Structural Breaks in the International Transmission of Inflation

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Abstract

To shed light on changes in international inflation, this paper proposes an iterative procedure to discriminate between structural breaks in the coefficients and the disturbance covariance matrix of a system of equations, allowing these components to change at different dates. Conditional on these, recursive procedures are proposed to analyze the nature of change, including tests to identify individual coefficient shifts and to discriminate between volatility and correlation breaks. Using these procedures, structural breaks in monthly cross-country inflation relationships are examined for major G-7 countries (US, Euro area, UK and Canada) and within the Euro area (France, Germany and Italy). Overall, we find few dynamic spillovers between countries, although the Euro area leads inflation in North America, while Germany leads France. Contemporaneous inflation correlations are generally low in the 1970s and early 1980s, but inter-continental correlations increase from the end of the 1990s, while Euro area countries move from essentially idiosyncratic inflation to co-movement in the mid-1980s.

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1. Introduction

Understanding the international transmission of price shocks is of obvious importance for the conduct of monetary policy, when the principal aim of that policy is to keep inflation at low and stable levels. Although monetary policy decisions are (generally) taken individually by central banks in each country, a number of recent contributions have uncovered evidence of a strong international dimension to the inflation experience of developed countries (Ciccarelli and Mojon, 2009, Monacelli and Sala, 2009, Mumtaz and Surico, 2008, Neely and Rapach, 2008). This finding challenges modern macroeconomic theories of monetary policy and the inflation process, which focus primarily on domestic factors and the role of the central bank. Consequently, Borio and Filardo (2007) refer to the need for inflation models to become more "global-centric", while Wang and Wen (2007) conclude that the strong international correlations observed are unlikely to be monetary phenomena.

There is, however, one important contemporary instance where monetary policy is taken on a cross-country basis, which is that of the Euro area where the common monetary policy regime came into operation from January 1999. Although Wang and Wen (2007) conduct a robustness analysis comparing these countries with other countries in their sample, this does not reflect Euro area actuality in that they conduct a constant-parameter analysis starting in the 1970s. Indeed, although some studies (including Ciccarelli and Mojon, 2009, Mumtaz and Surico, 2008, Neely and Rapach, 2008) allow for some time variation in cross-country correlations, a systematic examination of the date and nature of structural breaks in international inflation linkages has not yet been conducted.

In contrast to the relative lack of analyses of changes in international inflation linkages, an important focus of univariate cross-country comparisons has been the nature and timing of structural breaks, with changes in the level of inflation or its persistence often linked with monetary policy change; see, for example, Altissimo *et al*. (2006), Benati (2008), Cecchetti and Debelle (2006), Halunga, Osborn and Sensier (2009) or O'Reilly and Whelan (2005). Nevertheless, the clustering of break dates documented by Bataa *et al*. (2008), Corvoisier and Mojon (2005), Levin and Piger (2004), and others, suggests that in addition to changes in 'marginal' characteristics such as the level and persistence of inflation, changes may also have occurred in the inflation transmission process.

The current paper addresses this issue, by examining the changes in inflation linkages across the major countries of the international economy over the period from 1973 to 2007. In our analysis we specifically set out to disentangle the roles of spillovers versus contemporaneous correlations of price shocks. For this purpose, we build on the multivariate procedure of Qu and Perron (2007) to test for breaks in these cross-country interactions, as represented by the dynamic coefficients and conditional covariance matrix of a vector autoregressive (VAR) specification for the inflation series. A key feature of our approach is that it does not restrict the structural breaks in the VAR coefficients and in the covariance matrix to occur at the same dates, for which we develop a new iterative procedure taking account of both types of breaks separately. A further methodological contribution is that we decompose covariance matrix breaks into changes in volatility and correlations. Consequently, we are able to more accurately pinpoint the nature of changes in international inflation linkages.

The predominant econometric method in recent analyses of inflation linkages is based on dynamic factor models, including the contributions of Ciccarelli and Mojon (2009), Mumtaz and Surico (2008) and Neely and Rapach (2008). The extraction of factors, however, requires quite strong assumptions about the dynamics and covariances linking the series; typically the underlying parameters are assumed constant, although the analysis of Mumtaz and Surico (2008) permits them to evolve as random walk processes. By relating inflation in specific countries with the extracted factors, this methodology can provide important insights into the broad nature of change. However, it is not designed to either date breaks or to disentangle the roles of spillovers versus contemporaneous correlations. Our VAR-based methodology is better suited for this purpose, while it also is able to account for the presence of structural breaks in inflation volatility.

To be explicit, we examine links between inflation in Canada, the Euro area, the UK and the US on the one hand[1] and between France, Germany and Italy on the other. The latter group is, of course, of interest as it consists of the major Euro area countries, for which our analysis sheds particular light on the impact of monetary union. Our key findings can be summarized as follows. Firstly, spillovers are almost completely absent in international inflation. With few exceptions, price shocks occurring in one country are not transmitted to other countries. The most notable exception is that the Euro area appears to lead inflation in the US and Canada (but not the UK). Secondly, contemporaneous cross-correlations are practically zero before 1984, such that inflation during the first part of the sample period can be characterized as idiosyncratic. This finding is surprising, as the large oil price shocks during the 1970s are commonly thought to have affected price levels similarly at a global level. For the Euro area countries, correlations undergo substantial increases around 1984 and 1996, with the latter being attributed to the run-up to the introduction of the common currency. For the inter-continental analysis, we find that correlations remained stable (and close to zero) until 1999, but then jump sharply to levels between 0.3 and 0.6. Indeed, the US/Euro area correlation of 0.6 during the post-1999 period is even larger than the correlation between France and Germany. Thirdly, findings from univariate analyses that inflation persistence falls in all countries over this period while volatility is unstable are confirmed in our multivariate setting.

The paper is organized as follows. Section 2 discusses our data and some of its cross-country features. The methodology we employ is then outlined in Section 3, with further

[1] As in the international business cycle analysis of Doyle and Faust (2005), Japan is omitted since it behaved differently from these other countries during the 1990s. Although, ideally, a single VAR would be examined for all countries under study, this is infeasible given the computationally-intensive nature of the analysis undertaken.

details provided in an Appendix. Our principal results relating to changes in international inflation linkages are presented in Section 4. Finally, Section 5 concludes.

2. Data

Although previous studies generally use quarterly data, our analysis is at the monthly frequency in order to more accurately separate contemporaneous and dynamic inflation linkages. The monthly frequency is also relevant because it reflects the typical frequency of monetary policy decisions in the countries under study.

