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ON THE MACROECONOMIC CAUSES OF EXCHANGE RATES VOLATILITY

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On the macroeconomic causes of exchange rates volatility

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Abstract

What are the causes of exchange rate volatility? When second moments implications of theories of exchange rates determination are considered, long-term fundamental linkages between macroeconomic and exchange rate volatility can be envisaged. Moreover, as the exchange rate is an important determinant of aggregate demand, bidirectional causality should be expected. The results of the paper support the above intuitions pointing to important linkages and trade-offs relating exchange rate and macroeconomic volatility, with causality direction stronger from macroeconomic volatility to exchange rate volatility than the other way around. In particular, with a long-term perspective, Friedman (1953) conclusions on the macroeconomic sources of exchange rates instability and the impossibility of eliminating systemic volatility find full support in the empirical findings.

Key words: exchange rates volatility, macroeconomic volatility, long memory, structural change, fractional cointegration, cobreaking, fractionally integrated factor vector autoregressive model, G-7 area.

JEL classification: C22, F31, E44, E52.

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1 Introduction

What are the economic causes of exchange rates volatility? Despite the relevance of the question, surprisingly little work on this issue has been carried out in the literature so far. Recent contributions have pointed out that macroeconomic volatility is not an important source of exchange rates volatility for G-7 countries, since little evidence of exchange rate regime dependence is found in macroeconomic volatility, i.e., differently from exchange rates volatility, macroeconomic volatility does not tend to be higher in regimes of floating rates than in regimes of fixed rates. Moreover, little evidence of volatility conservation can also be found, i.e., apart from output volatility, no trade-offs between macroeconomic and exchange rates volatility can be found, suggesting that fixing exchange rates would not lead to higher macroeconomic volatility in general: excess volatility simply disappears (Flood and Rose, 1995, 1997).[1]

The above findings are however not inconsistent with second moments implications of fundamental models of exchange rate determination, predicting a linkage between exchange rate and macroeconomic volatility, for two main reasons. Firstly, once sticky prices are allowed for, only a weak response of macroeconomic variables to changes in exchange rates regimes and volatility can be expected in the short- to medium-term (Duarte, 2006). Secondly, other determinants than macroeconomic fundamentals may be important for exchange rate volatility in the short- to medium-term, as for instance excessive speculation (Flood and Rose, 1999), heterogeneous agents (Muller et. al., 1997), overshooting effects related to information problems (Faust and Rogers, 2003) and information flows (Andersen and Bollerslev, 1997), which, moreover, are responsible for the strong persistence of volatility shocks.[2] Yet, while the above factors may dominate macroeconomic shocks in the short- to medium-term, nothing prevent that in the long-term the role of fundamental (macroeconomic) volatility can be reasserted. Interestingly, no empirical evidence concerning the long-term has been provided in the literature so far, since all the available empirical evidence concerns the short-

[1] Similarly, Baxter and Stockman (1989) had previously found little evidence that macroeconomic volatility or trade flows are influenced by exchange rate regimes. Differently, Hutchinson and Walsh (1992) and Bayoumi and Eichengreen (1994) have found evidence consistent with the insulation properties of a flexible exchange rate regime, while both Arize et al. (2000) and Rose (2000) have found evidence of significant negative, yet small, effects of exchange rate volatility on trade flows.

[2] According to Muller et al. (1997), the interaction in the market of agents with different time horizon leads to long memory in exchange rates volatility. Differently, Andersen and Bollerslev (1997) explain long memory in exchange rates volatility as the consequence of the aggregation of a large number of news information arrival processes.

to medium-term. Hence, assessing long-term linkages between exchange rate and macroeconomic volatility is actually the scope of the paper.

The paper contributes to the literature under two respects.

Firstly, it provides empirical evidence on the linkage between exchange rates and macroeconomic volatility for the G-7 countries for both the short- to medium-term and the long-term, also focusing on the most recent float period (1980-2006), which has almost entirely been neglected in the literature so far.

Secondly, differently from the descriptive analysis carried out in Flood and Rose (1995, 1997), accurate modelling of the persistence properties of the data has been carried out in the framework of a new fractionally integrated factor vector autoregressive (FI-F-VAR) model. This latter model allows to investigate linkages across variables and countries involving both common deterministic and stochastic components, consistent with recent findings in the literature pointing to the presence of both structural change and stationary long memory in the volatility of financial asset returns and macroeconomic variables.[3] Hence, both long-term (cobreaking) and medium-term (fractional cointegration) relationships can be investigated in the current framework, controlling for short-term dynamic linkages as well. Moreover, conditioning is made relatively to a very large information set, since the analysis is carried out considering the entire G-7 macroeconomic structure jointly, allowing therefore for a fine control of the interrelations occurring across countries, currencies and macroeconomic factors.

The findings of the paper are clear-cut. Evidence of significant long-term linkages and trade-offs between macroeconomic and exchange rate volatility, particularly involving output and inflation volatility, and money growth volatility at a lower extent, has been found. Moreover, although causality is bidirectional, the linkage is much stronger from macroeconomic volatility to exchange rate volatility than the other way around. Interestingly, significant cross-country interactions have been found as well, i.e. foreign countries macroeconomic volatility may also be important for long-term exchange rate volatility. This latter finding is consistent with the evidence of global real and nominal dynamics found for the G-7 countries in the literature. Overall, relatively to previous work in the literature, the paper provides encompassing evidence for the short- to medium term, and new evidence for the so far neglected long-term case. Moreover, indirect empirical support is then found

[3]For financial asset returns, see for instance Granger and Hyung (1999) and Mikosch and Starica (1998) for structural breaks; Bailie et al. (1996), Andresen et al. (1997) for long memory; Lobato and Savin (1998), Morana and Beltratti (2004), Beine and Laurent (2000), Baillie and Morana (2007), Martens et al. (2004) for both features. For macroeconomic volatility see Beltratti and Morana (2006).

for fundamental models of exchange rate determination in the long-term.

Finally, two key policy implications follows from the results of the paper, consistent with the seminal views of Friedman (1953) on the case for flexible exchange rates: in a long-term perspective, focusing on macroeconomic stability may indeed be important to reduce excess exchange rates volatility; moreover, systemic volatility cannot be eliminated by fixing exchange rates, since the latter can only come at the cost of macroeconomic instability.

After this introduction the paper is organized as follows. In sections two the econometric methodology is introduced. Then, in section three the data are presented, while estimation is carried out in section four. Finally, in section five conclusions are drawn.

2 Econometric methodology

Consider the following fractionally integrated factor vector autoregressive (FI-F-VAR) model

$$
\begin{aligned}
(I - C(L))\,(x_t - \Lambda_\mu \mu_t - \Lambda_f f_t) &= v_t \quad (1)\\
D(L) f_t &= \eta_t, \quad (2)
\end{aligned}
$$

where x_t is a n-variate vector of stationary long memory processes ($0 < d_i < 0.5$, $i = 1, ..., n$)[4], f_t is a r-variate vector of stationary long memory factors, μ_t is an m-variate vector of common break processes, v_t is a n-variate vector of zero mean idiosyncratic i.i.d. shocks, η_t is a r-variate vector of common zero mean i.i.d. shocks, $E[\eta_t v_{is}] = 0$ all i, t, s, Λ_f and Λ_μ are $n \times r$ and $n \times m$, respectively, matrices of loadings, $C(L)$ is a matrix of polynomials in the lag operator of order p with all the roots outside the unit circle, i.e. $C(L) = C_1 L + C_2 L^2 + ... + C_p L^p$, C_i $i = 1, .., p$ is a square matrix of coefficients of order n, and $D(L) = diag\left\{(1-L)^{d_1}, (1-L)^{d_2}, ..., (1-L)^{d_r}\right\}$ is a diagonal matrix in the polynomial operator of order r. The fractional differencing parameters d_i, as well as the μ_t and f_t factors, are assumed to be known, although they need to be estimated. This is not going to affect the asymptotic properties of the estimator, since consistent estimation techniques are available for all the parameters and unobserved components.[5]

[4]See Baillie (1996) for an introduction to long memory processes.

[5]Alternatively, relying on the surplus lag theory presented in Bauer and Maynard (2006), estimation could be carried out without fractional differencing the series. In the latter case $D(L)$ would simply be a standard polynomial matrix in the lag operator, with all the roots outside the unit circle.

2.1 The fractional VAR form

By taking into account the binomial expansion[6] and substituting (2) into (1), the infinite order vector autoregressive representation for the factors f_t and the series x_t can be written as

$$\left[\begin{array}{c} f_t \\ x_t - \Lambda_\mu \mu_t \end{array}\right] = \left[\begin{array}{cc} \Phi(L) & 0 \\ E(L)\Phi(L) & C(L) \end{array}\right] \left[\begin{array}{c} f_{t-1} \\ x_{t-1} - \Lambda_\mu \mu_{t-1} \end{array}\right] + \left[\begin{array}{c} \varepsilon_{f_t} \\ \varepsilon_{x_t} \end{array}\right], \quad (3)$$

where $D(L) = I - \Phi(L)$, $\Phi(L) = \Phi_0 L^0 + \Phi_1 L^1 + \Phi_2 L^2 + ...$, Φ_i, $\forall i$, is a square matrix of coefficients of dimension r, $E(L) = [I - C(L)]\,\Lambda_f$,

$$\left[\begin{array}{c} \varepsilon_{f_t} \\ \varepsilon_{x_t} \end{array}\right] = \left[\begin{array}{c} I \\ \Lambda_f \end{array}\right] \eta_t + \left[\begin{array}{c} 0 \\ v_t \end{array}\right],$$

with variance covariance matrix

$$E\varepsilon_t \varepsilon_t' = \Sigma_\varepsilon = \left[\begin{array}{cc} \Sigma_\eta{}' & \Sigma_\eta{}'\Lambda_f' \\ \Lambda_f \Sigma_\eta{}' & \Lambda_f \Sigma_\eta{}'\Lambda_f' + \Sigma_v \end{array}\right],$$

where $E\eta_t\eta_t' = \Sigma_\eta$ and $Ev_t v_t' = \Sigma_v$.

2.2 Estimation

The estimation problem may be written as follows

$$\min_{\mu_1,\ldots,\mu_m, f_1,\ldots,f_r, \Lambda_f, \Lambda_\mu, C(L),} T^{-1} \sum_{t=1}^{T} \varepsilon_{x_t}' \varepsilon_{x_t},$$

where $\varepsilon_{x_t} = [I - C(L)L]\,[x_t - \Lambda_\mu \mu_t] - [E(L)\Phi(L)L]\, f_t$. Yet, since the infinite order representation cannot be handled in estimation, a truncation to a suitable large lag for the polynomial matrices $\Phi(L)$ is required. Hence, $\Phi(L) = \sum_{j=0}^{p} \Phi_j L^j$.

The estimation problem can then be solved following an iterative process, consisting of the following steps.

Step 1: initialization. Conditional to the presence of structural breaks and long memory in the series investigated[7], an initial estimate of the unobserved common deterministic and long memory factors can be obtained by

[6] $(1-L)^d = \sum_{j=1}^{\infty} \rho_j L^j$, $\rho_j = \frac{\sum_{k=0}^{\infty} \Gamma(j-d) L^j}{\Gamma(j+1)\Gamma(-d)}$, where $\Gamma(\cdot)$ is the gamma function.

[7] See the next section for details about the estimation of the break processes and the order of fractional differencing.

decomposing the series into their break process (b_t) and long memory components (l_t), i.e. $x_t = b_t + l_t$, and applying principal components analysis (PCA) to each set of components. Several possibilities are available at this stage, since the analysis could be carried out using different sub sets of components, in the place of the entire set, according to economic interpretability.

Considering for simplicity the entire data set and the case of equal fractional differencing parameters across series[8], the procedure is then as follows.

Firstly, the common break processes are estimated by means of PCA implemented using the estimated break process $\hat{b}_t$, yielding an estimate of the $n \times m$ loading matrix $\hat{\Lambda}_\mu = \hat{A}\hat{\Lambda}_b^{1/2}$, where $\hat{\Lambda}_b$ is the diagonal matrix of the estimated non zero eigenvalues of the reduced rank variance-covariance matrix of the (estimated) break processes $\hat{\Sigma}_{\hat{b}}$ (rank $m < n$), $\hat{A}$ is the matrix of the associated orthogonal eigenvectors, and $\hat{\mu}_t = \hat{\Lambda}_b^{-1/2}\hat{A}'\hat{b}_t$ is the $m \times 1$ vector of the standardized ($\hat{\Sigma}_{\hat{\mu}} = I_m$) principal components or common break processes.[9]

Secondly, the common long memory components can be obtained by means of PCA implemented using the estimated break-free processes $bf_t = x_t - \hat{b}_t$, yielding an estimate of the $n \times r$ loading matrix $\hat{\Lambda}_f = \hat{B}\hat{\Lambda}_{bf}^{1/2}$, where $\hat{\Lambda}_{bf}$ is the diagonal matrix of the estimated non zero eigenvalues of the reduced rank variance-covariance matrix of the (estimated) break-free processes $\hat{\Sigma}_{bf}$ (rank $r < n$), $\hat{B}$ is the matrix of the associated orthogonal eigenvectors, and $\hat{f}_t = \hat{\Lambda}_{bf}^{-1/2}\hat{B}'\hat{b}_t$ is the $r \times 1$ vector of the standardized ($\hat{\Sigma}_{\hat{f}} = I_r$) principal components or common long memory processes.

