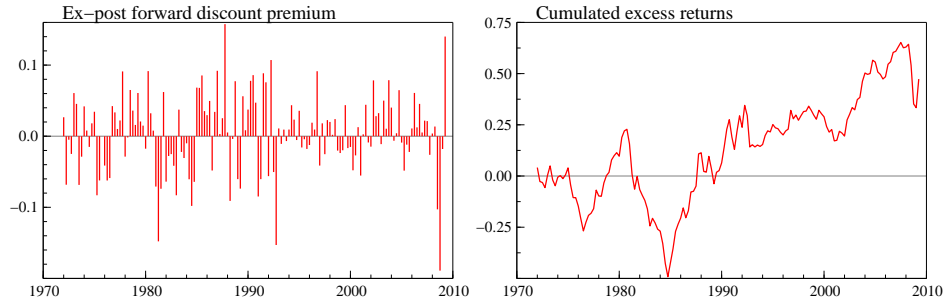


Figure 6 Deviations from UIP: Ex-post excess returns and their cumulation.

The UIP condition in (4) has been formulated for a one-period bond. We now consider its generalization to bonds with multi-period maturities. 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:

$$r_t^d = \frac{1}{T} \sum_{j=0}^{T-1} E_t i_{t+j}^d. \quad (6)$$

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

$$e_t = E_t e_{t+T} + \sum_{j=0}^{T-1} E_t i_{t+j}^d, \quad (7)$$

we get the multi-period form of UIP,

$$e_t = E_t e_{t+T} + T(r - r^*)_t, \quad (8)$$

which states that the exchange rate is determined by the long-term exchange rate expectation, $E_t e_{t+T}$, and T times the long-term interest rate differential. Note that this relation will not hold exactly in our data set due to the different type of bonds under consideration, in which case the impact of the bond yield differential is expected to be systematically smaller.

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:

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

where $\boldsymbol{\varepsilon}_t \sim \text{NID}(\mathbf{0}, \boldsymbol{\Sigma})$ 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:

$$\Delta \mathbf{y}_t = \boldsymbol{\nu} + \boldsymbol{\Pi} \mathbf{y}_{t-1} + \sum_{j=1}^{p-1} \boldsymbol{\Gamma}_j \Delta \mathbf{y}_{t-j} + \boldsymbol{\varepsilon}_t. \quad (10)$$

For a cointegrated vector process, the reduced-rank matrix, $\boldsymbol{\Pi}$, can be 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, i.e., $\boldsymbol{\Pi} = \boldsymbol{\alpha} \boldsymbol{\beta}'$. The Johansen procedure delivers unique estimates of $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ as a result of requiring $\boldsymbol{\beta}$ to be orthogonal and normalized. These estimates provide a value for the unrestricted log-likelihood function to be compared to the log-likelihood under economically meaningful overidentifying restrictions, $\boldsymbol{\beta}^r$:

$$\Delta \mathbf{y}_t = \boldsymbol{\nu} + \boldsymbol{\alpha} \boldsymbol{\beta}^r' \mathbf{y}_{t-1} + \sum_{j=1}^{p-1} \boldsymbol{\Gamma}_j \Delta \mathbf{y}_{t-j} + \boldsymbol{\varepsilon}_t, \quad (11)$$

with $\boldsymbol{\Sigma} = \text{E}[\boldsymbol{\varepsilon}_t \boldsymbol{\varepsilon}_t']$. The empirical modeling procedure for finding the cointegration relations follows 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, $\hat{\Sigma}$, of the system in (11). 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 orientation step of 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 acyclicity condition.

If all edges could be oriented, a directed acyclic graph (DAG) results. Based on the identified contemporaneous causal structure of the system, the VECM in (11) can be represented as a *recursive* structural vector equilibrium correction mechanism (SVECM). By a 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 $\hat{\Sigma}$, 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. In step (iv) we will consider *Gets* reductions of the SVECM to reduce the complexity of the model.

If the PC algorithm finds a link but has insufficient information to identify if, say, ‘A causes B’ or ‘B causes A’, 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. An additional modelling stage is then required for the selection of \mathbf{B}^r and, thus, the identification of the direction of causality. Having applied the model reduction step in (iv) to each SVECM associated with one of the found contemporaneous causal structures, the dominant econometric model is finally selected in (v).

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

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

$$\mathbf{B}^r \Delta \mathbf{y}_t = \boldsymbol{\delta} + \tilde{\boldsymbol{\alpha}} (\beta^{r'} \mathbf{y}_{t-1}) + \sum_{j=1}^{p-1} \boldsymbol{\Upsilon}_j \Delta \mathbf{y}_{t-j} + \boldsymbol{\eta}_t, \quad \boldsymbol{\eta}_t \sim \text{NID}(\mathbf{0}, \boldsymbol{\Omega}), \quad (12)$$

where \mathbf{B}^r is the lower-triangular matrix found by TETRAD and $\boldsymbol{\Omega}$ is a diagonal variance-covariance matrix. A single-equation based *Gets* reduction procedure such as *PcGets* can be applied to the equations in (12) 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 $\boldsymbol{\Upsilon}_j$ in the structural VECM. The result is a parsimonious structural vector equilibrium correction model denoted PSVECM, which is nested in (12) and defined by the selected $\boldsymbol{\delta}^*$, $\tilde{\boldsymbol{\alpha}}^*$ and $\boldsymbol{\Upsilon}_j^*$ with $j = 1, \dots, p - 1$.

(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. Thus the dominant design of the contemporaneous effects matrix according to information criteria such as Akaike or Schwarz would be used.

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 . 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.

Table 3 Misspecification tests for equations of the unrestricted VAR(4).

Test		π_t^d	y_t^d	i_t^d	r_t^d	e_t
AR 1-5	F(5, 124)	0.721 [0.609]	1.388 [0.233]	1.315 [0.262]	0.559 [0.732]	1.062 [0.385]
Normality	$\chi^2(2)$	92.059** [0.000]	26.045** [0.000]	8.129* [0.017]	19.331** [0.000]	1.134 [0.567]
ARCH 1-4	F(4, 87)	1.074 [0.372]	1.769 [0.140]	4.956** [0.001]	2.972* [0.022]	2.281 [0.065]
Hetero	F(22, 72)	2.000** [0.004]	1.498 [0.059]	2.347** [0.001]	4.437** [0.000]	1.175 [0.263]

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

The results of tests for misspecification are displayed in Table 3. There are no problems of autocorrelation in the equations. However, the normality test shows serious non-normality mainly due to excess kurtosis in all but the exchange rate equation. Also, heteroscedasticity and ARCH effects in the interest rate equations are detected. But overall the residuals are sufficiently well behaved to proceed with the system.⁵

We continue by analyzing the long-run properties of the system. The number of stable long-run relations $\beta' \mathbf{y}_t$, which is equal to the rank of the matrix $\mathbf{\Pi}$ of the vector equilibrium-correction mechanism

⁵Some of the non-normalities can be traced back to the reduction in volatility during the Great Moderation as well as outliers, for which dummy variables will be included in the PSVECM in §3.2.4.

in (10), is determined by the Johansen (1995) test for $I(1)$ cointegration. The eigenvalues and trace test results are shown in Table 4. We find that the long-run properties of the system are characterized by four cointegration relations, $r = \text{rank}(\mathbf{\Pi}) = 4$. With dimension $K = 5$ and rank $r = 4$ there is one unit root in the system.