We analyze consumer price inflation in individual G-7 countries and the Euro area, with inflation computed as 100 times the monthly difference of the log consumer price index (CPI) over the period 1973.03-2007.12. A detailed description of the dataset is provided in Bataa, Osborn, Sensier and van Dijk (2008), where a univariate analysis is undertaken. To stress the salient features of that analysis, these series show substantial communality in persistence break dates during the late 1980s and early 1990s, with almost universal volatility breaks in the early 1980s and many in the mid to late 1990s. For most countries (with the exception of Italy) inflation persistence falls to zero in the latter part of the sample. Inflation volatility decreases substantially in the early 1980s, but increases again for the US, Canada and France around 1999. This communality of break dates points to the value of a VAR analysis of the dynamics and co-movement of inflation, as undertaken below.

The results in the present paper are based on data series that have been cleaned from seasonality, together with outliers and mean breaks, in an iterative procedure that allows for time-varying seasonality, unconditional mean, persistence and conditional volatility, as well as outliers (see Bataa *et al*., 2008 for charts of each stage of this process). Removing mean breaks prior to forming a VAR follows the suggestion of Ng and Vogelsang (2002)[2]. It is also important in our context, since there is widespread agreement that inflation has experienced mean changes, with these often associated with monetary policy changes in individual countries, but the evidence for other breaks is less clear-cut (see, for example, Cecchetti and Debelle, 2006, Levin and Piger, 2004).

As noted in the Introduction, our analysis considers two systems, one consisting of Canada, the Euro area, the UK and the US, and the other of the individual Euro area countries of France, Germany and Italy. The former system sets out to capture the interactions between G-7 economies, with the Euro area treated as a single entity in order to abstract (as far as possible) from the changes in inflation linkages between G-7 countries that are now members of the European monetary union[3]. These intra-Euro area changes are the focus of the system for France, Germany and Italy. Japan is excluded from our analysis because, as documented by many previous analyses, it has been largely disconnected from other G-7 economies over this period.

3. Methodology

Our multivariate analysis focuses on changes in dynamic spillovers and contemporaneous correlations of inflation. Nevertheless, volatility changes may hamper such an analysis and hence we adopt a systematic approach in order to identify structural breaks in dynamics, volatility and correlations of inflation.

The framework for our analysis is a conventional VAR system for n countries

[2] This approach is similar to the one taken by Blanchard and Quah (1989) who remove the mean breaks a priori to the VAR analysis, but the difference is that we let the data determine endogenously the number and location of the break dates, not imposing them exogenously.
[3] While recognizing that it is a mis-representation of the nature of economic and monetary union, nevertheless for convenience we later often refer to the Euro area as a “country”.

$$\mathbf{y}_t = \sum_{i=1}^{p} \mathbf{A}_i \mathbf{y}_{t-i} + \mathbf{u}_t \tag{1}$$

where $\mathbf{y}_t = [y_{1,t},..., y_{n,t}]'$ and no intercept is required since all series are mean-corrected. The error term $\mathbf{u}_t$ in (1) has mean zero and covariance matrix $E(\mathbf{u}_t\mathbf{u}_t') = \mathbf{\Sigma}$, and is temporally uncorrelated. Further defining **D** to be the diagonal matrix containing the standard deviations of $\mathbf{u}_t$ and **P** to be the corresponding correlation matrix, then (by definition) $\mathbf{\Sigma} = \mathbf{D}\ \mathbf{P}\ \mathbf{D}$. Our methodology seeks to date structural breaks in each of the three components of (1), namely inflation spillovers as captured by the VAR coefficients $\mathbf{A}_i$ (i = 1, …, p), inflation volatility measured by **D**, and contemporaneous inflation correlations in **P**. In addition to dating any such breaks that may have occurred, we also examine the statistical significance of international relations by conducting inference on $\mathbf{A}_i$ and **P**.

Our analysis of structural breaks builds upon the recent methodology of Qu and Perron (2007) to test for mean and covariance breaks in a VAR system[4]. The Qu and Perron (2007) methodology provides us with tools to deal with three scenarios, namely breaks occurring simultaneously in both the VAR coefficients $\mathbf{A}_i$ and the covariance matrix $\mathbf{\Sigma}$, breaks occurring only in the VAR coefficients or breaks occurring only in the covariance matrix. Although there is widespread evidence suggesting the possibility of breaks in both components for international inflation, these need not occur at the same dates (and, consequently, the numbers of breaks need not even be the same). Indeed, given the previous literature concerning the univariate properties of inflation, volatility declines might be anticipated in the early 1980s (see, e.g., Sensier and Dijk, 2004) while those in dynamics may have more likely occurred in the early 1990s (see, e.g. Cogley and Sargent, 2002, 2005). In a univariate context, Pitarakis (2004) shows there is a potential problem of misspecified breaks

[4] This procedure improves upon the commonly used single-equation approach of Bai and Perron (1998, 2003) and the approach of Bai, Lumsdaine and Stock (1998) for breaks in VAR systems. The main advantages of the Qu and Perron (2007) methodology is that it is valid under more general assumptions, it is not a requirement that the regressors are independent of the errors at all leads and lags in the presence of heteroscedasticity and/or autocorrelation and in a multivariate setup it allows for multiple breaks in contrast to Bai, *et al.* (1998).

in dynamics affecting inferences on structural stability in volatility and *vice versa*; see also our univariate analysis in Bataa *et al*. (2008) which examines the same inflation series as the present paper.

For those reasons, we develop a new iterative procedure to test for (separate) breaks in the VAR coefficients and the covariance matrix. Since this procedure relies heavily on the Qu and Perron (2007) tests, these are first outlined in subsection 3.1. Subsection 3.2 then describes our iterative decomposition of breaks as changes in $\mathbf{A}_i$ and $\mathbf{\Sigma}$, followed by separation of volatility and correlation breaks for the latter. The nature of our hypothesis tests concerning the inflation linkages are then discussed in subsections 3.3 and 3.4. Further details of these procedures can be found in the appendix.