Step 2: starting the iterative procedure. Conditional to the estimate of the fractional differencing parameter, the lag truncation order, and the estimate of the deterministic and stochastic factors, the iterative estimation procedure is started by computing a preliminary estimates of the $C(L)$ polynomial matrix by means of the OLS estimation of the VAR model for the break and long memory free variables $x_t - \hat{\Lambda}_\mu\hat{\mu}_t - \hat{\Lambda}_f\hat{f}_t$.

[8] The case of equal fractional differencing parameters is consistent with the Engle and Granger (1987) definition of fractional cointegration, and relevant for the empirical analysis carried out in this paper. Generalizations considered in Marinucci and Robinson (2001) and Robinson and Yajima (2002) allow for the case of subsets of variables showing different orders of integration. If some of the higher order variables cointegrate within their own subset, the consequent order reduction may allow fractional cointegration to concern variables belonging to different subsets. The proposed approach allows to handle this latter case, by considering subsets of variables characterized by the same order of integration.

[9] Bierens (2000) has proposed a similar approach for the estimation of the common non linear deterministic components from a set of estimated individual non linear deterministic components. Yet, details cannot be found in the published version of his paper.

Then, a new estimate of the stochastic factors is obtained as the first r principal components of $\hat{\varepsilon}_{x_t} + \hat{E}(L)\hat{\Phi}(L)L\hat{f}_t = \left[I - \hat{C}(L)L\right][x_t - \hat{\Lambda}_\mu \hat{\mu}_t]$ and, conditional to the new estimated stochastic factors, a new estimate of the factor loading matrix Λ_f can be obtained by OLS estimation of the block of equations corresponding to x_t in (1).

Next, conditional to the new stochastic factors and factor loading matrix, the new deterministic common factors can be obtained as the first m principal components of the set of break processes extracted from the long memory free series $[x_t - \hat{\Lambda}_f \hat{f}_t]$. Then, the new factor loading matrix $\hat{\Lambda}_\mu$ is obtained from the normalized eigenvectors associated with the non zero eigenvalues of the variance covariance matrix for the new normalized eigenvectors.

The procedure is then iterated until convergence.

Step 3: the restricted estimation of the full model. Once the final estimates of f_t and μ_t are available, by employing the estimate of the $\Phi(L)$ matrix and the final estimates of the Λ_f, Λ_μ and $C(L)$ matrices, the restricted VAR in (3) can be estimated.[10]

The above approach can be understood as a generalization of the factor VAR approach proposed by Stock and Watson (2005), allowing for both deterministic and long-memory stochastic factors. Stock and Watson (2005) have provided details about the asymptotic properties, i.e. consistency and asymptotic normality, of the estimation procedure for the case of I(0) variables. Albeit no theoretical results are currently available for long memory processes, Monte Carlo evidence provided in Morana (in press) fully supports the use of the principal component analysis (PCA) for long memory processes.[11] Moreover, since the fractional differencing parameter can be consistently estimated, the asymptotic properties of the estimation method are not affected by the conditioning to the initial estimate of the persistence parameter. See the Appendix for details concerning the identification of the common and idiosyncratic shocks.

[10] It should be noted that the estimation of the short term structure of the model, i.e. the VAR part, is carried out conditional to the estimation of the fractional differencing parameter, which can be accurately estimated by semiparametric methods. This explains why an update of the $\Phi(L)$ matrix is never computed. Alternatively, at step 3 the final estimate of the $\Phi(L)$ matrix could be computed by OLS estimation of the VAR equations corresponding to the block of factor equations.

[11] Theoretical results also validate the use of PCA in the case of both weakly and strongly dependent processes. See for instance Bai (2003, 2004) and Bai and Ng (2004).

3 Data and modelling issues

Instead of considering Italy, Germany, and France as separate countries, euro area-12 aggregate data have been employed in the paper. This allows to focus on the most recent float period (1980-2006), which has almost entirely been neglected in the literature so far, despite being particularly suitable for investigation.[12] Monthly time series data for the five countries involved, i.e. the US, Japan, the Euro-12 area, the UK, and Canada, over the period 1980:1-2006:6, have then been employed. In addition to the four nominal exchange rate variables against the US\$, i.e. the €/US\$ rate, the yen/US\$ rate, the GBP£/US\$ rate, and the Canadian \$/US\$ rate, four macroeconomic variables for each country have been considered, i.e. the real industrial production growth rate, the CPI inflation rate, the nominal money growth rate and the nominal short-term interest rates.[13],[14] Monthly (log) volatility proxies for the above variables have then been constructed as the (log) absolute value of the innovations of the various series, obtained from the estimation of a standard VAR model for the 24 variables in the data set, with lag length set to two lags on the basis of misspecification tests and the AIC criterion. Although this yields noisy volatility proxies, the use of an effective noise filtering procedure grants reliability to the results obtained in the study.

The selection of the data set follows the monetarist models of exchange rate determination, from which the following reduced form exchange rate equation can derived

$$e = m^s - m^{s^*} - \phi\left(y - y^*\right) + \alpha\left(i - i^*\right) + \beta\left(\pi - \pi^*\right), \tag{4}$$

stating that the log level of the nominal exchange rate (e) is determined by differentials in the log money supplies (m^s), log real outputs (y), nominal interest rates (i) and inflation rates (π) between the domestic and foreign

[12] In the light of the literature on the great moderation, the period considered is homogeneous in terms macroeconomic volatility conditions. See for instance Stock and Watson (2003).

[13] Nominal money balances are given by M2 for the US, M2+CD for Japan, M3 for the euro area and Canada, and M4 for the UK. The aggregates employed are the one usually employed to measure broad money in each of the countries investigated. On the other hand, the short term rate refers to three-month government bills. The use of broad money is justified by country homogeneity, since, as far as Japan is concerned, in the view of the near liquidity trap experienced by this latter countries over the 1990s, the use of narrow money would have been problematic.

[14] Sinthetic euro area data are employed in this study. The author is grateful to the ECB, Monetary Policy Strategy Division, for data provision.

(starred variables) countries.[15]

Hence, a general long-run reduced form equation may be written as

$$e_t = \mathbf{z}_t'\boldsymbol{\delta} + \varepsilon_t, \tag{5}$$

where the vector $\mathbf{z}_t$ contains the macroeconomic fundamentals and ε_t is a zero mean stochastic disturbance term capturing non fundamental determinants, for instance related to speculative behavior in the exchange rate market. Hence, by moving to second moments, assuming orthogonal fundamentals and non fundamental determinants it follows

$$\sigma_{e_t}^2 = \sigma_{\mathbf{z}_t}^{2\prime}\boldsymbol{\delta}^2 + \sigma_{\varepsilon_t}^2, \tag{6}$$

pointing to a linkage between exchange rate ($\sigma_{e_t}^2$), fundamental ($\sigma_{\mathbf{z}_t}^2$) (macroeconomic) and non fundamental unconditional volatility ($\sigma_{\varepsilon_t}^2$).[16]

Consistent with general findings in the literature and the results of this study, both exchange rate and macroeconomic volatility is modelled as a stationary long memory process ($I(d)$, $0 < d < 0.5$), subject to structural change. Hence, for the generic ith exchange rate volatility process one has

$$\sigma_{i_t}^2 = b_t + P_t + NP_t, \tag{7}$$

where b_t is the deterministic break process (time-varying unconditional variance) of the series, i.e. the permanent or long-term component, expected to be related to fundamentals, i.e. $b_t = f(\sigma_{\mathbf{z}_t}^{2\prime}\boldsymbol{\delta}^2)$, P_t is the persistent (long memory, $I(d)$) or medium-term component, expected to be related to the non fundamental volatility component, i.e. $P_t = f(\sigma_{\varepsilon_t}^2)$, in the light of the explanations provided for long memory in volatility (see Andersen and Bollerslev, 1997; Muller et al., 1997), and NP_t is the non persistent or noise component ($I(0)$), with $E[P_t] = 0$ and $E[NP_t] = 0$.[17]

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[15] In particular, if $\alpha < 0$, $\beta > 0, |\beta| > |\alpha|$ the Frenkel real interest differential model is obtained; if $\alpha > 0$, $\beta = 0$ the flexible price monetarist model is obtained; if $\alpha = 0$, $\beta > 0$ the flexible price with hyperinflation monetarist model is obtained; if $\alpha < 0$, $\beta = 0$ the Dornbusch-sticky price monetarist model is obtained. Finally, assuming $\beta = 0$, and including in the equation the equlibrium real exchange rate, the equilibrium model is obtained. See Taylor (1995) for additional details.

[16] A similar relationship may be derived for the conditional variance. In fact by writing $e_t = E[e_t|I_{t-1}] + \boldsymbol{\varepsilon}_{zt}'\boldsymbol{\delta} + \varepsilon_t$, ad assuming the orthogonality of all the zero mean ε_i heteroskedastic innovations, it follows $\sigma_{e_t}^2 = \sigma_{\varepsilon_{\mathbf{z}t}}^{2\prime}\boldsymbol{\delta}^2 + \sigma_{\varepsilon_t}^2$, where the $\sigma_{i_t}^2$ are now conditional variances.

[17] Under covariance stationarity, the mean reversion property implies that the long-term forecast of a process converges to the unconditional mean. Hence, the breaak process b_t can be interpreted as the long-term forecast for the series $\sigma_{i_t}^2$, since $\lim_{s\to\infty} E_{t+s}\sigma_{i_t}^2 = b_{t+s}$,

Linkages among volatility series can then concern either the long-term or break process component b_t or the medium-term or long memory component P_t, or both. In the former case, two or more volatility series driven by the same break process component/s are then said to be cobreaking, and a long-term relationship can be found for the involved series which is free from structural change. Similarly, in the latter case, the series showing the same common long memory trend/s are said to be fractionally cointegrated, and a medium-term relationship can be found for the involved series which is free from long memory, or characterized by a lower order of integration than the involved series.[18]

The FI-F-VAR model employed in the paper allows to account for both kinds of linkages, controlling for short-term dynamic linkages as well. Moreover, conditioning is made relatively to a very large information set, i.e. the entire G-7 macroeconomic structure, allowing therefore for a fine control of the interrelations occurring across countries, currencies and macroeconomic factors.

3.1 Persistence properties

In the light of recent results in the literature pointing to the presence of both long memory and structural change in the volatility of financial assets, as well as in macroeconomic variables, the persistence properties of the data have been assessed by means of structural break tests and semiparametric estimators of the fractional differencing parameter. Structural change analysis has been carried out by means of the Dolado et al. (2004) test, modified to account for a general and unknown structural break process, with small sample critical values computed by means of the parametric bootstrap.[19]

given that $\lim_{s\to\infty} E_{t+s}NP_t = 0$ and, for $d < 0.5$, $\lim_{s\to\infty} E_{t+s}P_t = 0$. For this reason the permanent component is also denoted as the long-term component. On the other hand, the persistent component can be interpreted as the medium-term component, since for a sufficiently long, but finite forecast horizon k $\lim_{s\to k<\infty} E_{t+s}(\sigma^2_{i_t} - b_t) = \lim_{s\to k<\infty} E_{t+s}P_t$, given that $\lim_{s\to k<\infty} E_{t+s}NP_t = 0$. The non persistent component can finally be interpreted as the short-term component, possibly associated with measurement noise. See Morana (in press) for additional details on the permanent-persistent-non persistent decomposition for long memory processes.

[18] See Engle and Granger (1987), Robinson and Yajima (2002) and Marinucci and Robinson (2001) for seminal works on fractional cointegration. See Hendry (1996) for the seminal contribution on the theory of cobreaking and Hendry and Massmann (2007) for a recent survey.

[19] See Poskitt (2005) for a justification of the approach and the Appendix for a description of the modified Dolado et al. (2004) structural break test.

Since the computed log volatility series are likely to be characterized by measurement/observational noise, the null of pure long memory, i.e. long memory without structural change, may be expected to be favored in the above test. In fact observational noise may mask the non linearity determined by the break process, yet downward biasing the estimated persistence parameter as well. This could lead, for instance, to less significant t-test statistics in the Dolado et al. (2004) auxiliary regression equation. Then, in order to account for the presence of observational noise, the fractional differencing parameter employed in the test has been estimated by means of the Sun and Phillips (2003) non linear log periodogram estimator, which does not suffer from the downward bias affecting standard semiparametric estimators in this latter situation.