Table 4 Johansen likelihood ratio trace test of $H_0 : \text{rank} \leq r$.

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.

To identify the long-run structure of the system, we continue with the preliminary analysis of testing structural hypotheses regarding α and β . The test results for potential cointegration vectors are shown in Table 5. Here, we impose sequentially restrictions on one cointegration vector while leaving the others unconstrained. Hypotheses H_1 to H_5 test if the inflation spread, the interest rate spreads, the output gap spread or the nominal exchange rate constitute cointegration vectors, i.e., stationary relationships, on their own. According to the likelihood ratio (LR) test statistics, the output gap spread and the short-term interest rate spread are possible stationary cointegration relations. Hypotheses H_6 and H_7 state that real interest rate differentials are stationary. This is accepted with a p-value of 0.12 for the long-term but rejected for the short-term differential. H_8 rejects the stationarity of the country differences in the term structure, i.e., the spread between long and short-term interest rate differentials.

Table 5 Testing simple cointegration relations.

	π_t^d	y_t^d	i_t^d	r_t^d	e_t	$\chi^2(1)$	prob
H_1	1	0	0	0	0	10.60	[0.00]
H_2	0	1	0	0	0	1.52	[0.22]
H_3	0	0	1	0	0	2.34	[0.13]
H_4	0	0	0	1	0	10.80	[0.00]
H_5	0	0	0	0	1	4.55	[0.03]
H_6	1	0	-1	0	0	10.30	[0.00]
H_7	1	0	0	-1	0	2.37	[0.12]
H_8	0	0	1	-1	0	7.82	[0.01]

In Table 6 we test for long-run weak exogeneity of the variables of the system. Under the null hypothesis of a particular zero row in α , the corresponding variable is not adjusting towards the long-run equilibrium. The LR test results of the restrictions on α 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

Altogether, the tests of hypotheses H_1 to H_8 suggest three linearly independent cointegration vectors. Given a rank of four we will need to identify one further composite cointegration vector. Following the modelling approach suggested by Juselius (2006), the following cointegration vectors were identified

⁶In 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.

Table 6 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]

by paying attention not only to statistical acceptability but also to consistency with economic theory:⁷

(i) *Stationary output gap differential.*

$$y_t^d = y_t - y_t^* \sim I(0). \quad (13)$$

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.*

$$i_t^d = i_t - i_t^* \sim I(0). \quad (14)$$

This is somewhat surprising given our previous result 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.*

$$\rho_t^d = r_t^d - \pi_t^d = (r - \pi)_t - (r^* - \pi^*)_t \sim I(0). \quad (15)$$

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 differential⁸, which is driven by the same stochastic trend. It is also worth noting that,

⁷When analyzing a system consisting of the inflation differential as the only nominal variable and 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, $e_t + p_t^d$, the Johansen trace test indicated a cointegration rank of three rather than four in Table 4. The cointegration vectors were found to coincide with the relations (13) to (15) with the corresponding restrictions not rejected by the LR test at 36%. Furthermore, testing for weak exogeneity revealed that, at a marginal significance level of 0.31, the real exchange is one of two stochastic trends. The real exchange rate can also not be recovered as cointegration relation by a trend in the cointegration space. The restricted trend is statistically insignificant. When it is nevertheless included, the cointegration properties of the system break down as the Johansen procedure reports a rank of just one. In other words, while there is stable long-run relation between the *nominal* exchange rate and the *nominal* interest rate differential, no such relationship can be found for the real ones.

⁸Note that, while the ADF test indicated that the inflation differential is $I(0)$, π_t^d is not identified as one of the stationary components in the cointegration space. To understand these contradicting findings, we compare the standard ECM(2) equation used for the ADF test

$$\Delta\pi_t^d = \nu_\pi + \alpha_\pi\pi_{t-1}^d + \sum_{j=1}^2\gamma_{\pi,j}\Delta\pi_{t-j}^d + \varepsilon_{\pi,t}$$

with an augmented ECM(2) equation that also contains the short-run dynamics of all other system variables, $\mathbf{x}_t = (y_t^d, i_t^d, r_t^d, e_t)^d$, as in the inflation differential equation of the SVECM:

$$\Delta\pi_t^d = \nu_\pi^* + \alpha_\pi^*\pi_{t-1}^d + \sum_{j=1}^2\gamma_{\pi,j}^*\Delta\pi_{t-j}^d + \sum_{j=1}^2\gamma'_{\pi\mathbf{x},j}\Delta\mathbf{x}_{t-j} + \varepsilon_{\pi,t}^*.$$

Whereas the ADF test statistic, -3.53 , is significant at 1%, the t-value of α^* in the augmented ECM is with -2.81 greater than the 5% critical value of -2.88 . Hence, the conflicting results can largely be traced back to use of two different information sets.

due to (14) and (15), the UK and US term structures do not cointegrate.

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

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

$$e_t - 26.4(r - r^*)_t \sim I(0). \quad (16)$$

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:

$$e_t = E_t e_{t+40} + 40r_t^d. \quad (17)$$

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 (16) and (17), 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 7.

Table 7 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 (1.60)	-0.051 (-0.25)	0.733** (4.89)	0.0002 (-0.03)
y_t^d	1	0	0	0	-0.246** (-3.86)	0.106 (0.52)	0.098 (0.66)	0.0011 (0.19)
i_t^d	0	1	0	0	0.0218 (0.84)	-0.249** (-3.64)	-0.043 (-0.86)	0.0014 (-0.70)
r_t^d	0	0	1	-26.4 (3.6)	0.003 (-0.23)	-0.019 (-0.44)	0.010 (0.32)	0.0010 (0.85)
e_t	0	0	0	1	-0.383 (-1.16)	-0.288 (-0.27)	0.159 (0.20)	-0.0983** (-3.25)

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

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 β -coefficient is precisely estimated. In contrast, only few α -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 $e_t = 26.4r_t^d$, which is consistent with UIP and stable long-term exchange rate expectations $E_t e_{t+40}$. 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 four cointegrating relations are plotted in Figure 7. The upper panels are just the short-term interest rate and output gap differentials. In the lower left panel the real long-term interest rate differential

can be seen, which is dominated by the pattern of the inflation differential. The only new time series is the last diagram which shows the deviation of the exchange rate from its long-run equilibrium with the bond yield differential.

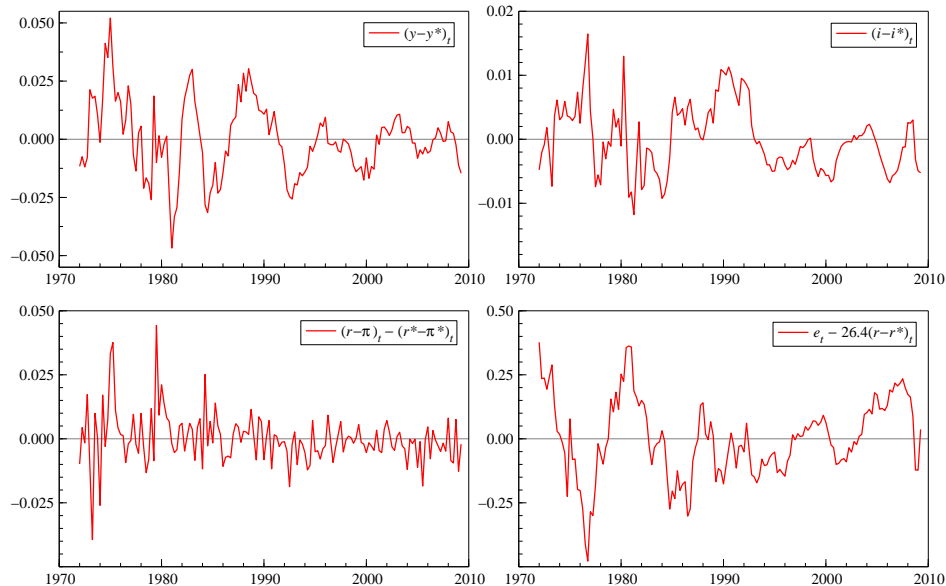


Figure 7 The four cointegrating vectors.