3.1 Tests for dynamic and covariance breaks

Prior to testing, the order p of the VAR in (1) is selected using the Hannan-Quinn criterion over the entire sample period. Then, using the procedure of Qu and Perron (2007), we check the stability of the VAR coefficients against the possibility of $m \leq M$ breaks, where m is unknown and the maximum number of breaks M is pre-specified. This is implemented as a test of the null hypothesis $H_0 : \mathbf{A}_{i,j} = \mathbf{A}_{i,0}$ (j = 1, ..., $m+1$; i = 1,...,p) in

$$\mathbf{y}_t = \sum_{i=1}^{p} \mathbf{A}_{i,j} \mathbf{y}_{t-i} + \mathbf{u}_t, \qquad (2)$$

for $t = T_{j-1} + 1, \ldots, T_j$, j = 1, …, m+1, where T_j denote the break dates marking the m subsamples, with $T_0 = 0$ and $T_{m+1} = T$; T being the total sample size, and where u_t can be heteroskedastic.

The overall null of no breaks is tested using the 'double maximum' statistic

$$WD\max F_T(M) = \max_{1 \leq m \leq M} a_m \left[\sup_{(\lambda_1,\ldots,\lambda_m \in \Lambda_\ell)} F_T(m, q, \varepsilon) \right], \qquad (3)$$

where λ_j (j = 1, …, m) indicate possible break dates as fractions of the sample size, with $0 < \lambda_1 < ... < \lambda_m < 1$ and $T_j = [T\lambda_j]$, and Λ_ε denotes all permissible sample partitions satisfying the requirement that a fraction of at least ε of the sample is contained in each segment, for some $0 < \varepsilon < 1$. The parameter $a_m = c(\alpha,1)/c(\alpha,m)$ with $c(\alpha,m)$ the asymptotic critical value (at a significance level of 100α percent) of the supremum statistic $\sup_{(\lambda_1,...,\lambda_m \in \Lambda_\ell)} F_T(m,q,\varepsilon)$ against a specific number of m breaks. For a total of q VAR coefficients in (1), all of which are allowed to change,

$$F_T(m,q,\varepsilon) = \left(\frac{T-(m+1)q}{m}\right)\hat{\boldsymbol{\beta}}'\mathbf{R}'[\mathbf{R}\hat{\mathbf{V}}(\hat{\boldsymbol{\beta}})\mathbf{R}']^{-1}\mathbf{R}\hat{\boldsymbol{\beta}}, \tag{4}$$

is a Wald-type test statistic for structural change at m known dates, $\hat{\boldsymbol{\beta}}$ is the stacked vector of estimated VAR coefficients given the m breaks with estimated robust covariance matrix $\hat{\mathbf{V}}(\hat{\boldsymbol{\beta}})$ [5], and $\mathbf{R}$ is the non-stochastic matrix such that $(\mathbf{R}\boldsymbol{\beta})' = (\boldsymbol{\beta}_1' - \boldsymbol{\beta}_2',...,\boldsymbol{\beta}_m' - \boldsymbol{\beta}_{m+1}')$ where $\boldsymbol{\beta}_j$ is the vector of coefficients in the j-th segment.

If the *WD*max test of (3) rejects the null of no breaks, a sequential *F*-type test is used to determine the number of breaks and their locations. In particular, this procedure makes use of the test statistic

$$SEQ_T(l+1|l) = \max_{1\le j\le l+1}\left[\sup_{\tau\in\Lambda_{j,\varepsilon}} F_T(\hat{T}_1,...,\hat{T}_{j-1},\tau,\hat{T}_j,...,\hat{T}_l) - F_T(\hat{T}_1,...,\hat{T}_l)\right], \tag{5}$$

where $\Lambda_{j,\varepsilon} = \{\tau; \hat{T}_{j-1} + (\hat{T}_j - \hat{T}_{j-1})\varepsilon \le \tau \le \hat{T}_j + (\hat{T}_j - \hat{T}_{j-1})\varepsilon\}$, and F_T is defined as in (4). The statistic in (5) can be used to test the null of l breaks against the alternative of l+1 breaks, by testing for the presence of an additional break in each of the segments defined by the break dates $(\hat{T}_1,\hat{T}_2,...,\hat{T}_l)$ obtained from estimating the model with l breaks. The test is applied sequentially for l = 0, 1, … until it fails to reject the null hypothesis of no additional break.

[5] As there are potential breaks in the variance covariance matrix in the residuals of (2), we use the Heteroskedasticity Consistent (HC) version of the tests when testing for the breaks in the conditional mean dynamics.

Having obtained a first estimate of the number of structural breaks using (5), the break dates and VAR coefficients are estimated by maximizing a Gaussian quasi-likelihood function using the efficient dynamic programming algorithm outlined in Bai and Perron (2003) and Qu and Perron (2007). This also allows the construction of confidence intervals for the break dates.

Testing for breaks in the conditional covariance matrix $\mathbf{\Sigma}$ proceeds along similar lines as the procedure for breaks in dynamics described above. First, the null hypothesis of no breaks, that is $H_0 : \mathbf{\Sigma}_j = \mathbf{\Sigma}_1$ (j = 2, ..., $m+1$) for an unknown $m \leq M$ number of breaks, is tested using a 'double maximum' likelihood ratio-type test statistic. In particular, the Sup$F$ statistic in (3) is replaced by the SupLR statistic defined as

$$\sup LR_T(m,q,\varepsilon) = \sup_{(\lambda_1,\ldots,\lambda_m \in \Lambda_\ell)} 2\ln\left(\frac{\hat{L}_T(T_1,\ldots,T_m)}{\widetilde{L}_T}\right), \tag{6}$$

where $\ln \hat{L}_T(T_1,\ldots,T_m) = -\frac{T}{2}(\ln 2\pi + 1) - \sum_{j=1}^{m+1} \frac{T_j - T_{j-1}}{2} \ln|\hat{\mathbf{\Sigma}}_j|$ and $\mathbf{\Sigma}_j = \frac{1}{T_j - T_{j-1}} \sum_{t=T_{j-1}+1}^{T_j} \hat{\mathbf{u}}_t \hat{\mathbf{u}}_t'$

with $\hat{\mathbf{u}}_t$ (t = 1, …., T) the residual vector from (2), while ~ represents the corresponding quantities computed under the null hypothesis of no covariance matrix breaks. Although we use m to denote the number of covariance matrix breaks, as for the VAR coefficient break test in (3), we emphasise that neither the number nor dates of these two types of breaks are restricted to be the same.