3.1.1 Structural break and long memory analysis

As shown in Table 1, Panel A, there is strong evidence of structural change in the volatility series investigated, since the null of pure long memory against the alternative of structural change is strongly rejected (at the 1% significance level) for all the series, apart from the euro area CPI inflation rate and the US short-term rate series.[20] The estimated volatility break processes are plotted in Figure 1, considering each sub group of variables at the time.[21] As shown in Figure 1, both similarities and differences can be noted in the plots. Firstly, in terms of descriptive statistics, homogeneity can be found within sets of variables but ethereogeneity across sets of variables, with exchange rate volatility showing both the highest average levels and variability (the sample means (with standard deviations) are 1.12%(0.19%), 1.37%(0.27%), 1.17%(0.35%) and 0.61%(0.19%), for the €/US$, the yen/US$, the GBP£/US$, and the Canadian $/US$, respectively). Among the macroeconomic variables, output growth rates are the most volatile variables, followed by money growth and inflation, and short-term rates (the average values and standard errors within categories are 0.46% (0.10%), 0.16% (0.04%), 0.10% (0.03%), and 0.02% (0.01%), respectively). Hence, in general macroeconomic volatility tends to be lower than exchange rates volatility when a series by series comparison is made. Secondly, evidence of clear-cut common dynamics can be found only for the interest rate volatility series, all of them showing the same downward trend over the time span considered. Yet, as shown by the princi-

[20] Following the Monte Carlo results reported in Cassola and Morana (2006), a fourth order trigonometric expansion has been employed for the computation of the structural break tests. Details are available upon request from the author.

[21] The volatility break processes have been obtained from the estimated log volatility break processes through Jensen transformation.

pal components analysis carried out within each set of variables, evidence of common dynamics can be found in all the cases. In fact, for all the subsets the bulk of volatility variance is explained by at most three principal components. For instance, for output growth volatility the first three principal components account for 92% of total variance, while figures for the other macroeconomic volatility variables are 81% for inflation volatility, 84% for money growth volatility, and 95% for the short-term interest rate volatility, although for this latter variable the first principal component alone explains already about 80% of total variance. On the other hand, as far as exchange rates volatility is concerned, the first three principal components (out of four) jointly account for 95% of total variability, with the first two components jointly accounting for about 80% of total variability. Finally, in all the cases the first principal component alone accounts for a large proportion of variance, i.e. about 40% for output growth, inflation and money growth volatility, while figures for interest rates and exchange rates volatility are 81% and 52%, respectively. Overall, the evidence of comovements in volatility is consistent with the evidence of comovement in first moments detected for the macroeconomic variables investigated, pointing to the relevance of both global and regional factors in explaining common macroeconomic dynamics in the G-7 area.[22]

Moreover, as shown in Panel B, Sun and Phillips (2003) non linear log periodogram estimation carried out on the break-free processes points to a moderate degree of long memory characterizing the break-free processes, ranging from 0.249(0.086) to 0.363(0.081), and to large inverse long-run signal to noise ratios, ranging from 16.960(5.162) to 30.113(7.236). Since in none of the cases the Robinson and Yajima (2001) test allows to reject the null of equality of the fractional differencing parameter, a single value for the fractional differencing parameters has then been obtained by averaging the twenty four available estimates, yielding a point estimate of 0.311(0.084).[23] Similarly, the average value of the inverse long-run signal to noise ratio is equal to 22.245(5.256). Given the size of the inverse long-run signal to noise ratios, the filtering of the volatility components is required before further analysis is carried out on the break-free series.[24] In the light of the above

[22] See for instance Bagliano and Morana (2006) and reference therein.

[23] The minimum p-value for the Robinson and Yajima (2002) equality test is equal to 0.423, which is very far apart from any usual significance value for the rejection of a simple or joint null hypothesis.

[24] In order to assess the robustness of the results, the break tests have been repeated considering trigonometric expansions of the first, second, and third order, allowing the fractional differencing parameter to take three different values, i.e. $d = 0.2$, 0.3, 0.4 in each case. The results point to the robustness of the structural break tests to both

results, the estimated candidate break processes have then been retained as non spurious for all the series.

3.1.2 Noise filtering

Noise filtering has been carried out using the approach of Morana (2007). The approach is based on flexible least squares estimation (Kalaba and Tesfatsion, 1989) and it has been found to perform very satisfactorily by Monte Carlo analysis, independently of the actual characteristics of the noisy stochastic process, i.e. deterministic versus stochastic persistence and long versus short memory, also when the inverse signal to noise ratio is very large. The key advantage of the approach, relatively to other available approaches suitable for long memory processes (Harvey, 1998; Beltratti and Morana, 2006; Morana, in press) is that it can be carried out directly on the actual series, also when structural instability characterizes the series, without requiring pretesting, and therefore the estimation of the order of integration of the process or the estimation of the actual break process. Hence, the noise filtering approach implemented is fully robust to the results of the persistence analysis carried out in the previous section.

Following the Monte Carlo results reported in Morana (2007), the Box-Pierce test optimal filtering based procedure has been employed for noise filtering, setting the significance level of the test at 1%. Concerning the robustness of the break process estimation procedure to observational noise, by comparing the break process estimated using the actual or noise filtered series, it can be concluded in favor of its robustness, since the largest root mean square error obtained from the comparison of the estimated break processes is just about equal to 0.02. The robustness of the findings points to the reliability of the filtering method, which is expected to remove only the non persistent noise dynamics from the processed series.[25]

4 The FI-F-VAR model

The minimum dimension of the FI-F-VAR model in the current data framework is twenty four equations, corresponding to the log volatility series for the twenty macroeconomic series and the four exchange rate series. Additional equations would refer the common long memory factors, which exis-

the order of the trigonometric expansion and the selection of the fractional differencing parameter. The results are available upon request from the author.

[25] Further support for the noise filtering methodology is provided by the robustness of the cobreaking, long-term causality and cointegration analyses carried out in the next sections.

tence has however to be determined through fractional cointegration/PCA analysis. Similarly, the existence of common break processes has to be determined through cobreaking/PCA analysis. Given the time series properties of the data investigated, showing both structural breaks and long memory, the presence of commonalities in the deterministic (cobreaking) and long memory (fractional cointegration) trends has then been investigated by means of principal components analysis (PCA), consistent with the first step required for the estimation of the FI-F-VAR model.

4.1 Cobreaking analysis

In order to investigate the linkages between macroeconomic and exchange rates volatility, the two-country specification implied by standard economic theories of bilateral exchange rates determination has been employed. Hence, linkages between macroeconomic and exchange rates volatility have been investigated with reference to a single bilateral exchange rate at the time and the macroeconomic variables of the involved countries. The results are reported in Table 2, Panels A and B.

As shown in Panel A, evidence of strong linkages among the volatility series investigated can be found for all the exchange rates, since just five factors are sufficient to account for about 95% of total variability for the nine variables involved in each case.[26] Additional similarities across currencies are worth noting.

Firstly, the first principal component accounts for a proportion of total variability in the range 30% to 43% across models, mostly affecting macroeconomic volatility, but exchange rate volatility only at a low extent (in the range 1% to 15%).[27] In particular, this latter component largely affects the short-term rate volatility in all of the cases. For instance, for the US figures are in the range 48% to 84%, while for the other countries figures are in the range 58% to 93%. The impact on output growth and money growth volatility is also large, i.e. in the range 11% to 77% and 1% to 20% for the US, respectively, and in the range 33% to 71% and 50% to 68%, respectively, for the other countries.[28] Finally, the impact on the inflation rate volatility

[26] The fraction of the total variance attributed to PC_j is given by $\lambda_j / (\sum_{i=1}^{n} \lambda_i)$, where λ_j is the j-th largest characteristic root of the sample variance-covariance matrix of the series.

[27] The proportion of variance of the i-th variable accounted by the j-th principal component can be computed as $\pi_{i,j} = d_{ij}^2 \lambda_j / (\sum_j d_{ij}^2 \lambda_j)$, where d_{ij} is the ijth entry in the $\mathbf{\Lambda}_\mu$ matrix.

[28] Only for the euro area the impact on output volatility is null, while only for the UK the impact on inflation volatility is negligible (2%).

is in general more modest, being noticeable only for π_{UK} and π_{JA} (84% and 38%, respectively).

Secondly, the second principal component accounts for a proportion of total variability in the range 21% to 27%, also accounting for 59% of the €/US$ exchange rate (e_{EA}) volatility, 68% for the yen/US$ exchange rate (e_{JA}) volatility, 44% of the GBP£/US$ exchange rate (e_{UK}) volatility, and 32% for the Canadian $/US$ (e_{CA}) exchange rate volatility. The second principal component also accounts for a non negligible proportion of macroeconomic volatility, although some interesting differences can be noted across exchange rates in terms of the macroeconomic variables involved. For instance, for e_{EA} the second principal component accounts for a large proportion of euro area output growth volatility (g_{EA}, 66%), while the proportion of explained US output growth volatility (g_{US}, 19%) is lower, albeit non negligible. Differently, for e_{UK} the second principal component accounts for a large proportion of US output growth volatility (g_{US}, 79%), but for a smaller proportion of UK (g_{UK}, 15%) output growth volatility. Similar output growth volatility figures can be found for the other exchange rates, i.e. 11% and 19% for g_{US} and g_{JA} for e_{JA}, and 10% and 13% for g_{US} and g_{CA} for e_{CA}. The proportion of inflation volatility variability explained by the second principal component is also large in all of the cases, apart from π_{UK}, ranging from 12% to 36% for π_{US} across exchange rates, while the other figures are 29% for π_{EA}, 14% for π_{JA}, and 47% for π_{CA}. On the other hand, short-term rate volatility and money growth volatility tend to be less affected by the component, albeit the proportion of explained variance for these latter variables is negligible only in few cases. In fact, for the US figures are in the range 3% to 24% for i_{US} and 0% to 47% for m_{US}. For the other countries figures are in the range 0% to 27% for the short-term rate volatility and 1% to 46% for money growth volatility.

Thirdly, the interpretation of the third, fourth and fifth components is less clear-cut, with the third and fourth components explaining about 17% of total variability each, and the fifth component only a residual 6%. Yet, these latter components seem to be important to account for additional, albeit minor, linkages among variables. The third and fourth components jointly account for 39% of e_{UK} variance and 44% of e_{CA} variance, but only for 20% and 8% of e_{EA} and e_{JA} variance, respectively. On the other hand, the fifth component only accounts for residual 19% and 8% of e_{EA} and e_{JA} variances, respectively. Moreover, as far as macroeconomic volatility is concerned, the third component accounts for a non negligible proportions of inflation volatility in all the cases, apart from the GBP£/US$ exchange rate (21% to 77% for π_{US}; 29% π_{EA}, 36% π_{JA}, 41% π_{CA}). The third component also accounts for a non negligible proportion of output volatility, i.e. g_{US} for the €/US$

(22%) and yen/US$ (29%). Differently, for the GBP£/US$ and Canadian $/US$ exchange rates the third component accounts for a non negligible proportion of money growth volatility, i.e. m_{US} (42% to 59%) and m_{UK} (33%). In addition, for the €/US$, the yen/US$ and the Canadian $/US$ is the fourth component to account for an important proportion of money growth volatility, particularly for the US (16% to 84% for m_{US}; 19% for m_{EA}, 14% for m_{JA}, and 26% for m_{CA}), also accounting for residual variability for g_{EA} (14%), g_{CA} (29%), and π_{US} (11% to 44%). Finally, the fifth component affects only e_{EA} and e_{JA} (19% and 9%, respectively), with minor effects on macroeconomic volatility (13% for π_{EA}, and about 20% for i_{US} and m_{JA}, respectively).

Overall the evidence points to strong and robust long-term linkages between macroeconomic and exchange rates volatility, with the bulk of long-run exchange rate volatility fluctuations associated with output growth and inflation volatility, and to money growth volatility at a lower extent.

4.1.1 Are cross-country interactions between macroeconomic and exchange rates volatility relevant?

The existence of cross-country interactions can be assessed by considering each exchange rate volatility series jointly with all the macroeconomic volatility series. As shown in Table 2, Panel B, findings are similar, since nine components are necessary to account for 100% of total variability, with seven factors accounting for over 95% of total variability. Moreover, for all the exchange rates the bulk of exchange rate volatility is explained by two principal components at most: figures are 61% for the €/US$, 82% for yen/US$, 84% for the GBP£/US$, 75% for the Canadian $/US$ exchange rates. These latter components also account for large proportions of macroeconomic volatility for all the countries, particularly output growth, inflation and money growth volatility (averages across countries are in the range 30% to 54% for output growth, 16% to 47% for inflation, 18% to 36% for money growth), while short-term rate volatility would seem to matter for the GBP£/US$ (80%; averages for the other countries are in the range 5% to 8%). Differently from the other exchange rates, three additional factors would be necessary to jointly account for 90% of total variance for the €/US$, while for the Canadian $/US$ exchange rate an additional factor would be necessary to jointly account for 91% of total variance. Adding an additional component in both cases leads to similar results in terms of explained variance, i.e. about 50% for output growth, money growth and inflation volatility, and about 80% for short-term rate volatility.