The adjustment matrix $\mathbf{\Pi} = \alpha\beta'$ reported in Table 8 determines how the system reacts to the state of the endogenous variables. The negative signs on the diagonal of the matrix indicate stable self-referencing feedback mechanisms for all variable apart from r_t^d , which in Table 6 was found to be weakly exogenous. The only statistically significant cross-equation feedbacks are the inflation differential and the exchange rate reacting to the bond yield differential, which are driven by the cointegration relations (15) and (16).

Table 8 Combined long-run effects $\mathbf{\Pi} = \alpha\beta'$, standard errors in brackets.

	π_{t-1}^d	y_{t-1}^d	i_{t-1}^d	r_{t-1}^d	e_{t-1}
$\Delta\pi_t^d$	-0.733** (0.150)	0.103 (0.064)	-0.051 (0.203)	0.738** (0.202)	0.0002 (0.0058)
Δy_t^d	-0.098 (0.149)	-0.256** (0.064)	0.106 (0.203)	0.069 (0.201)	0.0011 (0.0058)
Δi_t^d	0.043 (0.051)	0.018 (0.022)	-0.249** (0.069)	-0.007 (0.068)	0.0014 (0.0020)
Δr_t^d	-0.010 (0.031)	0.003 (0.013)	-0.018 (0.043)	-0.018 (0.042)	0.0010 (0.0012)
Δe_t	-0.159 (0.777)	-0.383 (0.331)	-0.288 (1.054)	2.756** (1.046)	-0.0983** (0.0302)

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

3.2.2 Assessing the underlying symmetry of the cointegration relations

The empirical relevance of our model analysis in country differences hinges on the existence of a symmetric long-run equilibrium as stated in (i)-(iv). In order to assess the evidence, we conclude the cointe-

gration analysis by testing for the symmetry of the cointegration relations in two subsystems of the full 9D two-country model. This approach avoids the problems associated with the cointegration analysis of high-dimensional systems. Its validity follows from the invariance of cointegration relations with regard to extensions of the information set. The first subsystem to be considered consists of the UK and US inflation rates, government bond yields and the nominal exchange rate. The statistical analysis was based on a VAR(4) of $(\pi_t^{uk}, \pi_t^{us}, r_t^{uk}, r_t^{us}, e_t)'$ with constant and trend. Two cointegration relations were found and could be identified as follows:

$$(r_t^{uk} - \pi_t^{uk}) - \beta_1(r_t^{us} - \pi_t^{us}) - \delta t \quad \text{and} \quad e_t - \beta_2 r_t^{uk} + \beta_3 r_t^{us}, \quad \text{with} \quad \hat{\beta}_1 = 1.19, \hat{\beta}_2 = 13.6, \hat{\beta}_3 = 16.2.$$

(0.14)
(5.5)
(7.2)

When imposing symmetry across countries on the cointegration space, $\beta_1 = 1$ and $\beta_2 = \beta_3$, the LR-test p-value increases from 0.06 to 0.10 and cointegration vectors corresponding to (iii) and (iv) result:

$$(r_t^{uk} - \pi_t^{uk}) - (r_t^{us} - \pi_t^{us}) - 0.000057t \quad \text{and} \quad e_t - 13.9(r_t^{uk} - r_t^{us}).$$

(0.000019)
(5.6)

Next, consider the subsystem including both output gap series, both short-term nominal interest rates and the nominal exchange rate. For an adequate VAR(4) of $(y_t^{uk}, y_t^{us}, i_t^{uk}, i_t^{us}, e_t)'$ with constant and trend, the trace test suggests a rank of three. With an LR-test p-value of 0.18, the three cointegration relationships can be identified in correspondence to vectors (i) and (ii): each output gap constitutes a cointegration relation, the third cointegration relationship is the trend-stationary short-term interest rate differential, $i_t^{uk} - \beta i_t^{us} + \delta t$ with $\hat{\delta} = 0.000071$, as $\hat{\beta} = 0.88$ with a standard error of 0.14 is not statistically significantly different from one.

3.2.3 Identifying instantaneous causality

The graph-theoretic identification of the contemporaneous causation in the VECM(3) with the identified cointegration vectors in (13) - (16) is based on the residual correlation matrix reported in Table 9. 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.

Table 9 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.

For further investigations of these issues, the correlation matrix in Table 9 is subjected to a graph-theoretical search for instantaneous causal relations. At the 5% and 1% significance levels⁹, the PC algorithm consistently finds only a single undirected edge linking the short and long-term interest rate differentials. This coincides with the only significant residual correlation. As here is one undirected edge, 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:

$$\mathbf{B}^{(1)} = \begin{pmatrix} 1 & \cdot & \cdot & \cdot & \cdot \\ 0 & 1 & \cdot & \cdot & \cdot \\ 0 & 0 & 1 & \cdot & \cdot \\ 0 & 0 & b_{ri} & 1 & \cdot \\ 0 & 0 & 0 & 0 & 1 \end{pmatrix} \quad \text{or} \quad \mathbf{B}^{(2)} = \begin{pmatrix} 1 & \cdot & \cdot & \cdot & \cdot \\ 0 & 1 & \cdot & \cdot & \cdot \\ 0 & 0 & 1 & b_{ir} & \cdot \\ 0 & 0 & \cdot & 1 & \cdot \\ 0 & 0 & 0 & 0 & 1 \end{pmatrix}, \quad (18)$$

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 §3.1. The selection for the design of the contemporaneous effects matrix is based on the statistical properties of the PSVECMs emerging from the *Gets* model reduction process using the conservative strategy of *PcGets* (see §3.2.4 for details) for the two competing contemporaneous causality matrices. As the SVECMs as well as the finally selected PSVECMs are mutually non-nested¹⁰, we select the structure with the best penalized likelihood in form of Akaike and Schwarz information criteria. According to both criteria, the PSVECM based on the candidate design matrix $B^{(1)}$ with an AIC of -36.099 and a SIC of -35.437 dominates the model with the contemporaneous matrix $B^{(2)}$ with AIC = -35.897 and SIC = -35.275 , which is slightly more parsimonious but offers a poorer characterization of the data. As $B^{(1)}$ is found to be the dominant design, we can conclude that it is the short-term interest rate differential, i_t^d , which instantaneously causes the bond yield differential, r_t^d . In the following we focus on the econometric modelling with the selected $B^{(1)}$ matrix.¹¹

3.2.4 The parsimonious structural vector equilibrium correction model

Having specified the SVECM in (12) with the cointegration relations found in §3.2.1 and the contemporaneous relations detected by the PC causal search algorithm in §3.2.3, the model reduction is performed with an automatic general-to-specific model reduction procedure. As the design of $B^{(1)}$ and the values of β^r are given, the model search is limited to the parameters of the short-run dynamics, $\Gamma_1, \dots, \Gamma_3$, and the long-run equilibrium adjustment, α , while it is ensured that the rank of the long-run matrix Π

⁹At 10% two additional edges are detected. The first is an undirected edge between short-term interest rate and inflation rate, the second a directed edge from the exchange rate to the inflation rate. However, both additional edges would drop out during the *PcGets* reduction of the SVECM. Thus, even at 10%, no further contemporaneous causal link would enter the PSVECM.