If the null hypothesis of no covariance matrix breaks is rejected, the number of breaks is obtained using a similar procedure to that for the VAR coefficients, with the sequential test in (5) replaced by

$$SEQ_T(l+1|l) = \max_{1 \leq j \leq l+1}\left[\sup_{\tau \in \Lambda_{j,\varepsilon}}\left(\ln\left(\frac{\hat{L}_T(T_1,\ldots,T_{j-1},\tau,T_j,\ldots,T_l)}{\hat{L}_T(T_1,\ldots,T_l)}\right)\right)\right] \tag{7}$$

Again the break dates are then estimated by maximizing a Gaussian quasi-likelihood function, which is also used for computing confidence intervals for these dates.

For the coefficient and covariance matrix analyses, the maximum number of breaks, M, needs to be specified, as well as the minimum fraction ε of the sample in each regime. Critical values of the tests depend on both the number of coefficients allowed to change and ε. In general ε has to be chosen large enough for the tests to have approximately correct size and small enough for them to have decent power. Moreover, when the errors are potentially heteroskedastic, ε has to be larger than when this feature is absent. In order to balance these issues in relation to our sample size, we set $\varepsilon = 0.15$ with M = 5. Finally, we use a significance level of 5 percent for all tests [6].

3.2 Disentangling dynamic, volatility and correlation breaks

We adopt an iterative procedure to disentangle breaks in the VAR coefficients and in the conditional covariance matrix, analogous to that in Bataa *et al*. (2008) in a single-equation context. This allows for the possibility that the numbers of breaks in $\mathbf{A}_i$ and $\mathbf{\Sigma}$ are different, and for breaks to occur at different dates. This approach initially examines breaks in the VAR coefficients using heteroskedasticity robust tests, as outlined in the previous subsection. Conditional on the estimated break dates for $\mathbf{A}_i$, we then test for breaks in the covariance matrix. Conditional on the estimated break dates for $\mathbf{\Sigma}$, breaks in VAR dynamics are again tested. However, rather than employing heteroskedasticity robust tests for $\mathbf{A}_i$, a feasible generalized least squares (GLS) procedure is now employed which exploits the covariance break information. This process is repeated, iterating between tests for breaks in $\mathbf{A}_i$ (i = 1, …, p) and in $\mathbf{\Sigma}$ until convergence, with the existence of identified breaks verified using finite sample inference; see Appendix A and subsection 3.4 below for details.

[6] Utilizing a 10 percent significance level yields results that are qualitatively very similar.

As already discussed, identified covariance breaks could originate from changes in either volatility or correlations. For example, an increase in covariance could result from an increase in correlation or from a decline in volatility. Since these have quite different implications in terms of the nature of international inflation linkages, identifying volatility or correlation as the source of a covariance break is of crucial importance. Indeed, correlation changes have been the focus of recent interest (see, for example, Neely and Rapach, 2008, Ciccarelli and Mojon, 2009, or Mumtaz and Surico, 2008).

Using the identity $\mathbf{\Sigma} = \mathbf{D}\ \mathbf{P}\ \mathbf{D}$, we distinguish between volatility and correlation changes, represented by $\mathbf{D}$ and $\mathbf{P}$ respectively, conditional on given covariance matrix break dates (obtained after iterating between dynamic and covariance breaks). Essentially, volatility is captured by squared residuals, with finite sample inference used to examine constancy of $\mathbf{D}^2$ over the specified covariance regimes, with a general to specific procedure used to eliminate any insignificant volatility breaks. Conditional on significant volatility breaks, the VAR residuals are standardized and breaks in the correlation matrix $\mathbf{P}$ are examined by applying finite sample bootstrap inference to the statistic of Jennrich (1970). The test is applied initially to each break date identified for $\mathbf{\Sigma}$. If not all breaks in $\mathbf{P}$ are significant (at five percent), the least significant is dropped and the procedure repeated until all remaining correlation breaks are significant. Details are again discussed in the Appendix.

3.3 Individual coefficients breaks, spillover and contemporaneous correlation tests

The coefficient breaks resulting from the analysis outlined in subsection 3.2 apply to the VAR system as a whole. However, it is also of interest to identify whether these relate to persistence changes in individual inflation series or in the causality pattern across countries. To shed light on the source of change, we employ a general to specific approach to test the equality of individual VAR coefficients across sub-samples. This is based on a conventional F-test conditional on the break dates, as in Doyle and Faust (2005). The test employs the statistic of (4), but with the restriction matrix **R** defined as $(\mathbf{R}\hat{\boldsymbol{\beta}})' = (0,..,\hat{a}^{(j)}_{(h,k),1} - \hat{a}^{(j-1)}_{(h,k),1}, \hat{a}^{(j)}_{(h,k),2} - \hat{a}^{(j-1)}_{(h,k),2},...,\hat{a}^{(j)}_{(h,k),p} - \hat{a}^{(j-1)}_{(h,k),p},...,0)$, where $\hat{a}^{(j)}_{(h,k),i}$ is the $(h,k)^{\text{th}}$ element of $\mathbf{A}_i$ matrix in the j^{th} regime. Note, therefore, that this test applies to the set of specific VAR coefficients for the impact of inflation in country *k* on that of country *h* at all lags $i = 1, ..., p$, with regime *j* compared to $j - 1$. For this purpose, the analysis is conditional on the estimated VAR break dates obtained from the entire system, with the general case following Doyle and Faust (2005) in allowing all coefficients not under study to change at these dates. However, in testing only adjacent regimes, individual *F*-tests may have relatively low power. Therefore, this *F*-test is employed in a recursive procedure in order to increase the parsimony of the model. Specifically, we compute the *F*-test for all *j* breaks and elements (*h*,*k*), and remove the specific break for a particular coefficient that renders the highest *p*-value, and then re-compute the other *F*-tests. We repeat this until all remaining breaks are significant at the five percent level.

In addition, spillovers (or Granger causality) between the inflation series are examined. Such an analysis could be applied to the sub-periods identified by the breaks in the autoregressive dynamics of the system, as discussed in subsection 3.2. However, since not all coefficients may change at any system break date, this would imply unnecessary sample splitting, thus reducing the power of the test. Therefore, this spillover (causality) analysis conditions on the significant breaks for individual coefficients, using the procedure just described.

International linkages are revealed through the correlations of the disturbances in (1), and it is relevant to examine whether a specific country is contemporaneously influenced by inflation shocks originating in other countries. Since correlation breaks may result in these changing from zero to nonzero (or vice versa), these tests are conducted for each regime for the correlation matrix **P** as identified by the correlation break dates. The test employed is the instantaneous causality test of Lütkepohl (2005).