Similar findings hold also when the cross-country interactions are isolated, i.e. when the US and the own country variables are neglected, since the proportion of accounted variance in this latter case falls in the range 25% to 63% for output growth, 17% to 67% for inflation, 6% to 88% for the short-term rate, and 19% to 59% for money growth. Hence, the key role of output and inflation volatility for long-term inflation volatility is confirmed by the cross-country analysis as well. Overall, the relevance of cross-country interactions should not be surprising, given the evidence of global macroeconomic factors for the G-7 economy, both real and nominal (see for instance Bagliano and Morana, 2006 and references therein).

4.1.2 Long-run causality analysis

While the principal components analysis carried out on the long-term volatility components has provided clear-cut evidence concerning the existence of linkages between macroeconomic and exchange rate volatility, still no conclusions concerning the direction of causality of these linkages can be drawn on the basis of PCA alone.

Hence, two long-run causality tests have been carried out. The first test is based on the regression of each volatility proxy on its lagged values and on the lagged values of the estimated break process for all the other variables, excluding the break process for the variable under testing. In the light of the results of the principal components analysis, in order to avoid perfect multicollinearity, the analysis has been carried out considering the first four principal components of the set of break processes (nineteen series) for macroeconomic volatility and the first three principal components of the set of break processes (four series) for exchange rate volatility. The selection grants that about 90% of total variance is jointly explained by the factors.[29] Moreover, by extracting the common break processes separately from the macroeconomic and exchange rate volatility series, no contaminations are allowed between these two potentially different causing forces. Finally, since the information concerning the own break process is fully disregarded when testing is carried out, no bias is imparted to the results. The long-run causality test is then carried out by assessing the statistical significance of the lagged break processes in the estimated regression equation, distinguishing between the contribution of macroeconomic and exchange rate volatility. Moreover, for multicollinearity reasons, only a single lag for each principal component is

[29] This figure (88%) is obtained by averaging the proportion of total variance explained by the first four principal components extracted by the set of twenty macroeconomic volatility series alone (82%) and the proportion of total variance explained by the first three principal components extracted by the set of four exchange rate volatility series alone (94%).

included in the specification at the time. Hence, twelve different regressions ($m = 1, ..., 12$) have been run for each case.[30]

Hence, a typical regression equation for Test 1 could be written as

$$y_{s,t} = \alpha + \sum_{j=1}^{12} \beta_j y_{s,t-j} + \sum_{i=1}^{3} \gamma_i bp_{i,t-m}^{PC_e} + \sum_{i=1}^{4} \delta_i bp_{i,t-m}^{PC_m} + \varepsilon_t \quad \text{with } m = 1, ..., 12$$

where $y_{s,t}$ is the volatility proxy for series s, $bp_{i,t-m}^{PC_e}$ is the ith principal component extracted from the exchange rate volatility series, and $bp_{i,t-m}^{PC_m}$ is the ith principal component extracted from the macroeconomic volatility series. The null of no causality is then rejected if the estimated γ_i or δ_i parameters of interest are found to be statistically significant.

On the other hand, Test 2 requires the assessment of the closeness between the estimated break process obtained by means of the flexible Gallant (1984) functional form and the break process obtained performing Test 1. In fact, from the above regression a new estimate of the break process for the series $y_{s,t}$ can be obtained as

$$\hat{b}p_{s,t} = \hat{\beta}(L)^{-1}(\hat{\alpha} + \sum_{i=1}^{3} \hat{\gamma}_i bp_{i,t-m}^{PC_e} + \sum_{i=1}^{4} \hat{\delta}_i bp_{i,t-m}^{PC_m}),$$

setting $\delta_i = 0 \ \forall i$ and $\delta_i = 0 \ \forall i$, alternatively, and compared with the one estimated using the Gallant flexible functional form specification, when performing the structural break tests, i.e.

$$\hat{b}_{s,t} = \hat{b}_0 + \hat{b}_1 t + \sum_{k=1}^{4} \left(\hat{b}_{s,k} \sin(2\pi kt/T) + \hat{b}_{c,k} \cos(2\pi kt/T) \right),$$

from the OLS regression $y_{s,t} = b_t + \varepsilon_t$, by means of correlation analysis. Hence, the different contribution provided by the macroeconomic and exchange rate break processes in determining the break component for each series may be assessed and a comparison between the explanatory power of the alternative specifications carried out.

The results of the long-run causality analysis are reported in Table 3. As shown in the table, according to the results of the first test, there is clear-cut evidence of bidirectional long-run causality between macroeconomic and exchange rate volatility. In fact, in only seven cases out of twenty the null

[30] The maximum lag length (12 lags) has been selected according to specification tests for the residuals, also allowing for a sufficient time lag for cross effects to manifest.

of no causality of exchange rate volatility for macroeconomic volatility is not rejected at any lag order at the 10% significance level[31], while in all of the cases the null of no long-run causality of macroeconomic volatility for exchange rate volatility can be rejected. Moreover, when the own explanatory power is assessed, the analysis points to the lack of any explanatory power of exchange rate volatility for e_{EA}, e_{UK} and e_{CA}. Differently, only the lack of any explanatory power of macroeconomic volatility for UK, Japan, and euro area inflation volatility and euro area and UK output growth volatility can be detected.

Yet, in the light of the results of the second test, the causality linkage seems to be stronger from macroeconomic volatility to exchange rate volatility than the other way around. In fact, as far as macroeconomic volatility is concerned, macroeconomic volatility it self has a stronger explanatory power (i.e. higher correlation coefficient) than exchange rate volatility in thirteen out of twenty cases and a lower explanatory power in only seven cases. On the other hand, as far as exchange rate volatility is concerned, macroeconomic volatility always has a stronger explanatory power than exchange rate volatility it self. The stronger explanatory power of macroeconomic volatility for exchange rate volatility is a clear-cut result, since the correlation coefficient for the estimated break processes tend to be much lower when estimation is performed using the exchange rate volatility principal components than the macroeconomic volatility principal components. Figures are in fact -0.48 and 0.64, respectively, for e_{EA}, 0.12 and 0.74, respectively, for e_{JA}, 0.16 and 0.82, respectively, for e_{UK}, and 0.25 and 0.76, respectively, for e_{CA}.[32]

[31] The non rejection cases are for output volatility for the US, Japan and the UK, inflation volatility for the euro area and the UK, the short term rate volatility for the UK and money growth volatility for the euro area.

[32] For robustness the analysis has been carried out on the noise-free series as well. The findings are virtually unchanged for both tests. Concerning the first test, the results point out that in only five cases out of twenty the null of no causality of exchange rate volatility for macroeconomic volatility is not rejected at any lag order at the 10% significance level, while in all of the cases the null of no long-run causality of macroeconomic volatility for exchange rate volatility can be rejected. On the other hand, the same findings, in term of relative performance in reconstructing the break process, are obtained independently of the noise filtering. Hence, the stronger causality linkage from macroeconomic to exchange rate volatility than the other way around is confirmed by the robustness analysis. Detailed results have not been reported for reasons of space, but are available upon request from the author.

4.1.3 Can the level of macroeconomic volatility account for the level of exchange rate volatility in the long-run?

In order to assess whether macroeconomic volatility can account for the long-term level of exchange rate volatility, a regression analysis exercise, involving the volatility break processes, computed from the corresponding log volatility break processes obtained by means of the Gallant flexible functional form, using Jensen transformation, has been carried out. For each exchange rate, the exchange rate volatility break process has been regressed on the break processes for the macroeconomic volatility break processes for the involved countries, following a bilateral approach, rather than a joint approach. In Figure 2 the actual and fitted exchange rate break processes are plotted. As shown in the plots, macroeconomic volatility can account not only for the dynamics of long-run exchange rates volatility, but also for its level, independently of the inclusion of a constant term in the specification.[33] Albeit this latter result cannot be taken as direct evidence in favour of monetarist models of exchange rate determination, since the size of the estimated parameters, as well as overshooting effects, should also be taken into account, it however provides additional support to the evidence of a strong linkage between macroeconomic and exchange rate volatility, and, in particular, to the evidence of the causal power of macroeconomic volatility for long-term exchange rate volatility. Actually, with a long-term view, focusing on macroeconomic volatility may then be important to reduce excess exchange rates volatility, since, consistent with Friedman (1953), exchange rate instability may indeed be determined by macroeconomic instability.

4.2 Cointegration analysis

As for cobreaking analysis, the linkage between break-free noise-free macroeconomic and exchange rate volatilities has been investigated in the bilateral exchange rate framework. As shown in Table 4, principal components analysis points to weaker linkages involving macroeconomic and exchange rate volatility in the medium-term than in the long-term. In fact, in all of the cases seven principal components out of nine are necessary to explain about 90% of total variability, with the proportion of total variability explained by the first principal component falling in the range 18% to 23% only. Yet, some interesting similarities can be detected across countries. For instance, for all the countries apart from Canada, some linkages between exchange

[33]Similar findings hold when the volatilities of the spreads for the macroeconomic variables are employed as regressors, rather than using unconstrained regressors. These latter results are available upon request from the author.

rates and macroeconomic volatility can be found, with the bulk of exchange rates volatility (73% to 81%) explained by at most three components. Moreover, in general, US macroeconomic volatility seems to matter more than other countries (euro area, Japan, UK) macroeconomic volatility. Another similarity concerns the linkage involving output growth and exchange rate volatility for the €/US$ and the yen/US$ exchange rates. In fact, in both cases the dominant component for exchange rates volatility is also dominant for output growth volatility (49% for e_{EA} and 35% and 25% for g_{US} and g_{EA}, respectively; 39% for e_{JA} and 46% and 12% for g_{US} and g_{JA}, respectively). Moreover, while interest rates volatility is related to exchange rates volatility for both currencies, money growth volatility is related to exchange rates volatility only for the €/US$ exchange rate case, while for inflation volatility the linkage is only relevant for the yen/US$ exchange rate. Differently, for the Canadian $/US$ exchange rate the linkage involves all the macroeconomic volatility series. Finally, for the GBP£/US$ exchange rate the linkage is extremely weak, being non negligible only concerning inflation volatility.

Overall, it can be concluded that medium-term linkages between macroeconomic and exchange rate volatility are weaker than long-term linkages, yet still involving output volatility in particular. This latter finding is possibly not surprising, since the linkages between macroeconomic and exchange rate volatility implied by economic theory should be mostly relevant for the long-term. The finding is also consistent with the available literature, pointing to other determinants than fundamentals, as for instance excessive speculation, heterogeneous agents, overshooting effects related to information problems, and information flows, for medium-term exchange rate volatility. Moreover, considering cross country interactions does not lead to different results.[34]

4.3 Forecast error variance decomposition

Although on the basis of the findings it is possible to conclude against fractional cointegration for the pure long memory, or break-free, volatility components, still an assessment of the short- to medium-term dynamic linkages between macroeconomic and exchange rate volatility may be worthwhile. The final specification of the FI-F-VAR model is then composed of just the twenty

[34]This finding is further supported by the analysis carried out on different sub sets of break-free noise-free variables involving only the macroeconomic volatility variables or only the exchange rates volatility series. No evidence of truly global factors can be found in none of the cases. Yet, some evidence of regional factors can be found, a for instance for output growth, interest rates, and exchange rates volatility for the English speaking countries, and for exchange rates volatility for Japan and the euro area. Detailed results are available upon request from the author.

four equations corresponding to the log volatilities for twenty four variables in the data set, i.e. real output growth, inflation, the nominal short-term rate, and nominal money growth for the five countries in the system, and the returns for the €/US$, the yen/US$, the GBP£/US$, and the Canadian $/US$ exchange rates. The model has been estimated without long memory prefiltering, exploiting the results of Bauer and Maynard (2006) concerning the properties of the surplus lag approach. Moreover, seven common deterministic break processes, but no common long memory factors, have been included in the specification. Thick estimation (Granger and Jeon, 2004) has been implemented by optimally selecting the lag order by information criteria and then allowing for two additional specifications of higher and lower lag orders. According to the Akaike and Hannan-Quin information criteria, two lags have been selected, while according to the Bayes information criterion, just one lag could be selected. Hence, median estimates have been obtained by considering lag orders up to the third order. Monte Carlo replications have been set to 1000 for each case, considering two different orders for the variables, i.e. the volatility series have been ordered as output, inflation, short-term rate, money growth, and exchange rates, with countries selected as the US, euro area, Japan, the UK and Canada in the first case, and inverting the previous order in the second case. The median estimates have therefore been obtained from cross-sectional distributions counting 6000 units.[35]