¹⁰When commencing the model search from a VECM nesting $B^{(1)}$ and $B^{(2)}$, i.e., allowing for both contemporaneous effects, b_{ir} and b_{ri} (note that the *Gets* selection process is largely unaffected by lack of identification in the unrestricted system as shown in Hendry and Krolzig, 2004) and estimating the selected model with Full Information Maximum Likelihood, the short-term interest rate differential affects the long-term interest rate differential contemporaneously, but the bond rate differential is found to be insignificant when included as a contemporaneous effect in the short-term interest rate equation. This is consistent with the notion that monetary policy is directly affecting the bond market but not vice versa.

¹¹The results for the PSVECM with $B^{(2)}$ can be requested from the authors.

is unchanged by the constraints on α . 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 rate as measure of inflation expectations:

$$\begin{aligned} \widehat{\Delta\pi_t^d} = & \quad 0.746 (r_{t-1}^d - \pi_{t-1}^d) + \quad 0.17 y_{t-1}^d - \quad 0.264 \Delta\pi_{t-1}^d - \quad 0.186 \Delta\pi_{t-2}^d \\ & \quad (0.098) \quad \quad \quad (0.039) \quad \quad \quad (0.085) \quad \quad \quad (0.064) \\ & - \quad 0.795 \Delta r_{t-1}^d - \quad 0.044 \text{I1973Q2}_t + \quad 0.0424 \text{I1979Q3}_t, \\ & \quad (0.32) \quad \quad (0.008) \quad \quad \quad (0.008) \end{aligned} \quad (19)$$

$$\hat{\sigma} = 0.00772, \bar{R}^2 = 0.65.$$

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, $\pi_t^{e,d} = \pi_t^e - \pi_t^{*e}$, 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.¹²

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. In combination with the Phillips-curve term in the inflation equation. The output gap puzzle in form of a positive effect of the short-term interest on the output gap enforces the price puzzle phenomenon observed earlier. Also

¹²Note that the coefficients on $r_{t-1}^d - \pi_{t-1}^d$ and Δr_{t-1}^d sum up to approximately zero. Thus equation (19) could be written in a more parsimonious way as:

$$\begin{aligned} \widehat{\Delta\pi_t^d} = & \quad - \quad 0.815 \pi_{t-1}^d + \quad 0.903 r_{t-2}^d + \quad 0.158 y_{t-1}^d - \quad 0.192 \Delta\pi_{t-1}^d \\ & \quad (0.086) \quad \quad (0.151) \quad \quad (0.045) \quad \quad (0.064) \\ & + \text{det.terms}, \quad \hat{\sigma} = 0.00774, \bar{R}^2 = 0.651. \end{aligned}$$

While there is no significant loss in quality of fit, $\chi^2(1) = 0.8429[0.359]$, we will use the systematically derived model in (19) for the impulse response analysis.

a weak short-run Mundell effect can be found:

$$\begin{aligned} \widehat{\Delta y_t^d} = & - 0.156 y_{t-1}^d + 0.184 \Delta \pi_{t-3}^d + 0.805 \Delta i_{t-1}^d + 0.0279 \text{I1974Q3}_t \\ & (0.039) \quad (0.051) \quad (0.19) \quad (0.008) \\ & + 0.0372 \Delta \text{I1979Q2}_t, \quad \hat{\sigma} = 0.00793, \bar{R}^2 = 0.37. \\ & (0.0057) \end{aligned} \quad (20)$$

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

$$\begin{aligned} \widehat{\Delta i_t^d} = & - 0.169 i_{t-1}^d - 0.0528 \Delta \pi_{t-2}^d + 0.339 \Delta r_{t-2}^d - 0.0106 \text{I1977Q1}_t \\ & (0.047) \quad (0.018) \quad (0.11) \quad (0.003) \\ & - 0.0086 \Delta_4 \text{I1980Q4}_t + 0.00101, \quad \hat{\sigma} = 0.0029, \bar{R}^2 = 0.28. \\ & (0.0021) \quad (0.00037) \end{aligned} \quad (21)$$

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.¹³

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 :

$$\begin{aligned} \widehat{\Delta r_t^d} = & 0.299 \Delta i_t^d + 0.118 \Delta i_{t-1}^d - 0.294 \Delta r_{t-1}^d + 0.00776 \Delta \text{I1974Q4}_t \\ & (0.031) \quad (0.037) \quad (0.061) \quad (0.00094) \\ & - 0.00562 \text{I1979Q1}_t - 0.00432 \text{I1980Q1}_t, \quad \hat{\sigma} = 0.0012, \bar{R}^2 = 0.63. \\ & (0.0013) \quad (0.0013) \end{aligned} \quad (22)$$

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:

$$\begin{aligned} \widehat{\Delta e_t} = & - 0.0916 (e_{t-1} - 26.4 r_{t-1}^d) + 1.15 \Delta \pi_{t-1}^d + 1.38 \Delta \pi_{t-2}^d + 1.1 \Delta \pi_{t-3}^d \\ & (0.026) \quad (0.38) \quad (0.44) \quad (0.38) \\ & - 5.82 \Delta r_{t-3}^d + 0.299 \Delta e_{t-1} - 0.218 \Delta e_{t-2} + 0.238 \Delta e_{t-3} + 0.0394, \quad (23) \\ & (1.9) \quad (0.079) \quad (0.08) \quad (0.082) \quad (0.013) \\ & \hat{\sigma} = 0.04679, \bar{R}^2 = 0.23. \end{aligned}$$

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%

¹³We will return to this issue in §4.

adjustment of the exchange rate per quarter, there is more predictability than allowed under UIP.¹⁴ 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.

3.2.5 Testing for the validity and congruency of the model

The efficiency of the single-equation reduction procedure depends on the lack of correlation among the error terms of the model. To investigate the orthogonality assumption we could consider two statistical tests. Firstly, for the exactly identified SVECM, we could test for the joint significance of the nine over-identifying restrictions on the contemporaneous matrix $\mathbf{B}^{(1)}$. Secondly, for the selected PSVECM, we can use the likelihood ratio principle to confront our model with its superimposed orthogonal errors against the alternative of an unrestricted variance covariance matrix. To test the involved 10 independent restrictions on $\mathbf{\Omega}$, we compare the log likelihood of the restricted system estimated by OLS with the unrestricted system estimated with Full Information Maximum Likelihood estimation (FIML). Here we utilize the fact that under the condition of a diagonal variance matrix, $\mathbf{\Omega}$, the FIML estimation under normality collapses to OLS. In support of our previous analysis, the LR test with a test statistic of $\chi^2(10) = 7.250$ and a p-value of [0.702] does not reject the hypothesis of a diagonal covariance matrix. That implies all contemporaneous effects have been captured by the causal search algorithm, the model is valid and can be efficiently estimated by OLS.

The congruency of the model is investigated in Table 10. For the highly reduced model, there are no signs of dynamic misspecification. While there is a huge reduction in non-normality for π_t^d , y_t^d and r_t^d , some issues of non-normality and heteroscedasticity remain. A re-estimation of the equations with GARCH errors appears appealing, but the estimation results are similar to the results in §3.2.4 and therefore not reported here.