3.4 Finite sample inference

The initial analysis of dynamic and covariance breaks in the VAR system of (1) employs the asymptotic critical values provided by Qu and Perron (2007). However, conditional on these dates, all breaks (for both the VAR coefficients and the covariance matrix) are confirmed by a finite sample bootstrap analysis. As discussed in more detail in the Appendix, if any individual break yields an empirical *p*-value for the system test that is greater than 5 percent, then the maximum number of breaks is reduced appropriately and the asymptotic analysis of Qu and Perron (2007) is re-applied. Although this finite sample analysis is conditional on the break dates identified at a given stage, nevertheless building it into the iterative procedure that identifies (separate) breaks in $\mathbf{A}_i$ and $\mathbf{\Sigma}$ provides some assurance that the asymptotic procedure does not lead to spurious breaks.

Other test results are also based on finite sample bootstrap analyses. To take account of possible conditional heteroskedasticity of unknown form, as well as avoiding excessive reliance on asymptotic distributions in potentially modest or small sub-samples, tests applied to specific VAR coefficients (including constancy tests applied jointly over lags $i = 1, \ldots, p$) employ a wild bootstrap form of the heteroskedasticity-robust test statistic of (4), as in Hafner and Herwartz (2009). The wild bootstrap has been shown to yield reliable finite sample inference even when applied to data that are homoskedastic (Gonçalves and Kilian, 2004).

4. Results

This section reports results for both a four-country VAR, consisting of Canada, the Euro area, the UK and the US, and also for a Euro area VAR comprising the individual countries of France, Germany and Italy. The lag order selected by the Hannan-Quinn criterion is $p = 1$ in the former case and $p = 2$ in the latter, with the residuals being substantially free from serial correlation and conditional heteroskedasticity[7].

All subsequent results build on Table 1, which shows the dynamic and covariance matrix breaks identified by our iterative procedure (outlined in subsection 3.2). Although relatively few iterations are required for convergence (namely, four for the international VAR and three for the Euro area), nevertheless these iterations and the use of finite sample inference play a role in these results. In particular, the use of the asymptotic procedure of Qu and Perron (2007) with a robust covariance matrix produces no coefficient breaks for the international VAR and five for the Euro area VAR. In the former case, iteration leads to the detection of one coefficient break, while in the latter a break is dropped when finite sample inference (conditional on break dates) is employed. In addition to the break dates, 90 percent confidence intervals, computed using the method of Qu and Perron (2007), are presented in Table 1. It may be noted that these confidence intervals are relatively tight, which may be indicative of abrupt change rather than gradual change over an extended period.

– Table 1 about here –

Interestingly, more changes apparently occur in the covariances capturing contemporaneous inflation linkages across the international system than in the dynamics of the VAR system of (1), with only one break being found for the latter compared with three for the covariances. However, the converse is true for the Euro area countries, where four breaks

[7] The maximum lag order considered is $\text{int}[12(T/100)^{0.25}] = 17$. Diagnostic tests applied are a multivariate extension of the Godfrey and Tremayne (2005) Lagrange Multiplier test for serial correlation robust to (conditional) heteroskedasticity, together with individual equation and system tests for conditional heteroscedasticity, with these applied to the subsamples identified by the coefficient and covariance matrix breaks respectively. Detailed results for these are available from the authors on request.

are uncovered in dynamics and only two in covariances. The more frequent dynamic breaks in the Euro area may reflect the changing monetary affiliations between these countries over our sample period, an issue explored through the further analysis below.

4.1. VAR coefficients and spillover tests

The late 1990 date for the change in international inflation dynamics provided in Table 1 largely coincides with the persistence breaks for the Euro area and the UK found in our univariate analysis of these series (Bataa *et al*., 2008). However, univariate analysis cannot shed light on persistence versus spillover effects.

To study these effects, the restriction that the persistence/spillover effect is unchanged, that is $a_{(h,k),i}^{(j)} = a_{(h,k),i}^{(j+1)}$, $i = 1, \ldots, p$, is imposed whenever the potential break relating to the effect of economy k on economy h at this date is not significant, using the recursive general to specific procedure described in Section 3.3. The results shown in Panel A of Table 2 indicate that individual coefficient changes are almost always insignificant, except those relating to inflation persistence (measured by the sum of the autoregressive coefficients). Indeed, inflation persistence significantly declines for all countries, except Canada. Turning to Panel B, whereas persistence is around 0.4-0.5 and highly significant for the US, UK and the Euro area prior to 1990.12, persistence becomes insignificant for all four countries after this date. Although this is in line with our univariate results (Bataa *et al*., 2008), the finding here indicates that this is not a spurious consequence of ignoring cross-country links. Another notable feature of the results in Panel B is the general lack of international consumer price inflation spillovers, with the notable exception of significantly positive effects from the Euro area inflation to both Canada and the US throughout the period. Interestingly, the results for Canada indicate off-setting changes between own and US inflation around 1990.

– Table 2 about here –

Table 3 shows the corresponding results for the Euro area, for which four dynamic breaks are considered (see Table 1). However, the recursive analysis finds no individual VAR coefficient changes significantly (at 5%) at the final identified VAR break in 2001; consequently, this break is ignored for the spillover analysis of Panel B[8]. Further, although substantial changes sometimes occur in terms of the magnitude of specific VAR coefficients (for example, the effects of France on Italy in both 1980 and 1990, or Germany on France in 2001), these are often not significant at a 5% level when finite sample inference is employed. In fact, only one individual coefficient change is identified in 1980, and two changes in each of 1985 and 1990. While the single significant coefficient change in 1980 concerns the effect of lagged French inflation in the equation for Germany, those detected in 1985 and 1990 all relate to changes in persistence. As for the international VAR, the overall pattern is one of substantial reductions in persistence, although Germany goes through an intermediate stage of a persistence increase. From 1990, persistence is either negative (for France) or insignificantly different from zero (Germany and Italy).

– Table 3 about here –

Once again, and in line with the international VAR of Table 2, there are very few significant inflation spillovers in Panel B of Table 3. Indeed, with the exception of the lagged coefficient for France in the equation for Germany prior to 1985 and the lagged coefficient for Germany in the equation for France for the complete sample period, no spillover effect is significant at 5 percent in these results.

[8] However, this break is significant for the coefficients of both Germany and the own coefficients in the equation for France, when the general model is examined for all coefficients and all adjacent regimes across the VAR coefficient break dates indicated in Table 1.