4.3.1 Results

As shown in Table 5, the results of the forecast error variance decomposition are clear-cut, pointing to only weak short- to medium-term linkages between macroeconomic and exchange rate volatility and to mostly idiosyncratic long memory dynamics for all the variables. In fact, the own shock explains the bulk of volatility fluctuations for all the variables at all the horizons, with the contribution of the other shocks increasing as the forecast horizon increases. For instance, on average output volatility shocks explains between 77% and 85% of output volatility variance at the selected horizons. Similarly, inflation volatility shocks explains between 75% and 85% of inflation volatility variance. On the other hand, figures for the short-term rate and the money growth volatility are in the range 82% to 88% and 82% to 87%, respectively. Moreover, for the exchange rates figures are in the range 77% to 85%. In all the cases it is always the own volatility shock to explain the bulk of variability for the log volatility series. Finally, while some contribution to

[35] Detailed results have not been reported for reason of space. A full set of results is however available upon request from the author.

the explanation of macroeconomic volatility is provided by the exchange rate shocks, a much stronger role seems to be played by the macroeconomic shocks. In fact, on average the exchange rate volatility shocks explains only about 5% of macroeconomic volatility variance at all the forecasting horizons considered, while the average contribution of macroeconomic volatility shocks to exchange rate volatility variance is about 20%.[36] Hence, the stronger causality linkage from macroeconomic volatility to exchange rate volatility than the other way around, detected by the long-term causality analysis, is confirmed to hold in the short- to medium-term as well.[37]

4.4 Is there a trade-off between macroeconomic and exchange rates volatility?

In the light of the previous results, the linkage predicted by economic theory between macroeconomic and exchange rate volatility appears to be a robust finding, being stronger in the long-term than in the medium-term, albeit more than a transmission mechanism may be at work. When the factor loading matrices, obtained from the bilateral exchange rate long-term (Λ_μ) and medium-term (Λ_f) principal components analysis, are inspected, evidence of a trade-off between macroeconomic and exchange rates volatility can be noted for all the countries, for both the horizons. In Table 6, the percentage of variance involved in the trade-off for each volatility variables is reported.

As is shown in the table, apart from US money growth and output growth volatility, on average the trade-off is relatively stronger in the medium-term than in the long-term. Moreover, on average the variables which are more affected by the trade-off at both horizons are output growth and money growth volatility (the average proportion of variance for g_{US} is 53% and 63% in the long- and medium-term, respectively; for the other countries figures are 55% and 53%, respectively; for m_{US} figures are 64% and 51% in the long- and medium-term, respectively; for the other countries figures are 50% and 75%, respectively). On the other hand, short-term rate volatility is the variable for which the trade-off is weakest (the proportion of involved variance is in the range 29% to 39% for all the countries at both horizons), followed by

[36] The evidence of strongly idiosyncratic macroeconomic volatility shocks is consistent with previous results of Iwata and Wu (2005), pointing that exchange rates volatility does not account for more than 5% of output and inflation volatility, while output and inflation volatility can account for up to 67% of exchange rates volatility.

[37] Qualitatively similar results hold when the short- to medium-term causality analysis is carried out by means of standard Granger causality tests as in Chen (2006) or Bauer and Maynard (2006). Results are available upon request from the auhor.

inflation volatility (average figures are in the interval 31% to 37% and 51% to 67% in the long- and medium-term, respectively).

Interesting differences can also be found across exchange rates at the same horizon. For instance, the long-term trade-off, independently of the macroeconomic variable, is weakest for the UK, while it is strongest for Japan for output growth and inflation volatility, for Canada for interest rate volatility, and for the euro area for money growth volatility. Moreover, the US tends to be strongly affected by the trade-off, ranking always second. On the other hand, the euro area, as the UK, in general tends to be weakly affected by the trade-off, while Canada and Japan fall in an intermediate ranking. Differently, the medium-term trade-off is in general strongest for output growth for the UK and for money growth for the US. Moreover, for these two latter variables the trade-off is weakest for Japan. In addition for inflation and the short-term rate the trade-off is strongest for Canada and the US/Japan, respectively, and weakest for the UK. In general, the US is a country strongly affected by the trade-off, followed by the UK and the euro area, while Canada and Japan fall in an intermediate ranking.

Hence, in the light of the PCA analysis and the trade-off results, Friedman (1953) view that systemic volatility cannot be reduced by switching from floating to fixed exchange rates, but only traded-off, would seem to be supported by the empirical evidence, particularly in a long-term perspective.

5 Conclusions

What are the macroeconomic causes of exchange rate volatility? The paper provides a clear-cut answer to the above question. By means of a new fractionally integrated factor vector autoregressive (FI-F-VAR) model, evidence of significant long-term linkages and trade-offs between macroeconomic and exchange rate volatility have been found for the G-7 countries, involving output and inflation volatility in particular, and money growth volatility at a lower extent.

Moreover, although evidence of bidirectional causality has been found, linkages are much stronger from macroeconomic volatility to exchange rate volatility than the other way around. Interestingly, significant cross-country long-term interactions have been found as well, i.e. foreign countries macroeconomic volatility may be important for long-term exchange rate volatility as well. This latter finding is consistent with the evidence of global real and nominal dynamics found for the G-7 countries in the literature.

Hence, while other factors than macroeconomic fundamentals may be important determinant of exchange rates volatility in the short- to medium-term, neglecting the impact of macroeconomic volatility on long-term exchange rate volatility would however be inappropriate. Albeit indirectly, empirical support for fundamental models of exchange rate determination in the long-term has then been found.

Finally, two key policy implications follows from the results of the paper, consistent with the seminal views of Friedman (1953) on the case for flexible exchange rates: in a long-term perspective, focusing on macroeconomic stability may indeed be important to reduce excess exchange rates volatility; moreover, systemic volatility cannot be eliminated by fixing exchange rates, since the latter can only come at the cost of macroeconomic instability.

Overall, the paper not only provides encompassing evidence for the available results for the short- to medium-term, but new stylized facts for the long-term as well.

6 Appendix: the identification of structural shocks in the FI-F-VAR model

Given the stationarity of the long memory components the vector moving average form (VMA) for the $x_t - \Lambda_\mu \mu_t$ process

$$x_t - \Lambda_\mu \mu_t = G(L)\eta_t + F(L)v_t,$$

where $G(L) = F(L)[\Lambda_f + [E(L)\Phi(L)L]\left[I - \Phi(L)L\right]^{-1}]$ and $F(L) = [I - C(L)L]^{-1}$.

Following Bagliano and Morana (2006), the identification of the structural shocks in the FI-F-VAR model can be carried out as follows. Denoting by ξ_t the vector of the r structural global shocks, the relation between reduced form and structural form global shocks can be written as $\xi_t = H\eta_t$, where H is square and invertible. Then the identification of the structural shocks amounts to the estimation of the elements of the H matrix. It is assumed that $E\left[\xi_t \xi_t'\right] = I_r$, and hence $H\Sigma_\eta H' = I_r$. Moreover, by denoting ψ_t the n structural idiosyncratic shocks, the relation between reduced form and structural form idiosyncratic shocks can be written as $\psi_t = \Theta v_t$, where Θ is square and invertible. The identification of the structural idiosyncratic shocks then amounts to the estimation of the elements of the Θ matrix. It is assumed that $E\left[\psi_t' \psi_t\right] = I_n$, and hence $\Theta \Sigma_v' \Theta = I_n$.

The VMA representation of the factor model in structural form can then be written as

$$X_t = G^*(L)\eta_t + F^*(L)\psi_t,$$

where $G^*(L) = G(L)H^{-1}$, $F^*(L) = F(L)\Theta^{-1}$, and $E\left[\psi_{i,t}\xi_{j,t}'\right] = 0$ any i, j. Given r factors, then $r(r-1)/2$ restrictions need to be imposed in order to exactly identify the structural global shocks. Moreover, exact identification of the n structural idiosyncratic shocks requires the imposition of additional $n(n-1)/2$ zero restrictions.

The imposition of the exactly identifying restrictions is easily achieved by following a double Cholesky strategy, guided by economic theory, carried out as follows. Firstly, the structuralization of the factor or common shocks is achieved by assuming a lower triangular structure for the H matrix, with ordering of the variables set according to economic theory.[38] The H matrix is then written as

$$H = \begin{bmatrix} H_{11} & & \\ \vdots & \ddots & \\ H_{r1} & \cdots & H_{rr} \end{bmatrix},$$

and estimated by the Choleski decomposition of the matrix $\hat{\Sigma}_\eta$, i.e. from $\xi_t = H\eta_t$ we have $E[\eta_t\eta_t'] = H\Sigma_\eta H' = I$, and hence, $\hat{H} = chol(\hat{\Sigma}_\eta)$. The identification scheme performed allows for exact identification of the r structural global shocks, imposing $r(r-1)/2$ zero restrictions on the contemporaneous impact matrix.

Secondly, the matrix G_0^* is identified by imposing a lower triangular structure, with each non-zero block on the main diagonal showing a lower triangular structure as well, i.e.

$$G_0^* = \begin{bmatrix} G_{0_{11}}^* & & \\ \vdots & \ddots & \\ G_{0_{1r}}^* & \cdots & G_{0_{rr}}^* \end{bmatrix},$$

where

$$G_{0_{jj}}^* = \begin{bmatrix} g_{0_{ij,11}}^* & & \\ \vdots & \ddots & \\ g_{0_{ij,1m}}^* & \cdots & g_{0_{ij,mm}}^* \end{bmatrix},$$

and $n = mr$, with m equal to the number of units[39] in the sample. Again economic theory is called for to guide the ordering of the different units in each block.[40]

The estimation of the Θ matrix is then carried out as follows:

1) regress $\hat{\varepsilon}_{x,t}$ on $\hat{\xi}_t$ by OLS and obtain $\hat{v}_t$ as the residuals;

[38] For instance, standard economic assumptions concerning the speed of adjustment to shocks, i.e. slow (output, inflation), intermediate (interest rates, money growth), and fast (exchange rates) variables, could be employed.

[39] Countries, for instance, as in the current paper.

[40] For instance the distinction in large and small countries, in terms of GDP, could be employed.

2) then from $\psi_t = \Theta v_t$ we have $E[\psi_t \psi_t'] = \Theta \Sigma_v' \Theta = I$. Hence, $\hat{\Theta} = chol(\hat{\Sigma}_v)$.

The identification scheme performed allows for exact identification of the n structural idiosyncratic shocks, imposing $n(n-1)/2$ zero restrictions on the contemporaneous impact matrix.

Alternatively, in order to ensure robustness to variable ordering, estimation may be carried out by following a thick modelling estimation approach[41], consisting of estimating the model by using both the ordering suggested by economic theory and its inverse, different lag truncation orders, as well as simulating the model by Monte Carlo methods. From the cross sectional distribution of the relevant statistics, i.e. impulse response functions and forecast error variance decomposition, median estimates and 95% confidence levels for the parameters of interests can then be obtained. Finally, policy analysis can also be carried out by means of generalized impulse response analysis (Pesaran and Shin, 1998), which, by construction, is not affected by variables ordering. Therefore, estimation strategies allowing to draw robust conclusions not only to the ordering of the variables, but also to potential misspecification of the econometric model, are available for the proposed approach.

7 Appendix: A generalization of the Dolado, Gonzalo and Mayoral (2004) test

Dolado et al. (2004) have proposed a generalization of the Dickey-Fuller test to the I(d) case. The generalization allows for testing the null of I(d), $0 < d \leq 1$, against the alternative of trend stationarity I(0), with or without structural breaks.

As far as the case of no structural breaks is concerned, the process assumed under the alternative by Dolado et al. (2004) is

$$y_t = \mu + \frac{\varepsilon_t 1_{(t>0)}}{\Delta^d - \phi L}, \tag{8}$$

$$y_t = \mu + \beta t + \frac{\varepsilon_t 1_{(t>0)}}{\Delta^d - \phi L}, \tag{9}$$

with $\varepsilon_t \tilde{} iid(0, \sigma_\varepsilon^2)$, with $0 < \sigma^2 < \infty$, and $d_0 \in (0, 1]$.

Hence, under H_1

$$\Delta^d y_t = \alpha + \Delta^d \delta + \phi y_{t-1} + \varepsilon_t \tag{10}$$

[41] See Granger and Jeon (2004).

$$\Delta^d y_t = \alpha + \Delta^d \delta + \gamma t + \Delta^{d-1}\varphi + \phi y_{t-1} + \varepsilon_t, \tag{11}$$

where $\alpha = -\phi\mu$, $\delta = \mu$, $\gamma = -\phi\beta$ and $\phi = \beta$, $\varepsilon_t = \varepsilon_t 1_{(t>0)}$. The null hypothesis of I(d) implies $\phi = 0$ in both cases, while the alternative of I(0) stationarity implies $\phi < 0$. In the case the error term ε_t is a stationary I(0) process, the augmented form of the auxiliary test regression can be written as

$$\Delta^d y_t = \alpha + \Delta^d \delta + \phi y_{t-1} + \sum_{j=1}^{s} \Delta^d y_{t-j} + v_t, \tag{12}$$

$$\Delta^d y_t = \alpha + \Delta^d \delta + \gamma t + \Delta^{d-1}\varphi + \phi y_{t-1} + \sum_{j=1}^{s} \Delta^d y_{t-j} + v_t, \tag{13}$$

with $v_t \tilde{} iid(0, \sigma_v^2)$.