Table 10 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 [0.239]	1.099 [0.364]	0.813 [0.543]	1.193 [0.316]	0.495 [0.779]
Normality	9.546** [0.009]	6.917* [0.032]	21.482** [0.000]	3.841 [0.146]	1.462 [0.481]
ARCH 1-4	5.826** [0.000]	1.223 [0.304]	2.604* [0.039]	2.675* [0.035]	2.680* [0.034]
Hetero	3.709** [0.000]	1.867 [0.062]	2.699** [0.006]	1.113 [0.357]	1.029 [0.432]
RESET	6.588* [0.011]	1.610 [0.207]	0.668 [0.414]	0.923 [0.338]	4.257* [0.041]

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

¹⁴The short-run dynamics of (23) could be further simplified as:

$$\begin{aligned} \widehat{\Delta e}_t = & \quad 0.039 \quad - \quad 0.0906 \quad (e_{t-1} - 26.4r_{t-1}^d)_{t-1} \quad - \quad 5.75 \quad \Delta r_{t-3}^d \quad + \quad 1.12 \quad \Delta_3 \pi_{t-1}^d \\ & (0.012) \quad (0.025) \quad (1.85) \quad (0.314) \\ & + \quad 0.304 \quad \Delta e_{t-1} \quad - \quad 0.238 \quad \Delta^2 e_{t-2}, \quad \hat{\sigma} = 0.04691, \quad \bar{R}^2 = 0.237. \\ & (0.077) \quad (0.0607) \end{aligned}$$

The reduction is accepted with $\chi^2(3) = 0.795[0.851]$.

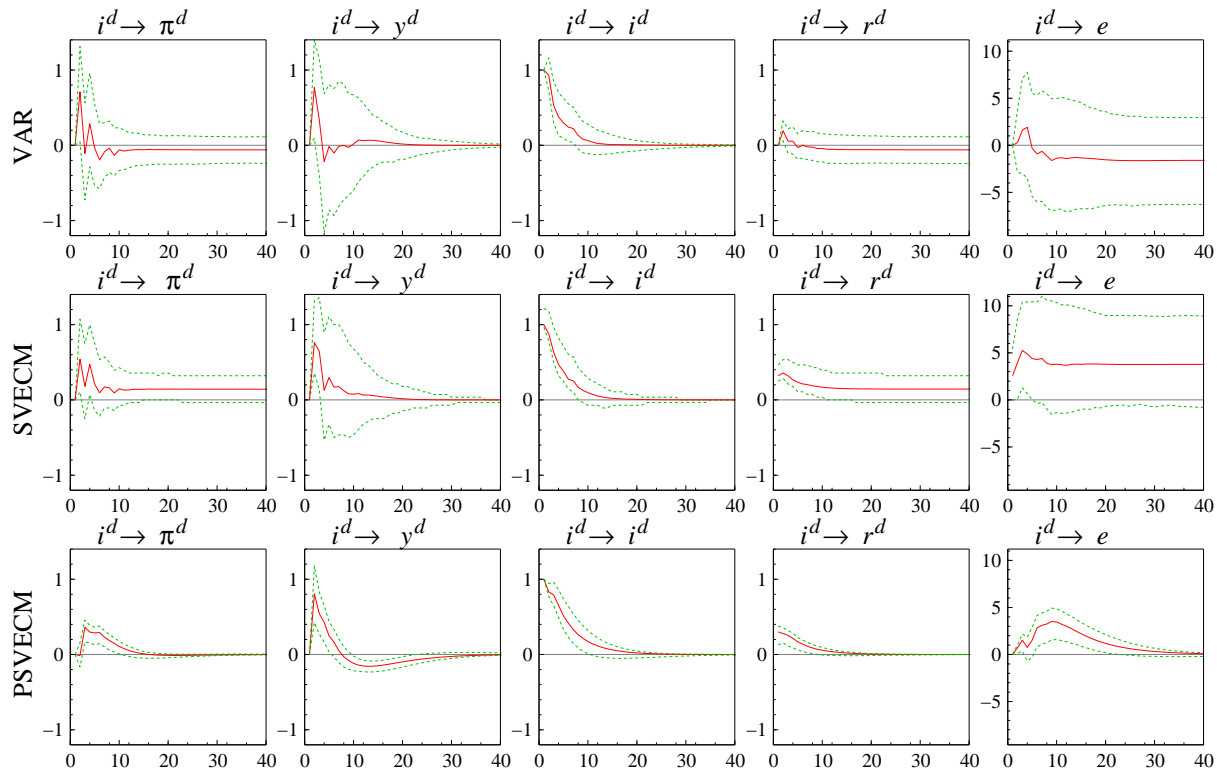


Figure 8 Responses to an asymmetric monetary policy shock in the cointegrated reduced-form VAR, the just-identified and parsimonious SVECM.

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 differential¹⁵, $i_t^d = i_t - i_t^*$. 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 8 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 parsimonious SVECM (bottom panel) presented in §3.2.4. For the just-identified SVECM, the causal order $\pi_t^d \rightarrow y_t^d \rightarrow i_t^d \rightarrow r_t^d \rightarrow e_t$ was imposed to ensure consistency with the TETRAD results in §3.2.3 and making the SVECM nest the PSVECM.

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 parsimonious SVECM

¹⁵It is worth pointing out that, due to the construction of the interest rate variable as the quarterly log return of a 3M treasury bond, a unit shock in i_t^d corresponds to an approximately 400 basis point increase in the 3M treasury bill interest rate differential.

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_t^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.¹⁶

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 disequilibrium with an abnormally high nominal bond yield spread for the first six periods. The increases in the bond yield spread and the resulting undervaluation of the currency relative to its long-run equilibrium, $\nabla e_h - \nabla 26.4r_h^d$, are the main drivers of the steady appreciation of the currency (see the decomposition of ∇e_h in Figure 10, which we will discuss in detail in §4.2). The nominal exchange rate is appreciating for several quarters, before persistently depreciating back to the 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 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 §3.2.3, 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. 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). 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. The robustness of this results for the $\$/\pounds$ exchange rate will be examined in §4.2, where we add a contemporaneous link between short-term interest rate differential and exchange rate into the PSVECM. The key question is whether the delayed overshooting puzzle will emerge *even if* we allow the exchange rate to jump despite the lack of any statistical evidence for this behaviour.

4.2 Allowing the exchange rate to jump

In the following we investigate the case where the exchange rate is jumping in response to unpredicted variations of the short-term interest rate differential. Recall that the contemporaneous effect was found to be statistically insignificant. In order to find a parsimonious model for e_t conditional on the contemporaneous i_t^d (and the past information set), we impose the restriction $b_{ei} \neq 0$ onto the selection process. Forcing the selection of i_t^d , irrespectively of its statistical significance, the conservative strat-

¹⁶In §4.5, we show that the possibly resulting identification issues do not affect the delayed overshooting finding.