4.2 Volatility and correlations

After imposing the subsample and causality restrictions implied by the results of Tables 2 and 3, with the latter setting coefficients insignificant at 5% in Panel B of those tables to zero, the results for tests of volatility and correlation breaks are reported in Tables 4 and 5 for the international and Euro area systems, respectively. The break dates considered are those detected for $\mathbf{\Sigma}$ in Table 1.

– Table 4 about here –

Considering first the international VAR, Table 4 provides strong evidence that volatility changes at each of the three identified covariance matrix break dates. Although the variance of inflation declines at the 1983 break, the 1992 break appears to be due primarily to a further reduction for the US. In contrast, the 1999 break leads to substantially higher volatility, particularly in the US and Canada. These results are in line with the univariate ones in Bataa *et al*. (2008).

One of our primary interests is whether cross-country correlation shocks, as measured by the contemporaneous residual correlations, change over time. For the international VAR (Table 4) this occurs only at the third covariance matrix break, in 1999, with all pairs experiencing very substantial increases in correlation at this time. Therefore, this analysis implies that the earlier covariance changes (in 1983 and 1992) are due entirely to volatility shifts. Although the zero correlation test results indicate that none of these countries is isolated from the contemporaneous effects of international inflation post-1999, the increased exposure of the US is particularly notable, having cross-correlations with the Euro area and Canada of 0.59 and 0.65, respectively. Prior to this date, Canada and more particularly the UK appear relatively isolated from such effects.

– Table 5 about here –

The picture for the three largest Euro area countries, shown in Table 5, is a little different, with the conditional inflation volatilities for all three countries being essentially constant after the first covariance matrix break in 1984. Although the volatility changes at the second break date in 1996 are statistically significant, the magnitudes of change are relatively small. More importantly, however, the contemporaneous correlations change significantly at both covariance break dates, with the general pattern being of increasing correlations. In the first correlation subsample (to 1984), all three countries are relatively immune from current price shocks in the other countries. Although Italy and France are still close to being uncorrelated after the first break in 1984, the correlations of both countries with Germany increase substantially from this date. However, the correlation between France and Italy overtakes that between Italy and Germany at the second break in 1996, with a large increase to 0.32. At the same time, the correlation between Germany and Italy remains largely unchanged, while Germany and France experience a further increase in correlation to 0.56 after this break. Note that this correlation is remarkably similar to the US/Euro area correlation of 0.59, despite the very different economic relationships between these.

The high correlations in Tables 4 and 5 at the end of the sample are not a consequence of the imposition of the restrictions on the VAR coefficients, since very similar results are obtained when an unrestricted breaks specification is employed, where all VAR coefficients are allowed to be non-zero and to change at each of the VAR coefficient break dates. This is evidenced by the robustness analysis presented for the two systems in Table 6, where results are shown for the extreme cases of constant VAR coefficients and all VAR coefficients are allowed to change. In short, for the international VAR, although the assumption of constant VAR coefficients (not surprisingly) leads to the identification of 1992 as an additional break date for the contemporaneous international inflation correlations, the post-1999 correlations are almost identical to those reported in Table 4.

– Table 6 about here –

Similarly, Table 6 shows that correlation breaks within the Euro area are generally robust to assumptions about the VAR coefficients. However, with either imposition of constant VAR coefficients or the use of an unrestricted structural break VAR specification, the latter with all coefficients allowed to break at the dates indicated by Table 1 and no restrictions imposed on persistence/spillover effects, the 1996 correlation break becomes insignificant at 5%, albeit marginally so. Nevertheless, the overall pattern of strongly increased correlations is robust across VAR coefficient specifications, with these countries experiencing essentially idiosyncratic contemporaneous inflation movements until 1984, while becoming strongly (positively) correlated with the movement towards monetary integration.

5. Conclusions

This paper provides evidence on the nature of change in cross-country links for monthly consumer price inflation for major industrialized countries, over the period of floating exchange rates (1973 to 2007). For this purpose, we embed the recently developed procedures of Qu and Perron (2007) for determining system breaks within a novel iterative procedure that allows coefficient and covariance matrix breaks to occur at different dates. Another new feature of our analysis is the use we make of finite sample bootstrap inference to identify the individual coefficients that change over time, whether coefficients and contemporaneous correlations differ significantly from zero, and to distinguish volatility breaks from changes in (contemporaneous) correlations.

Three broad conclusions can be drawn from our results. Firstly, although structural stability tests reject constancy of the coefficients in VAR representations of cross-country inflation linkages, dynamic spillover effects play a relatively modest role in these, both in terms of the extent of such spillovers and also in that changes over time are relatively rare. Indeed, the key cross-country effects that we find in the VAR coefficients indicate that Euro area inflation (at an aggregate level) leads inflation in both Canada and the US, while that in Germany leads France, with these effects applying throughout the sample period. At least in these terms, monetary integration in Europe appears to have played no role in altering international inflation dynamics.

The second general conclusion is that significant persistence and volatility changes occur for inflation, verifying the results from univariate studies of Altimisso *et al*. (2006), Benati (2008), Cecchetti and Debelle (2006), O'Reilly and Whelan (2005), and others. Indeed, although mean breaks have been removed from our series, persistence breaks are found for all countries, in both the international and Euro area VARs, with persistence (as measured by the sum of the own lagged coefficients) almost always insignificant in the latter part of the sample.

Our third conclusion is, however, the most important in terms of the light it sheds on the nature of changes in international inflation linkages. That is, after allowing for dynamic effects and volatility shifts, the globalization of inflationary experiences is very evident in the large increases in the contemporaneous correlations of inflation. In the case of the Euro area countries of France, Germany and Italy, the pattern of shifts around 1984 and 1996 may be due to the progress towards monetary union. In this context, the increased correlation of Italy with France from 1996, from whom its inflation was previously isolated, is notable, in addition to the increased correlation between the two leading economies of France and Germany. Nevertheless, some of the most marked correlation changes relate not to individual Euro area countries, but rather to the integration of the US (and Canada) into world inflationary influences. The date at which we find this change occurs, namely 1999, is of course a milestone in terms of European monetary integration. However, if monetary integration is the cause, then more evidence of changed correlations between inflation in Euro area countries themselves may have been anticipated.

While our results agree those of Borio and Filardo (2007), Ciccarelli and Mojon (2009), Mumtaz and Surico (2008) and Neely and Rapach (2008), who all conclude that inflation is a global phenomenon, our analysis emphasizes that this applies to even the largest world economy, namely the US. Although our study is limited to inflation alone, our findings point to the inadequacy of the predominant paradigm whereby economic models treat the US as being closed to world influences.