On the other hand, for the structural break case the process under the alternative is

$$y_t = A_B(t) + \frac{\varepsilon_t 1_{(t>0)}}{\Delta^d - \phi L}, \tag{14}$$

where $A_B(t)$ is a linear deterministic trend function showing breaks of various forms at unknown points in time. In line with the literature, three cases are considered, i.e. the crash hypothesis

$$A_B^1(t) = \mu_0 + (\mu_1 - \mu_0) DU_t(\lambda),$$

with $DU_t(\lambda) = 1_{(T_{B+1} \leq t \leq T)}$, the changing growth hypothesis

$$A_B^2(t) = \mu_0 + \beta_0 t + (\beta_1 - \beta_0) DT_t^*(\lambda),$$

with $DT_t^*(\lambda) = (t - T_B) 1_{(T_{B+1} \leq t \leq T)}$, and the crash plus changing growth hypothesis

$$A_B^3(t) = \mu_0 + \beta_0 t + (\mu_1 - \mu_0) DU_t(\lambda) + (\beta_1 - \beta_0) DT_t(\lambda),$$

with $DT_t(\lambda) = t 1_{(T_{B+1} \leq t \leq T)}$ and $\lambda = T_B / T$.

The process under H_1 is then

$$\Delta^d y_t = \Delta^d A_B^i(t) - \phi A_B^i(t-1) y_{t-1} + \phi y_{t-1} + \varepsilon_t \qquad i = 1, 2, 3 \tag{15}$$

The null hypothesis of I(d) again implies $\phi = 0$, while the alternative of I(0) stationarity plus structural change implies $\phi < 0$. In the case the error term ε_t is a stationary I(0) process, the augmented form of the auxiliary test regression can then be written as

$$\Delta^d y_t = \Delta^d A_B^i(t) - \phi A_B^i(t-1) y_{t-1} + \phi y_{t-1} + \sum_{j=1}^{s} \Delta^d y_{t-j} + v_t \qquad i = 1, 2, 3 \quad (16)$$

with $v_t \tilde{} iid(0, \sigma_v^2)$. The critical values for all the cases above are tabulated by Dolado et al. (2004).

A simple generalization of the model, allowing for a general non linear deterministic break process, can be obtained by specifying the trend function according to the Gallant (1984) flexible functional, as in Enders and Lee (2004), i.e.

$$A_B^{NL}(t) = \mu_0 + \beta_0 t + \sum_{k=1}^{p} \left(\beta_{s,k} \sin(2\pi kt/T) + \beta_{c,k} \cos(2\pi kt/T) \right),$$

allowing for a suitable order (p) for the trigonometric expansion. Critical values can be easily computed, case by case, by the parametric bootstrap. Monte Carlo evidence supporting the use of the above adaptive approach for structural break estimation can be found in Cassola and Morana (2006). Details concerning the selection of the order of the trigonometric expansion can also be found in Cassola and Morana (2006).

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Table 1: Persistence analysis

Panel A: Dolado et al. (2004) structural break tests

g_{US}	-4.381*	π_{US}	-4.447*	s_{US}	-1.132	m_{US}	-4.409*		
g_{EA}	-4.358*	π_{EA}	-1.849	s_{EA}	-3.447*	m_{EA}	-6.128*	e_{EA}	-5.688*
g_{JA}	-4.598*	π_{JA}	-4.079*	s_{JA}	-4.352*	m_{JA}	-4.792*	e_{JA}	-4.996*
g_{UK}	-4.980*	π_{UK}	-5.211*	s_{UK}	-4.818*	m_{UK}	-4.087*	e_{UK}	-3.655**
g_{CA}	-5.152*	π_{CA}	-4.513*	s_{CA}	-4.248*	m_{CA}	-3.849*	e_{CA}	-4.761*

Panel B: Break processes, descriptive statistics

	m (s)	PC_i		m (s)	PC_i		m (s)	PC_i		m (s)	PC_i		m (s)	PC_i
g_{US}	0.31 (0.10)	**0.40**	π_{US}	0.10 (0.03)	**0.40**	s_{US}	0.02 (0.01)	**0.81**	m_{US}	0.12 (0.04)	**0.40**			
g_{EA}	0.45 (0.08)	**0.33**	π_{EA}	0.06 (0.01)	**0.25**	s_{EA}	0.02 (0.00)	**0.08**	m_{EA}	0.12 (0.02)	**0.29**	e_{EA}	1.12 (0.19)	**0.52**
g_{JA}	0.57 (0.11)	**0.19**	π_{JA}	0.12 (0.03)	**0.17**	s_{JA}	0.01 (0.00)	**0.06**	m_{JA}	0.11 (0.02)	**0.14**	e_{JA}	1.37 (0.27)	**0.27**
g_{UK}	0.45 (0.09)	**0.06**	π_{UK}	0.10 (0.03)	**0.16**	s_{UK}	0.02 (0.01)	**0.03**	m_{UK}	0.24 (0.06)	**0.12**	e_{UK}	1.17 (0.35)	**0.15**
g_{CA}	0.54 (0.10)	**0.02**	π_{CA}	0.13 (0.03)	**0.02**	s_{CA}	0.02 (0.01)	**0.02**	m_{CA}	0.23 (0.06)	**0.04**	e_{CA}	0.61 (0.19)	**0.06**

Panel C: Fractional differencing parameter estimation

g_{US}	0.346 (0.106) [15]	π_{US}	0.346 (0.061) [26]	s_{US}	0.320 (0.111) [15]	m_{US}	0.310 (0.073) [23]		
g_{EA}	0.363 (0.081) [19]	π_{EA}	0.303 (0.090) [19]	s_{EA}	0.303 (0.067) [26]	m_{EA}	0.309 (0.091) [19]	e_{EA}	0.325 (0.080) [21]
g_{JA}	0.338 (0.078) [23]	π_{JA}	0.328 (0.074) [22]	s_{JA}	0.274 (0.076) [24]	m_{JA}	0.249 (0.086) [23]	e_{JA}	0.277 (0.106) [17]
g_{UK}	0.356 (0.047) [33]	π_{UK}	0.310 (0.106) [16]	s_{UK}	0.311 (0.088) [19]	m_{UK}	0.284 (0.056) [32]	e_{UK}	0.300 (0.076) [23]
g_{CA}	0.304 (0.095) [18]	π_{CA}	0.299 (0.099) [18]	s_{CA}	0.353 (0.080) [20]	m_{CA}	0.249 (0.118) [17]	e_{CA}	0.304 (0.070) [25]

Panel A reports the value of the Dolado et al. (2004) test. "$*$" denotes significance at the 1% level. In Panels B descriptive statistics, i.e. sample mean (m) and standard deviation (s, in brackets), for the estimated break processes are reported. The percentage of total variance explained by each principal component computed using the sub sets of data is also reported (PC_i). In Panel C the estimated fractional differencing parameters obtained using the Sun and Phillips (2003) non linear log periodogram estimator, with standard errors ($(\cdot)$) and selected ordinates ($[\cdot]$), are reported. The log volatility series investigated are real output growth rates (g), inflation rates (π), short-term nominal interest rates (s), nominal money growth rates (m), and nominal exchange rates returns for the euro, the Japanese yen, the British pound and the Canadian dollar against the US dollar (e). The countries investigated are the US (US), the euro area (EA), Japan (JA), the UK (UK), and Canada (CA).

Table 2: Principal components analysis (long-term)

Panel A: Bilateral exchange rates analysis

$€/US\$$	PC_1	PC_2	PC_3	PC_4	PC_5	PC_6	PC_7	PC_8	PC_9
all	0.30	0.23	0.17	0.16	0.06	0.04	0.03	0.01	0.00
g_{US}	0.43	0.19	0.22	0.03	0.05	0.01	0.07	0.00	0.00
g_{EA}	0.00	0.66	0.00	0.14	0.02	0.17	0.00	0.01	0.00
π_{US}	0.01	0.12	0.77	0.03	0.05	0.00	0.00	0.00	0.00
π_{EA}	0.10	0.29	0.29	0.08	0.13	0.12	0.00	0.00	0.00
s_{US}	0.84	0.05	0.01	0.00	0.01	0.00	0.01	0.06	0.00
s_{EA}	0.59	0.16	0.10	0.05	0.00	0.06	0.02	0.01	0.00
m_{US}	0.01	0.01	0.03	0.84	0.03	0.01	0.05	0.01	0.00
m_{EA}	0.68	0.01	0.02	0.19	0.00	0.00	0.07	0.02	0.00
e_{EA}	0.01	0.59	0.12	0.08	0.19	0.00	0.00	0.00	0.00
$YEN/US\$$	PC_1	PC_2	PC_3	PC_4	PC_5	PC_6	PC_7	PC_8	PC_9
all	0.43	0.21	0.15	0.09	0.06	0.03	0.02	0.01	0.00
g_{US}	0.52	0.11	0.29	0.00	0.01	0.02	0.06	0.00	0.00
g_{JA}	0.71	0.19	0.01	0.01	0.00	0.06	0.01	0.00	0.00
π_{US}	0.00	0.25	0.58	0.11	0.02	0.02	0.02	0.00	0.00
π_{JA}	0.38	0.14	0.36	0.05	0.02	0.02	0.00	0.02	0.00
s_{US}	0.48	0.24	0.04	0.00	0.20	0.01	0.03	0.00	0.00
s_{JA}	0.88	0.01	0.01	0.01	0.00	0.06	0.01	0.01	0.00
m_{US}	0.20	0.25	0.01	0.52	0.01	0.01	0.00	0.01	0.00
m_{JA}	0.60	0.01	0.01	0.14	0.22	0.00	0.01	0.00	0.00
e_{JA}	0.08	0.68	0.08	0.00	0.09	0.06	0.00	0.01	0.00
$GBP£/US\$$	PC_1	PC_2	PC_3	PC_4	PC_5	PC_6	PC_7	PC_8	PC_9
all	0.38	0.27	0.14	0.10	0.05	0.03	0.02	0.00	0.00
g_{US}	0.11	0.79	0.04	0.01	0.04	0.00	0.00	0.00	0.00
g_{UK}	0.65	0.15	0.01	0.00	0.01	0.18	0.00	0.00	0.00
π_{US}	0.01	0.46	0.02	0.44	0.05	0.02	0.00	0.00	0.00
π_{UK}	0.84	0.00	0.04	0.01	0.04	0.05	0.01	0.00	0.00
s_{US}	0.60	0.12	0.09	0.00	0.12	0.00	0.06	0.00	0.00
s_{UK}	0.93	0.00	0.01	0.02	0.00	0.00	0.03	0.01	0.00
m_{US}	0.16	0.00	0.59	0.16	0.06	0.00	0.02	0.00	0.00
m_{UK}	0.02	0.46	0.33	0.00	0.16	0.01	0.02	0.00	0.00
e_{UK}	0.13	0.44	0.10	0.29	0.00	0.04	0.00	0.00	0.00
$CA\$/US\$$	PC_1	PC_2	PC_3	PC_4	PC_5	PC_6	PC_7	PC_8	PC_9
all	0.35	0.26	0.17	0.11	0.06	0.03	0.01	0.01	0.00
g_{US}	0.77	0.10	0.08	0.01	0.00	0.02	0.01	0.01	0.00
g_{CA}	0.33	0.13	0.05	0.29	0.15	0.03	0.01	0.00	0.00
π_{US}	0.09	0.36	0.21	0.07	0.25	0.01	0.00	0.01	0.00
π_{CA}	0.01	0.47	0.41	0.03	0.00	0.04	0.04	0.00	0.00
s_{US}	0.70	0.03	0.13	0.01	0.08	0.02	0.00	0.03	0.00
s_{CA}	0.58	0.27	0.09	0.01	0.01	0.01	0.01	0.03	0.00
m_{US}	0.03	0.47	0.42	0.00	0.00	0.04	0.04	0.00	0.00
m_{CA}	0.50	0.16	0.00	0.26	0.07	0.01	0.00	0.00	0.00
e_{CA}	0.15	0.32	0.16	0.28	0.00	0.08	0.01	0.00	0.00

Panel A reports the results of the long-term principal components (PC) analysis carried out for each exchange rate on the log volatility for the relevant macroeconomic variables, i.e. real output growth rates (g), inflation rates (π), short-term nominal interest rates (s), nominal money growth rates (m), and nominal exchange rates returns for the euro, the Japanese yen, the British pound and the Canadian dollar against the US dollar (e). The countries investigated are the US (US), the euro area (EA), Japan (JA), the UK (UK), and Canada (CA). For each set the first row shows the fraction of the total variance explained by each PC_i $(i = 1, \ldots 9)$; the subsequent nine rows display the fraction of the variance of the individual series attributable to each PC_i.