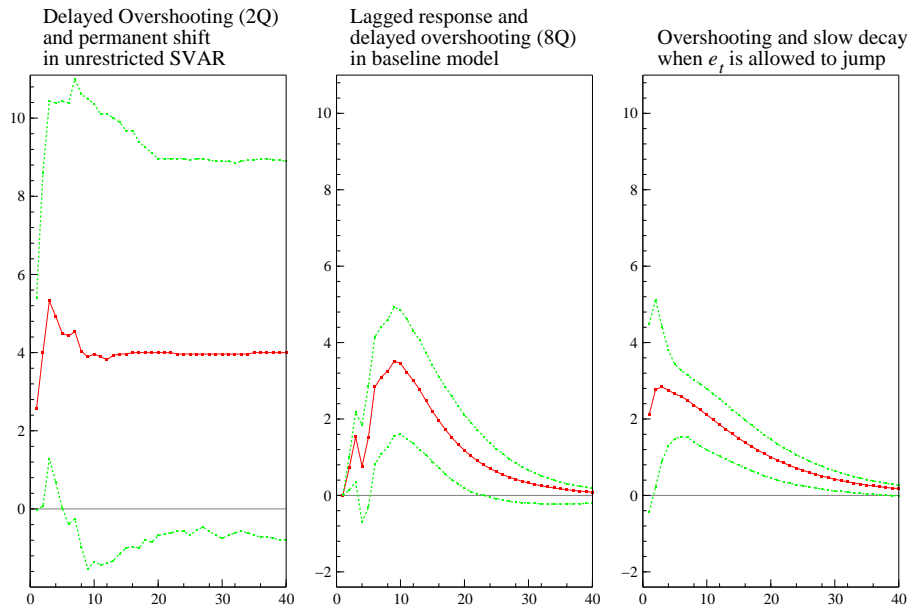


Figure 9 Effects of a monetary policy shock: response of the exchange rate to a one percentage point increase in the 3-month interest rate differential (with 95% confidence interval).

egy of *PcGets* delivers the following exchange rate equation:

$$\begin{aligned} \widehat{\Delta e}_t = & -0.0803 (e_{t-1} - 26.4r_{t-1}^d) + 2.11 \Delta i_t^d + 0.258 \Delta e_{t-1} \\ & (0.027) \qquad\qquad\qquad (1.23) \qquad\qquad\qquad (0.081) \\ & -0.182 \Delta e_{t-2} + 0.0339, \quad \hat{\sigma} = 0.050, \bar{R}^2 = 0.12. \\ & (0.082) \qquad\qquad\qquad (0.013) \end{aligned} \tag{24}$$

The new equation has changed greatly when compared to the baseline model in (23) 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. Figure 9 seeks 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 SVAR and the baseline model. The panel on the left displays the response of the exchange rate in the exactly identified SVAR. 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 parsimonious SVAR. The second panel corresponds to 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. Across models, there is strong evidence for delayed overshooting over the sample period.

Figure 10 gives some insight into the dynamic adjustment processes caused by the shock in i_t^d by

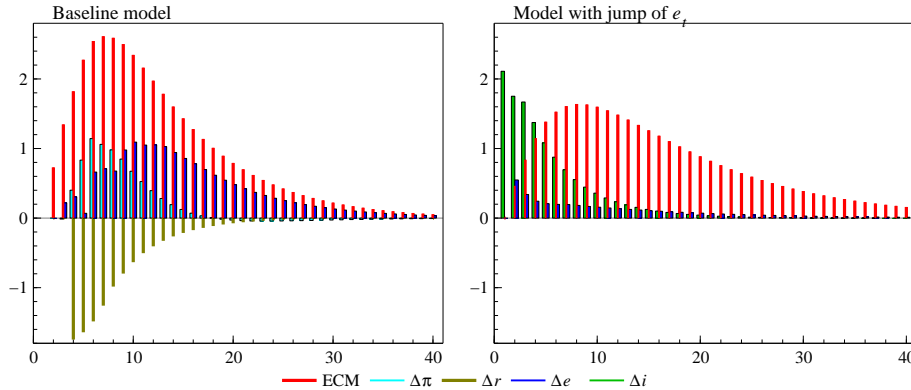


Figure 10 Decomposition of the response of the exchange rate to a monetary policy shock: Direct interest rate effects (Δi^d), equilibrium correction (ECM), momentum (Δe) as well as spill-overs from inflation ($\Delta \pi^d$) and bond yield changes (Δr^d). The baseline model is depicted in the left, the alternative where e_t is allowed to jump on the right panel.

decomposing the response of the exchange rate to a monetary policy shock into direct interest effects, which are only relevant when e_t is allowed to jump, equilibrium correction, momentum (Δe) as well as spill-overs from inflation ($\Delta \pi^d$) and bond yield changes (Δr^d). It can be seen that the equilibrium adjustment is the major force behind the delayed overshooting phenomenon as it induces a hump-shaped response of the exchange rate. In contrast, the direct interest rate effects in the model with an instant reaction of exchange rate are largely consistent with conventional overshooting.

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. As in Figure 9, we 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 11 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:

$$\nabla e_h^{\text{UIP}} = \nabla e_T + \sum_{s=h}^{T-1} \nabla i_s^d \quad (25)$$

where in the plots above $T = 150$ was used.¹⁷ 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 UIP with the ex-ante one-period excess return series:

$$\nabla \xi_h = \nabla i_h^d + \Delta \nabla e_{h+1}. \quad (26)$$

The plots reveal, consistently among models, excess returns for UK treasury bonds after a tightening of

¹⁷As the exchange rate converges much earlier, the results are insensitive to the choice of T .

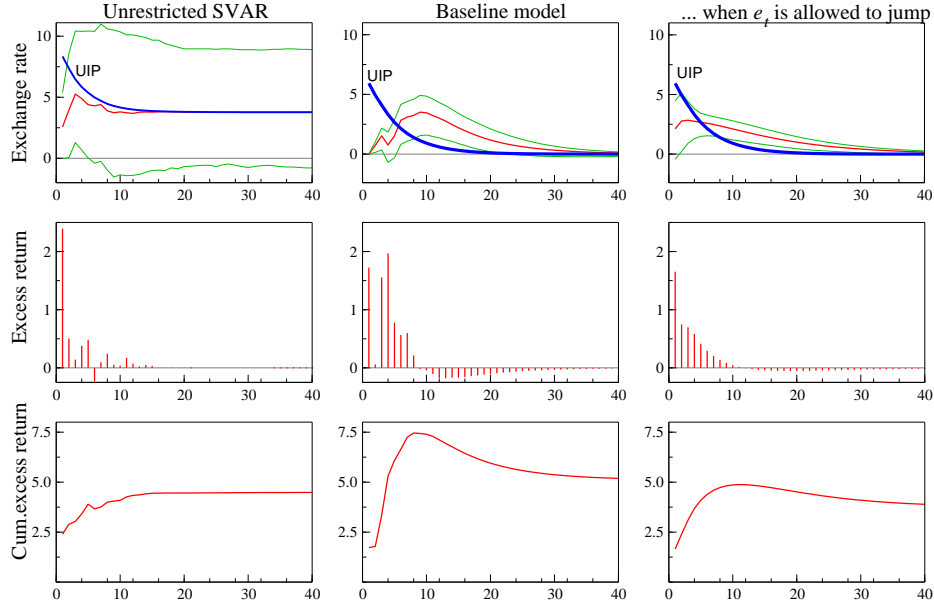


Figure 11 Effects of a monetary policy shock: responses of exchange rate, ∇e_{t+h} , the excess return, $\nabla \xi_h$ and cumulative excess return, $\nabla \xi_{0,h}$, to a one percentage point increase in i_t^d .