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Appendix: Algorithms Employed for Inference

A conventional n-equation VAR(p) model for the zero mean vector $\mathbf{y_t}$ can be written as

$$\mathbf{y}_t = (\mathbf{I} \otimes \mathbf{z}_t')\mathbf{S}\boldsymbol{\beta} + \mathbf{u}_t \qquad \text{(A.1)}$$

where $\underset{T\times np}{\mathbf{z}_t} = [\mathbf{y}_{t-1}', ..., \mathbf{y}_{t-p}']'$, $\underset{n\times n}{\mathbf{I}}$ and $\underset{n^2p\times n^2p}{\mathbf{S}}$ are identity matrices and $\underset{n^2p\times 1}{\boldsymbol{\beta}}$ is a vectorized coefficient matrix, $\mathbf{u}_t$ are assumed to be temporally uncorrelated with $E(\mathbf{u}_t\mathbf{u}_t') = \boldsymbol{\Sigma}$. If ℓ breaks occur in the VAR coefficients at dates $T_1^{(B)}, ..., T_\ell^{(B)}$, (A.1) can be written as

$$\mathbf{y}_t = (\mathbf{I} \otimes \mathbf{z}_t')\mathbf{S}\boldsymbol{\beta}_k + \mathbf{u}_t \qquad \text{(A.2)}$$

for $t = T_{k-1}^{(B)} + 1, ..., T_k^{(B)}$ (k = 1, …, ℓ+1; $T_0^{(B)} = 0; T_{l+1}^{(B)} = T;$) and, moreover, if there m covariance matrix breaks at dates $T_1^{(C)}, ..., T_m^{(C)}$, then $E(\mathbf{u}_t\mathbf{u}_t') = \boldsymbol{\Sigma}_j$ for $t = T_{j-1}^{(C)} + 1, ..., T_j^{(C)}$ (j = 1, …, m+1; $T_0^{(C)} = 0; T_{m+1}^{(C)} = T;$). The VAR coefficient and covariance matrix break dates do not necessarily coincide. Prior to the breaks analysis, the VAR lag length p is selected by the conventional HQ information criterion applied to the entire sample period.

The algorithm for detecting breaks in the parameters of (A.1) proceeds in two steps, firstly disentangling breaks in $\boldsymbol{\beta}$ from those in $\boldsymbol{\Sigma}$ (discussed in section A.1) and secondly separating volatility and correlation breaks that contribute to changes in $\boldsymbol{\Sigma}$ (section A.2). In addition, section A.3 describes the test undertaken for zero contemporaneous correlation. All tests are undertaken using a 5 percent empirical significance level.

A.1 Disentangling VAR Coefficient and Covariance Matrix Breaks

The algorithm proceeds as follows:

1. (a) Initialize the coefficient breaks by applying the asymptotic Qu and Perron (2007) multiple breaks procedure using the Eicker-White heteroskedasticity (HC) robust covariance matrix, for a maximum of M = 5 breaks and $100\varepsilon = 15\%$ trimming, to test for breaks in the VAR coefficients $\boldsymbol{\beta}$. Say ℓ coefficient breaks are detected, at dates $T_1^{(B)}, ..., T_\ell^{(B)}$. Obtain $\hat{\boldsymbol{\beta}}_k$ for each regime k = 1, …, ℓ+1, estimated from

observations $t = T_{k-1}^{(B)}+1,\ldots,T_k^{(B)}$ and corresponding residual series $\hat{\mathbf{u}}_t = \mathbf{y}_t - (\mathbf{I}\otimes\mathbf{z}_t')\mathbf{S}\hat{\boldsymbol{\beta}}_k$, ($k$ = 1, …, ℓ+1).

(b) Verify the significance of each break k, k = 1, …, ℓ, identified in (a), through a test of the null hypothesis $\boldsymbol{\beta}_k = \boldsymbol{\beta}_{k+1}$, conditional on all other ℓ-1 breaks. Use the statistic

$$F_B = \left(\frac{T-(\ell+1)pn^2}{\ell}\right)\hat{\boldsymbol{\beta}}'\mathbf{R}'[\mathbf{R}\operatorname{cov}(\hat{\boldsymbol{\beta}})\mathbf{R}']^{-1}\mathbf{R}\hat{\boldsymbol{\beta}} \qquad \text{(A.4)}$$

where $\hat{\boldsymbol{\beta}}$ consists of the vectorized coefficient matrix $\hat{\boldsymbol{\beta}}_k$ for all k = 1, …, ℓ+1, $\mathbf{R}$ is the non-random restriction matrix such that $(\mathbf{R}\hat{\boldsymbol{\beta}})' = (\mathbf{0}',..,\hat{\boldsymbol{\beta}}'^{(k+1)} - \hat{\boldsymbol{\beta}}'^{(k)},...,\mathbf{0}')$ and $\operatorname{cov}(\hat{\boldsymbol{\beta}})$ is the HC robust covariance matrix. Denote the computed statistic as F_B^*. Inference is conducted by comparing this to the empirical distribution of F_B obtained from a bootstrap data generating process (DGP) that employs the restricted VAR parameters, with $\boldsymbol{\beta}_k = \boldsymbol{\beta}_{k+1}$, and a wild bootstrap process for $\boldsymbol{u}_t$. Based on N samples generated using this bootstrap DGP, the empirical p-value is given by $p = \frac{\# I(F_B^* \geq F_B)}{N}$, where I is the indicator function taking the value unity when the condition is satisfied and zero otherwise.

(c) If one or more breaks k = 1, …, ℓ is not significant in step 1(b), the number of coefficient breaks is reduced to $\ell^* = \ell$-1.

(d) Set $\ell = \ell^*$ and estimate the new break dates $T_1^{(B)},\ldots,T_\ell^{(B)}$ via global optimization, using the Qu and Perron (2007) HC procedure. Return to step 1(b) and iterate until the empirical p-values for all ℓ coefficient breaks are individually significant.

2. (a) Based on the residuals $\hat{\mathbf{u}}_t$ obtained from Step 1 (or 3 below), apply the Qu and Perron (2007) asymptotic test for multiple breaks in the covariance matrix $\boldsymbol{\Sigma}$. Say m breaks are found, at dates $T_1^{(C)},\ldots,T_m^{(C)}$, and obtain $\hat{\boldsymbol{\Sigma}}_j$ for each regime j = 1, …, m+1, estimated from observations $t = T_{j-1}^{(C)}+1,\ldots,T_j^{(C)}$.