Table 2: Principal components analysis (long-term)

Panel B: Joint analysis

€/US$	PC_1	PC_2	PC_3	PC_4	PC_5	PC_6	PC_7	PC_8	PC_9
all	0.37	0.17	0.15	0.12	0.06	0.05	0.05	0.03	0.02
g_{US+EA}	0.15	0.46	0.19	0.09	0.00	0.04	0.02	0.05	0.00
g_O	0.32	0.16	0.19	0.16	0.08	0.03	0.01	0.05	0.01
π_{US+EA}	0.02	0.23	0.04	0.41	0.05	0.11	0.07	0.02	0.05
π_O	0.54	0.05	0.19	0.02	0.05	0.04	0.04	0.04	0.02
s_{US+EA}	0.68	0.02	0.08	0.08	0.05	0.02	0.01	0.02	0.04
s_O	0.84	0.03	0.01	0.02	0.01	0.04	0.01	0.02	0.01
m_{US+EA}	0.16	0.05	0.48	0.11	0.01	0.07	0.10	0.02	0.01
m_O	0.17	0.32	0.11	0.12	0.10	0.05	0.11	0.01	0.01
e_{EA}	0.09	0.40	0.01	0.09	0.21	0.09	0.00	0.11	0.00
$YEN/US\$$	PC_1	PC_2	PC_3	PC_4	PC_5	PC_6	PC_7	PC_8	PC_9
all	0.37	0.17	0.16	0.12	0.06	0.05	0.04	0.02	0.02
g_{US+JA}	0.40	0.20	0.31	0.02	0.04	0.01	0.00	0.00	0.01
g_O	0.16	0.29	0.24	0.07	0.06	0.00	0.13	0.05	0.00
π_{US+JA}	0.28	0.10	0.22	0.14	0.07	0.02	0.10	0.05	0.02
π_O	0.36	0.23	0.10	0.12	0.09	0.03	0.02	0.01	0.04
s_{US+JA}	0.78	0.07	0.02	0.01	0.03	0.03	0.02	0.01	0.03
s_O	0.75	0.04	0.03	0.07	0.02	0.03	0.01	0.02	0.01
m_{US+JA}	0.25	0.16	0.06	0.19	0.00	0.29	0.03	0.00	0.01
m_O	0.13	0.08	0.32	0.30	0.11	0.02	0.04	0.01	0.00
e_{JA}	0.06	0.69	0.02	0.08	0.13	0.01	0.00	0.01	0.00
$GBP£/US\$$	PC_1	PC_2	PC_3	PC_4	PC_5	PC_6	PC_7	PC_8	PC_9
all	0.36	0.19	0.14	0.11	0.05	0.05	0.04	0.02	0.02
g_{US+UK}	0.39	0.43	0.04	0.04	0.00	0.03	0.03	0.04	0.00
g_O	0.17	0.21	0.29	0.17	0.04	0.05	0.04	0.02	0.02
π_{US+UK}	0.43	0.11	0.03	0.16	0.03	0.15	0.03	0.04	0.00
π_O	0.26	0.11	0.21	0.18	0.05	0.07	0.04	0.03	0.04
s_{US+UK}	0.72	0.05	0.08	0.04	0.05	0.00	0.01	0.01	0.04
s_O	0.81	0.02	0.01	0.05	0.04	0.01	0.01	0.03	0.02
m_{US+UK}	0.06	0.14	0.30	0.18	0.16	0.04	0.12	0.00	0.01
m_O	0.25	0.25	0.22	0.09	0.04	0.05	0.10	0.01	0.01
e_{UK}	0.00	0.75	0.00	0.00	0.09	0.06	0.01	0.08	0.00
$CA\$/US\$$	PC_1	PC_2	PC_3	PC_4	PC_5	PC_6	PC_7	PC_8	PC_9
all	0.39	0.17	0.14	0.11	0.06	0.05	0.04	0.02	0.02
g_{US+CA}	0.16	0.30	0.13	0.27	0.02	0.02	0.08	0.01	0.00
g_O	0.30	0.31	0.22	0.02	0.03	0.01	0.07	0.04	0.01
π_{US+CA}	0.09	0.12	0.32	0.17	0.18	0.01	0.09	0.02	0.02
π_O	0.51	0.15	0.02	0.16	0.02	0.05	0.03	0.03	0.05
s_{US+CA}	0.74	0.01	0.08	0.01	0.00	0.10	0.02	0.02	0.03
s_O	0.81	0.03	0.00	0.08	0.00	0.03	0.01	0.03	0.01
m_{US+CA}	0.13	0.33	0.30	0.07	0.05	0.06	0.05	0.01	0.01
m_O	0.18	0.11	0.22	0.18	0.13	0.14	0.03	0.00	0.00
e_{CA}	0.52	0.23	0.00	0.06	0.16	0.02	0.00	0.00	0.02

Panel B reports the results of the joint long-term principal components (PC) analysis carried out, for each exchange rate, using the log-volatility for all the macroeconomic variables and the log volatility of each nominal exchange rate at the time. For each set the first row shows the fraction of the total variance explained by each PC_i $(i = 1, ...9)$; the subsequent nine rows display the average fraction of the variance of the series for the involved countries (A+B) and for the other countries (O) attributable to each PC_i. For instance, in the €$/US\$$ sub table g_{US+EA} is the average proportion of variance for output growth volatility for the US and the euro area, while g_O is the average proportion of variance for output growth volatility for Japan, the UK and Canada.

Table 3

Long-term causality analysis

	$ev \rightarrow$		$mv \rightarrow$	
	ρ	*p-value*	ρ	*p-value*
g_{US}	0.21	0.383	0.88	0.041
g_{EA}	0.75	0.063	0.17	0.130
g_{JA}	0.21	0.131	0.84	0.014
g_{UK}	0.71	0.320	0.85	0.155
g_{CA}	0.42	0.029	0.62	0.021
π_{US}	0.38	0.030	0.58	0.002
π_{EA}	0.09	0.986	0.66	0.192
π_{JA}	0.77	0.040	0.69	0.141
π_{UK}	0.82	0.282	0.74	0.285
π_{CA}	0.30	0.004	0.57	0.005
s_{US}	-0.28	0.001	0.72	0.000
s_{EA}	0.76	0.034	0.70	0.045
s_{JA}	0.83	0.011	0.79	0.016
s_{UK}	-0.39	0.113	0.94	0.000
s_{CA}	-0.43	0.000	0.81	0.000
m_{US}	-0.28	0.069	0.77	0.002
m_{EA}	0.06	0.206	0.56	0.082
m_{JA}	0.78	0.002	0.46	0.030
m_{UK}	0.15	0.000	0.44	0.000
m_{CA}	0.65	0.000	0.24	0.002
e_{EA}	-0.48	0.246	0.68	0.014
e_{JA}	0.12	0.041	0.74	0.001
e_{UK}	0.16	0.385	0.82	0.063
e_{CA}	0.25	0.104	0.76	0.004

The table reports the minimum p-value for the long-term causality test (Test 1) and the correlation coefficient (ρ) for the actual break process and the break process reconstructed on the basis of the macroeconomic or exchange rates volatility principal components (Test 2). The first two columns refer to the case in which the direction of causality is from exchange rate volatility to macroeconomic volatility, and the break process is estimated using the first three principal components extracted from the exchange rate volatility series. The last two columns refer to the case in which the direction of causality is from macroeconomic volatility to exchange rate volatility, and the break process is estimated using the first four principal components extracted from the macroeconomic volatility series. The variable investigated are log volatilities for real output growth rates (g), inflation rates (π), short-term nominal interest rates (s), nominal money growth rates (m), and bilateral nominal exchange rates returns for the euro, the Japanese yen, the British pound and the Canadian dollar against the US dollar (e). The countries investigated are the US (US), the euro area (EA), Japan (JA), the UK (UK), and Canada (CA).

Table 4: Principal components analysis (medium-term)

Panel A: Bilateral exchange rates models

€/US$	PC_1	PC_2	PC_3	PC_4	PC_5	PC_6	PC_7	PC_8	PC_9
all	0.19	0.16	0.15	0.14	0.10	0.08	0.07	0.06	0.05
g_{US}	0.06	0.35	0.12	0.24	0.02	0.01	0.04	0.05	0.11
g_{EA}	0.19	0.25	0.14	0.07	0.03	0.03	0.21	0.08	0.00
π_{US}	0.13	0.04	0.40	0.01	0.20	0.08	0.10	0.01	0.04
π_{EA}	0.36	0.04	0.31	0.01	0.08	0.00	0.02	0.05	0.13
s_{US}	0.29	0.21	0.02	0.07	0.04	0.26	0.04	0.00	0.08
s_{EA}	0.25	0.00	0.18	0.03	0.39	0.00	0.08	0.08	0.00
m_{US}	0.12	0.03	0.13	0.36	0.02	0.24	0.00	0.09	0.00
m_{EA}	0.26	0.06	0.00	0.36	0.09	0.00	0.12	0.09	0.03
e_{EA}	0.03	0.49	0.02	0.10	0.07	0.14	0.02	0.10	0.04
$YEN/US\$$	PC_1	PC_2	PC_3	PC_4	PC_5	PC_6	PC_7	PC_8	PC_9
all	0.18	0.16	0.14	0.13	0.12	0.08	0.07	0.07	0.05
g_{US}	0.04	0.04	0.46	0.05	0.20	0.15	0.02	0.00	0.06
g_{JA}	0.00	0.52	0.12	0.08	0.01	0.01	0.18	0.00	0.08
π_{US}	0.12	0.28	0.10	0.20	0.02	0.00	0.03	0.24	0.01
π_{JA}	0.13	0.34	0.01	0.13	0.09	0.09	0.13	0.04	0.04
s_{US}	0.25	0.00	0.02	0.08	0.42	0.09	0.07	0.00	0.05
s_{JA}	0.35	0.10	0.14	0.00	0.03	0.25	0.10	0.01	0.01
m_{US}	0.63	0.03	0.03	0.00	0.03	0.02	0.01	0.11	0.14
m_{JA}	0.09	0.10	0.01	0.55	0.00	0.06	0.01	0.17	0.00
e_{JA}	0.01	0.06	0.39	0.10	0.26	0.00	0.08	0.01	0.08
$GBP£/US\$$	PC_1	PC_2	PC_3	PC_4	PC_5	PC_6	PC_7	PC_8	PC_9
all	0.20	0.16	0.14	0.13	0.11	0.09	0.07	0.06	0.05
g_{US}	0.07	0.20	0.07	0.16	0.00	0.45	0.01	0.03	0.00
g_{UK}	0.00	0.43	0.08	0.21	0.00	0.02	0.23	0.02	0.00
π_{US}	0.16	0.00	0.01	0.55	0.10	0.00	0.05	0.13	0.01
π_{UK}	0.36	0.11	0.22	0.02	0.01	0.04	0.05	0.12	0.07
s_{US}	0.60	0.09	0.03	0.02	0.01	0.00	0.00	0.07	0.17
s_{UK}	0.47	0.01	0.03	0.13	0.01	0.13	0.07	0.14	0.02
m_{US}	0.04	0.56	0.07	0.02	0.03	0.10	0.11	0.00	0.08
m_{UK}	0.03	0.05	0.06	0.08	0.71	0.04	0.01	0.00	0.00
e_{UK}	0.02	0.00	0.72	0.01	0.08	0.02	0.06	0.01	0.08
$CA\$/US\$$	PC_1	PC_2	PC_3	PC_4	PC_5	PC_6	PC_7	PC_8	PC_9
all	0.23	0.16	0.14	0.13	0.10	0.09	0.07	0.04	0.04
g_{US}	0.06	0.10	0.23	0.12	0.44	0.01	0.01	0.04	0.00
g_{CA}	0.13	0.12	0.31	0.00	0.12	0.25	0.00	0.01	0.05
π_{US}	0.00	0.16	0.28	0.31	0.02	0.12	0.03	0.06	0.00
π_{CA}	0.35	0.22	0.01	0.20	0.00	0.05	0.03	0.05	0.09
s_{US}	0.44	0.26	0.00	0.02	0.03	0.05	0.02	0.12	0.05
s_{CA}	0.42	0.04	0.18	0.00	0.08	0.05	0.12	0.07	0.04
m_{US}	0.04	0.36	0.00	0.12	0.09	0.21	0.17	0.00	0.00
m_{CA}	0.17	0.19	0.09	0.35	0.10	0.00	0.00	0.00	0.10
e_{CA}	0.42	0.01	0.12	0.01	0.02	0.10	0.27	0.04	0.01

The table reports the results of the medium-term principal components (PC) analysis carried out for each exchange rate on the break-free noise-free log-volatility for the relevant macroeconomic variables, i.e. real output growth rates (g), inflation rates (π), short-term nominal interest rates (s), nominal money growth rates (m), and nominal exchange rates returns for the euro, the Japanese yen, the British pound and the Canadian dollar against the US dollar (e). The countries investigated are the US (US), the euro area (EA), Japan (JA), the UK (UK), and Canada (CA). For each set the first row shows the fraction of the total variance explained by each PC_i ($i = 1, ...9$); the subsequent nine rows display the fraction of the variance of the individual series attributable to each PC_i.