Bank of England policy. In the bottom panels, the cumulated excess returns,

$$\nabla \xi_{0,h} = \sum_{j=0}^h \nabla \xi_j = \sum_{j=0}^h \nabla i_j^d + \Delta_h \nabla e_{h+1}, \quad (27)$$

are plotted. Over the first two years excess returns of up to 7.5 percent are observed constituting major violations of the UIP hypothesis. Altogether, the statistical results discussed in this paper strongly support the presence of a delayed overshooting puzzle for the $\$/\pounds$ exchange rate.

4.4 Exploring asymmetries in the exchange rate response to UK versus US policy shocks

Due to the underlying symmetry of the model, an increase in short-term interest rate differential, $i_t^d = i_t^{uk} - i_t^{us}$, could either be associated with an asymmetric tightening in the stance of monetary policy of the Bank of England or an asymmetric loosening by the Fed. To gain insight into possible asymmetries in the response of the exchange rate to policy shocks in the UK versus the US, we consider here an extension of the baseline model, where the short-term interest differential, i_t^d , is replaced by the UK and the US interest rate. The resulting system consisting of the six variables, $(\pi_t^d, y_t^d, i_t^{us}, i_t^{uk}, r_t^d, e_t)'$, embraces both country differences and country variables. This unbalanced nature of the 6D system limits the scope for congruent econometric modelling. Having previously established evidence for the symmetry of cointegration space, we proceed by imposing the cointegration relations in (13) and (14) of the model in country differences discussed in §3.2.1. It is worth noting that the model setup implies that the two short-term interest rates are cointegrated: while their differential i_t^d is stationary, the interest rates themselves are integrated.

Commencing from a VECM(3) of the 6D system with the four cointegration relationships of the baseline model superimposed, the analysis continued by following steps (iii)-(v) of the modelling procedure outlined in §3.1. The graph-theoretic search by TETRAD for the contemporaneous causal relations

results in a DAG corresponding to a contemporaneous matrix of the form:

$$\mathbf{B} = \begin{pmatrix} 1 & \cdot & \cdot & \cdot & \cdot & \cdot \\ 0 & 1 & \cdot & \cdot & \cdot & \cdot \\ 0 & b_{i^{us}y} & 1 & \cdot & \cdot & \cdot \\ 0 & 0 & b_{i^{uk}i^{us}} & 1 & \cdot & \cdot \\ 0 & 0 & b_{r^{i^{us}}} & b_{r^{i^{uk}}} & 1 & \cdot \\ 0 & 0 & 0 & 0 & 0 & 1 \end{pmatrix}. \quad (28)$$

In the temporary macroeconomic equilibrium, the US short-term interest rate, which is instantaneously reacting to the output gap differential, affects the UK short-term interest rate but not vice versa. Both short-term interest rates affect the bond yield differential. However, there is no impact effect on the exchange rate implying again delayed overshooting. To preserve the link between the short-term interest rates, which is found by TETRAD even at the 1% significance level, the model reduction was performed with the liberal strategy of *PcGets*.

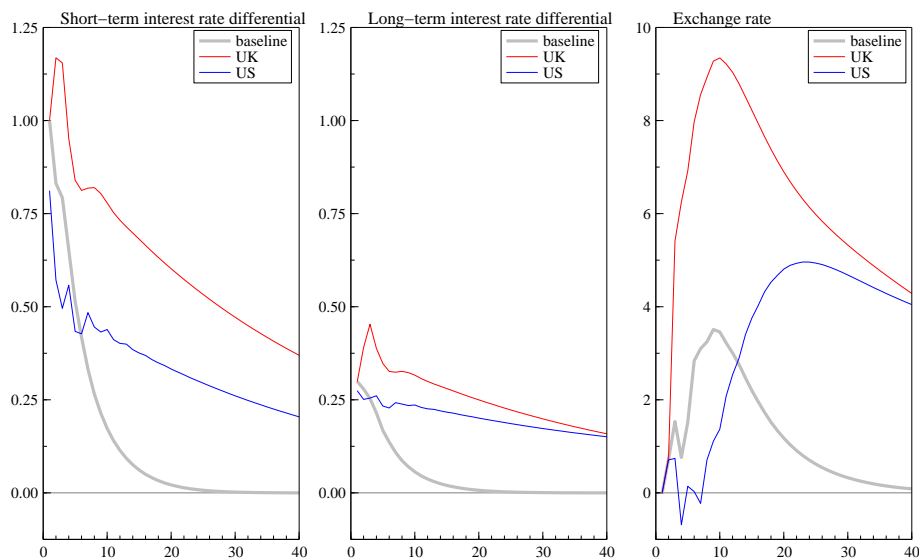


Figure 12 Responses of interest rate differentials and the exchange rate to monetary policy shocks in UK (unit increase in i^{uk}) and in the US (unit decrease in i^{us}) in the 6D-PSVECM in comparison to the baseline impulse response function (unit increase in i^d).

The results of an impulse response analysis of the selected PSVECM are presented in Figure 12. Focussing on the short- and long-term interest rate differentials as well as the exchange rate, we compare the responses to (i) a unit short-term interest rate rise in the UK and (ii) a unit decrease in the US rate to the impulse response function of the baseline model. In line with the latter, a delayed overshooting of the exchange rate can be observed for both the UK and US monetary policy shock. However, there are pronounced quantitative differences in the exchange rate response. Most strikingly, the $\$/\mathcal{L}$ exchange rate reacts much stronger with regard to the shock in the UK than to the US shock. This result is directly linked to the contemporaneous causality in the system. Since a monetary policy shock in the US instantaneously affects the UK money market, the reaction of both short-term and long-term interest differentials are much deflated when compared to the idiosyncratic UK shock. When discounting the long-run results in light of the mixed nature of the system with its two short-term interest rates and country differences otherwise, the baseline model is a good approximation of the delayed overshooting

of the exchange rate in response to a mixture of UK and US interest rate shocks. In practice, we would expect UK shocks to have a bigger impact than their US counterparts.

4.5 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. The adverse reaction of inflation and output gap in the quarters following a short-term interest rate shock raises serious questions regarding the identification of the monetary shock. It is perfectly conceivable that the shock we previously analysed is not an autonomous change in the stance of monetary policy but rather the reaction of monetary authorities to the anticipation of future inflationary shocks. Indeed, we proposed to interpret the role of the bond yield differential in equation (19) as that of inflation expectations in an augmented Phillips curve. A lack of identification would, in turn, bring into question the presence of a delayed overshooting puzzle, particularly the extent to which this result has been contaminated by the price and output gap puzzles.

Following Sims (1992), the standard procedure of establishing robustness of the model results with regard to the exchange rate behaviour would involve the integration of proxies of future inflation (e.g., commodity prices) or inflation expectations (i.e., survey information) into the system. However, given the lack of available data¹⁸ for inflation expectations in the UK over the full sample period and the conceptual limits on using commodity prices as a measure of future inflation *divergences* of the two countries, this route is not an option here. It is also worth noting that, as shown Demiralp *et al.* (2009) for the US price puzzle using methodologically closely related approach, the inclusion of measures of future inflation may not resolve the price puzzle after all.