Table 5
Forecast error variance decomposition

	Horizon (months)	macroeconomic shocks					exchange rate shocks
		output	inflation	short rate	money	**all**	**all**
g_{US}	6	88.71	1.75	3.31	3.05	**96.83**	**3.17**
	24	78.95	3.58	7.41	4.00	**93.93**	**6.07**
π_{US}	6	5.12	87.92	0.97	1.70	**95.70**	**4.30**
	24	6.71	83.32	1.32	4.10	**95.45**	**4.55**
s_{US}	6	1.77	3.94	88.80	2.10	**96.61**	**3.39**
	24	2.03	6.55	87.15	1.09	**96.81**	**3.19**
m_{US}	6	3.09	1.44	2.15	89.60	**96.29**	**3.71**
	24	4.32	3.10	3.67	85.01	**96.10**	**3.90**
g_{EA}	6	80.68	12.03	3.21	2.82	**98.73**	**1.27**
	24	70.60	15.84	4.03	4.22	**94.69**	**5.31**
π_{EA}	6	5.60	80.95	3.96	6.25	**96.75**	**3.25**
	24	6.17	72.35	5.21	9.89	**93.63**	**6.37**
s_{EA}	6	2.47	1.07	89.07	3.56	**96.17**	**3.83**
	24	4.92	1.55	81.08	6.02	**93.57**	**6.43**
m_{EA}	6	2.81	4.47	2.23	87.40	**96.91**	**3.09**
	24	4.66	7.34	4.13	79.04	**95.17**	**4.83**
g_{JA}	6	87.53	4.65	3.82	2.03	**98.03**	**1.97**
	24	78.71	6.38	7.19	3.16	**95.43**	**4.57**
π_{JA}	6	1.88	84.45	5.56	3.78	**95.66**	**4.34**
	24	1.06	77.65	12.48	5.86	**97.06**	**2.94**
s_{JA}	6	6.58	1.35	88.09	1.59	**97.61**	**2.39**
	24	9.91	3.85	82.54	1.85	**98.15**	**1.85**
m_{JA}	6	2.50	7.49	2.87	80.06	**92.93**	**7.07**
	24	3.08	8.66	5.11	73.41	**90.26**	**9.74**
g_{UK}	6	84.66	3.49	4.18	4.84	**97.17**	**2.83**
	24	80.62	3.93	6.57	6.39	**97.51**	**2.49**
π_{UK}	6	1.53	80.91	8.37	0.88	**91.69**	**8.31**
	24	2.66	59.71	10.58	2.13	**75.08**	**24.92**
s_{UK}	6	2.97	2.64	88.80	0.17	**94.59**	**5.41**
	24	5.35	3.41	80.81	0.19	**89.76**	**10.24**
m_{UK}	6	2.67	1.05	4.49	90.03	**98.23**	**1.77**
	24	3.93	2.58	5.66	86.07	**98.24**	**1.76**
g_{CA}	6	85.56	5.66	3.69	2.51	**97.43**	**2.57**
	24	77.29	9.67	7.59	1.33	**95.87**	**4.13**
π_{CA}	6	2.16	90.44	4.49	2.05	**99.14**	**0.86**
	24	3.74	83.12	7.72	3.11	**97.69**	**2.31**
s_{CA}	6	2.69	4.27	86.19	3.25	**96.41**	**3.59**
	24	4.37	5.61	78.03	5.11	**93.13**	**6.87**
m_{CA}	6	3.70	0.63	3.32	88.25	**95.91**	**4.09**
	24	4.84	1.06	2.51	85.13	**93.54**	**6.46**

(continued)

(Table 5 continued)

	Horizon (months)	macroeconomic shocks					exchange rate shocks
		output	inflation	short rate	money	**all**	**all**
e_{EA}	6	8.47	4.65	1.87	3.63	**18.91**	**81.09**
	24	8.41	7.13	4.55	2.32	**22.40**	**77.60**
e_{JA}	6	6.54	2.56	1.75	3.21	**14.07**	**85.93**
	24	10.24	6.17	1.97	5.57	**23.96**	**76.04**
e_{UK}	6	1.58	1.90	6.85	2.09	**12.41**	**87.59**
	24	1.78	1.93	12.49	4.45	**20.65**	**79.35**
e_{CA}	6	5.19	1.09	7.12	2.87	**16.26**	**83.74**
	24	9.16	2.86	11.04	3.88	**26.94**	**73.06**

The table reports for each log volatility variable, i.e. real output growth rate (g), inflation rate (π), short-term nominal interest rate (s), nominal money growth rate (m) for the five countries investigated (the US (US), the euro area (EA), Japan (JA), the UK (UK), and Canada (CA)), and nominal exchange rate returns for the euro, the Japanese yen, the British pound and the Canadian dollar against the US dollar (e), the median forecast error variance decomposition at the six-month and two-year horizons obtained from the structural VMA representation of the $FI-FVAR$ model, following the thick modelling estimation strategy. For each log volatility variable the table shows the percentage of forecast error variance attributable to each macroeconomic shock ("output", "inflation", "short rate" and "money") together with their sum ("all", in bold). The last column reports the percentage of the forecast error variance attributable to all the exchange rate shocks ("exchange rates", "all", in bold).

Table 6: Trade-off analysis

	Long-term			
	€$/US\$$	$YEN/US\$$	$GBP£/US\$$	$CA\$/US\$$
g_{US}	0.25	0.19	0.88	0.81
g_i	0.33	0.93	0.18	0.77
π_{US}	0.16	0.37	0.54	0.17
π_i	0.31	0.58	0.11	0.48
s_{US}	0.03	0.29	0.21	0.85
s_i	0.11	0.04	0.04	0.96
m_{US}	0.88	0.45	0.78	0.46
m_i	0.69	0.22	0.50	0.58
	Medium-term			
	€$/US\$$	$YEN/US\$$	$GBP£/US\$$	$CA\$/US\$$
g_{US}	0.70	0.48	0.70	0.63
g_i	0.50	0.22	0.87	0.51
π_{US}	0.83	0.47	0.31	0.41
π_i	0.53	0.68	0.60	0.86
s_{US}	0.40	0.49	0.17	0.51
s_i	0.34	0.49	0.36	0.29
m_{US}	0.47	0.82	0.20	0.54
m_i	0.77	0.70	0.80	0.73

The table reports the proportion of variance for each macroeconomic volatility variable involved in the trade-off. Hence, i.e. $i = EA$ for the €$/US\$$, JA for the $YEN/US\$$, UK for the $GBP£/US\$$, CA for the $CA\$/US\$$. The variables investigated are log volatilities for real output growth rates (g), inflation rates (π), short-term nominal interest rates (s), nominal money growth rates (m). Columns correspond to the four exchange rates.

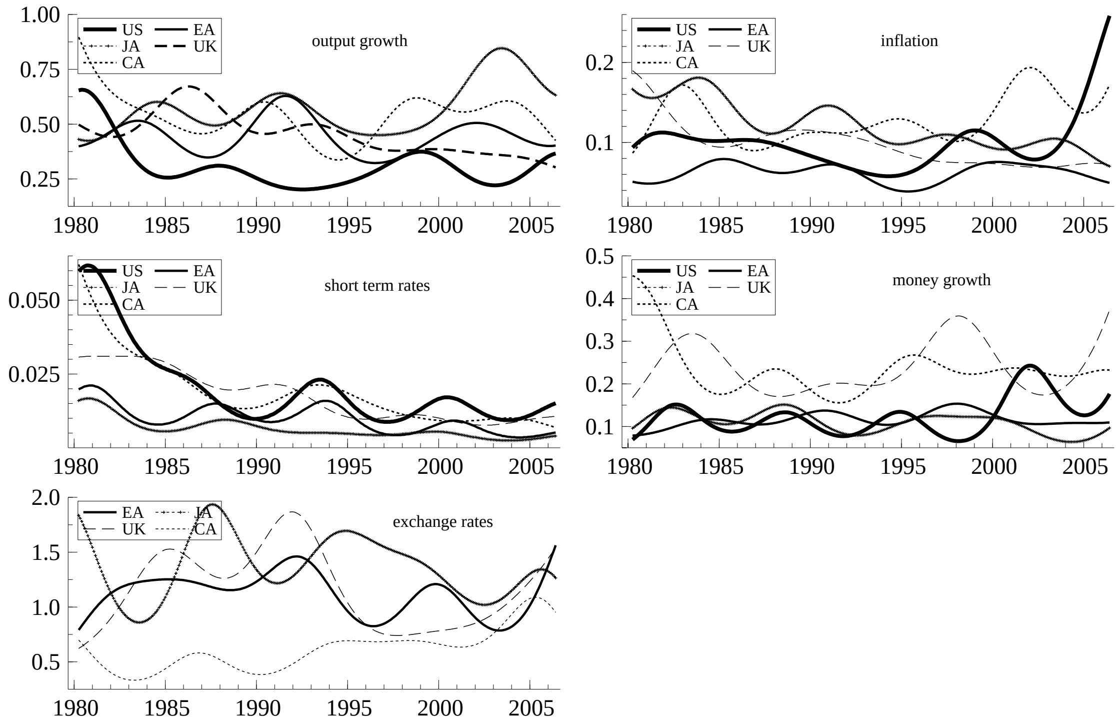

Figure 1: Estimated break processes for output growth, inflation, short term rate, nominal money growth, and exchange rate volatility; United States (US), euro area (EA), Japan (JA), the United Kingdom (UK), Canada (CA).

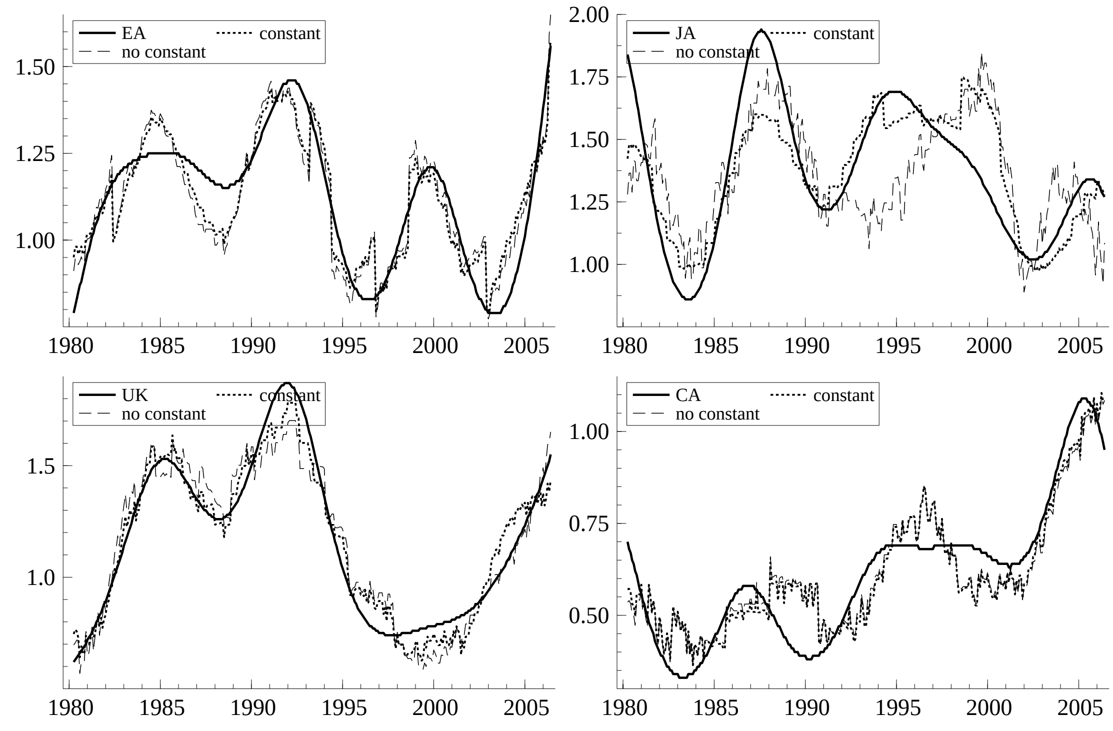

Figure 2: Estimated exchange rates volatility break processes: actual and reconstructed processes, including (constant) or excluding (no constant) an intercept component in the OLS regression (€/US$ exchange rate (EA), yen/US$ exchange rate (JA), GBP£/UK$ exchange rate (UK), Canadian $/US$ (CA)).