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. We start by evaluating the relative importance of the price and output puzzle for the emergence of the delayed overshooting puzzle. In order to dissect the contribution of the price and output puzzles for the impulse response of the exchange rate with regard to the monetary policy shock, we run a thought experiment in which the feedback chains from the real sector of the economy, $(\pi_t^d, y_t^d)'$, to the financial markets in form of $(i_t^d, r_t^d, e_t)'$ are broken. In other words, we rerun the impulse response analysis while fixing the inflation and output gap differentials. The result of this experiment are shown in the upper panels of Figure 13. We can see that the exchange movements are qualitatively unchanged when discarding the feedback from the real economy. Actually, without interaction of e with the π^d and y^d , the peak response of the exchange rate is even further delayed than in the PSVECM. When the exchange is allowed to jump, the result is also quantitatively indistinguishable from the baseline scenario. These findings demonstrate that the price puzzle does not invalidate the delayed overshooting results of the paper. As the price puzzle does not cause the delayed overshooting puzzles, the presence of a delayed overshooting puzzle should be invariant to model variations not exhibiting the price and output gap puzzles.

To investigate the identification issue further, we shift our focus onto two theory-driven model interventions designed to tackle the sources of the price and output gap puzzles by revising the links from

¹⁸While the Survey of Professional Forecasters would provide quarterly one-year-ahead inflation expectations for the US, no such time series exists for the UK over the whole sample period. For the subsample from 1985Q1 to 2009Q2, we have been able to create a rough measure of the country differences in expected inflation by combining the US survey data with an implied expected inflation series derived from inflation-indexed bond data for the UK, which one of the first OECD countries to issue this type of government bonds in 1981. The US followed in 1997 with the introduction of Treasury Inflation-Protected Securities. We are using Bank of England data for the implied four-year-ahead inflation rate from the zero coupon yields of inflation-indexed bonds beginning in 1985.

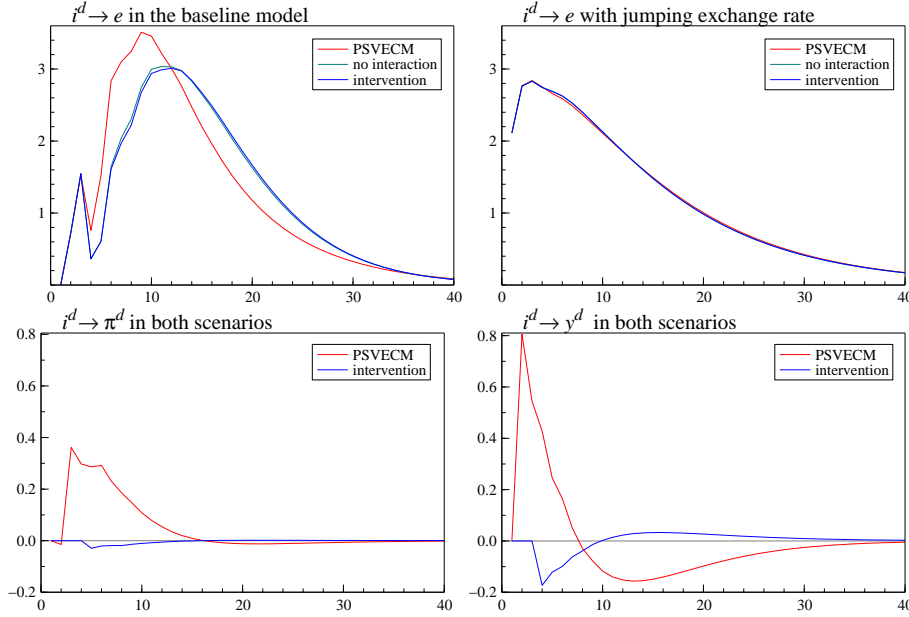


Figure 13 Impulse responses of the inflation differential, the output gap differential and of the exchange rate to a monetary policy shock in the baseline model, the model without feedback from the real sector of the economy, and the intervention model with a sign-restricted IS curve and an adaptive inflation expectation mechanism.

the money and bond markets to the real economy:

- (i) To overcome the output gap puzzle, caused by the positive sign of the Δi_{t-1}^d term in the output gap equation in (20), we first re-parameterize the unrestricted SVECM equation for y_t^d and then rerun the model search under sign restrictions: In the reformulated unrestricted equation for the output gap differential, the interest and inflation rates only enter as *real* interest rate differentials, i.e., $\Delta(i^d - \pi^d)_{t-j}$ and $\Delta(r^d - \pi^d)_{t-j}$ for $j = 1, \dots, 3$. The model search is constrained to include real interest rate terms only in case of a negative sign of the respective total effect, i.e., the sum of the corresponding coefficients. With the sign restrictions imposed (see Krolzig, 2008, for details), *PcGets* selects the following equilibrium-correction equation for the output gap differential:

$$\begin{aligned} \widehat{\Delta y_t^d} = & - 0.128 y_{t-1}^d - 0.173 (\Delta i_{t-3}^d - \Delta \pi_{t-3}^d) + 0.0210 \Delta_4 I1974 Q3_t \\ & \quad (0.040) \quad \quad (0.053) \quad \quad (0.006) \\ & + 0.0346 \Delta I1979 Q2_t, \quad \hat{\sigma} = 0.00824, \bar{R}^2 = 0.32. \\ & \quad (0.0059) \end{aligned} \quad (29)$$

As in the original equation, it is the short-term interest rate that drives the output gap but now as the real rate and with the expected sign. Due to the link with the inflation rate equation via the Phillips curve term, the change in sign also contributes to the circumvention of the price puzzle.

- (ii) To resolve the price puzzle, we have to deal with the positive sign of the bond yield differential in the equilibrium-correction equation for the inflation differential. As pointed out in the discussion of (19), the long-term interest rate acts here as proxy for the missing forward-looking inflation expectations. This nexus is the main root for the price puzzle. To facilitate the identification of an autonomous monetary policy shock, it is essential to break it up. For the impulse response analysis, we therefore replace the bond yield differential by a new variable measuring the country differences in expected inflation. Thus, the equilibrium correction term, $r_{t-1}^d - \pi_{t-1}^d$, is represented by the deviation of the expected rate of inflation from the realized one: $\pi_{t-1}^{e,d} - \pi_{t-1}^d$. In

view of the highlighted lack of suitable survey data, we close the model by an adaptive learning mechanism:

$$\Delta\pi_t^{e,d} = \alpha(\pi_{t-1}^d - \pi_{t-1}^{e,d}). \quad (30)$$

For the calibration of α , we constructed a $\pi^{e,d}$ time series using US survey and UK indexed bond data from 1985Q2 to 2009Q2 (see footnote 18) and estimated (30) by OLS as $\hat{\alpha} = 0.07$, which (0.025)

will be used in the following.

The results of the impulse response analysis with the intervention superimposed on the model are shown in Figure 13. The reactions of the inflation and output gap differential to a monetary policy shock, depicted in the lower panels, clearly demonstrate that the interventions were sufficient to resolve the price puzzle. Having dissolved the concerns about the identification of the monetary policy shock, we can now focus on the delayed overshooting results. As can be seen in the upper left panels, the behaviour of the exchange rate follows closely the impulse response function of the empirical baseline model developed in this paper. This is consistent with our earlier result that the existence of the delayed overshooting puzzle is unrelated to the price and output puzzle observed in the same model. Indeed the spillovers from the adverse short-term effects of a monetary policy on the inflation and output gap differentials on the exchange rate are neglectable. Thus, our 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 but also reproduces our earlier findings with regard to reaction to the response of the exchange rate qualitatively and, when the exchange rate is allowed to jump, quantitatively.

5 Conclusion

In our investigation into the presence of a ‘delayed overshooting puzzle’ in the response of the $\$/\mathcal{L}$ 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, $e_t = 26.4r_t^d$, which is consistent with UIP and stationary long-term exchange rate expectations, $E_t e_{t+40}$. 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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