(b) For each covariance break j = 1, …, m identified in (a), test the null hypothesis $\boldsymbol{\Sigma}_j = \boldsymbol{\Sigma}_{j+1}$, using the usual quasi-likelihood ratio statistic, LR (as employed by Qu and

Perron, 2007), yielding the value $LR*$ and apply a finite sample bootstrap test for equality of the covariance matrices in adjacent regimes j and j+1. The residual vectors $\hat{\boldsymbol{u}}_t$ for $t = T_{j-1}^{(C)}+1, ..., T_j^{(C)}, ..., T_{j+1}^{(C)}$ are randomly i.i.d. resampled, with a wild bootstrap employed in other regimes[9], to create the bootstrap residuals $\hat{\mathbf{u}}_t^*$. Using these bootstrap residuals, together with the ℓ VAR coefficient breaks and the associated estimates found in Step 1 (or 3), a pseudo dataset $\mathbf{y}_t^*$ is generated recursively from a set of randomly chosen starting values. The VAR coefficients and residuals are re-estimated using $\mathbf{y}_t^*$, but for computational feasibility the coefficient break dates are assumed known at $T_1^{(B)}, ..., T_\ell^{(B)}$. Obtain the empirical p-value as $p = \dfrac{\# I(LR^* \geq LR)}{N}$ when N bootstrap replications are undertaken.

(c) If one or more of the covariance matrix breaks examined in step 2(b) is not significant, the number of breaks is reduced to $m^* = m - 1$.

(d) Set $m = m^*$ and obtain new covariance break dates $T_1^{(C)}, ..., T_m^{(C)}$ via global optimization. Return to step 2(b) and iterate until the empirical p-values for all m covariance breaks are individually significant.

(e) Based on the m breaks identified in step (d), obtain $\hat{\mathbf{\Sigma}}_j$ ($j = 1, ..., m$+1).

3. Re-estimate the VAR coefficient breaks using a feasible generalized least squares (GLS) approach, which is achieved by premultiplying each covariance matrix block $t = T_{j-1}^{(C)}+1, ..., T_j^{(C)}$ in (A.1) by $\hat{\mathbf{\Sigma}}_j^{-1/2}$, j = 1, …, m+1, where $\hat{\mathbf{\Sigma}}_j^{-1/2}$ is the inverse square root of the corresponding estimated covariance matrix. Follow the procedure as in step 1, but apply the Qu and Perron (2007) multiple breaks procedure to the VAR coefficient vector $\boldsymbol{\beta}$ assuming a constant disturbance covariance matrix[10].

4. Iterate between steps 2 and 3 until the break dates do not change. (The residuals are always computed from the observed, not standardized, data).

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[9] The wild bootstrap sets $\hat{\mathbf{u}}_t^* = \delta \hat{\mathbf{u}}_t$ where δ is randomly chosen as +1 or -1 with equal probabilities. The use of the wild bootstrap here allows the covariance matrix to differ across regimes, with constant variance imposed by the i.i.d. bootstrap only for regimes j and j+1.

[10] For each break detected, the finite sample test for $\boldsymbol{\beta}_k = \boldsymbol{\beta}_{k+1}$ employs a wild bootstrap of the residuals, since Gonçalves and Kilian (2004) show the cost of using the wild bootstrap instead of an i.i.d. bootstrap is minimal even in the case of i.i.d disturbances.

A.2 Decomposing Covariance Matrix Breaks into Variance and Correlation Breaks

Having estimated coefficient breaks in the VAR of (A.2), with restrictions imposed on the coefficients as indicated by the equality over regimes and persistence/spillover tests on individual VAR coefficients, the residuals are obtained for each regime, so that $\hat{\mathbf{u}}_t = \mathbf{y}_t - (\mathbf{I} \otimes \mathbf{z}_t')\mathbf{S}\hat{\boldsymbol{\beta}}_i$, i = 1, …, ℓ+1,. To start the recursive procedure, set $T_j^{(Cor)} = T_j^{(Vol)} = T_j^{(C)}$, *j=1,…,m,* where $T_j^{(Cor)}$ and $T_j^{(Vol)}$ are correlation and volatility break dates respectively.

Now, consider the identity **Σ** = **D P D**, where **D** is the diagonal matrix of standard deviations of $\mathbf{u}_t$ and **P** is the correlation matrix. Breaks in **D** are considered through the vector of squared residuals. That is, we test $H_0 : \boldsymbol{\mu}_1 = \boldsymbol{\mu}_2$ vs. $H_A : \boldsymbol{\mu}_1 \neq \boldsymbol{\mu}_2$, where $\boldsymbol{\mu}_1$ and $\boldsymbol{\mu}_2$ are the means of vector of squared residuals before and after a single break, respectively. With T_1 and T_2 observations in the respective subsamples, the test statistic used is

$$T^2 = \frac{T_1 T_2}{T_1 + T_2}(\overline{\boldsymbol{\mu}}_1 - \overline{\boldsymbol{\mu}}_2)'\mathbf{S}_{pl}^{-1}(\overline{\boldsymbol{\mu}}_1 - \overline{\boldsymbol{\mu}}_2), \qquad \text{(A.5)}$$

where $\mathbf{S}_{pl} = \frac{1}{T_1 + T_2 - 2}(\mathbf{W}_1 + \mathbf{W}_2)$, $\overline{\boldsymbol{\mu}}_h$, are the sample estimates of $\boldsymbol{\mu}_h$ defined as $\overline{\boldsymbol{\mu}}_h = \sum_{j=k}^{T_h} \hat{\mathbf{u}}_j^2 / T_h$, $\mathbf{W}_h = \sum_{j=k}^{T_h} (\hat{\mathbf{u}}_j^2 - \overline{\boldsymbol{\mu}}_h)(\hat{\mathbf{u}}_j^2 - \overline{\boldsymbol{\mu}}_h)'$, h = 1, 2, k = 1 if h = 1 and $k = T_1 + 1$ otherwise. Under normality, the test is asymptotically distributed as Hotelling's T^2 with n and $T_1 + T_2 - 2$ degrees of freedom, but we report finite sample inference using a bootstrap (see Rencher, 2002, p. 122).

The finite sample algorithm to identify the volatility breaks is proceeds as follows:

1. For each volatility break j = 1, …, m calculate the statistic T^2 for the null hypothesis $\boldsymbol{\mu}_j = \boldsymbol{\mu}_{j+1}$ in adjacent regimes j and j+1 as in (A.5), using residuals calculated from the VAR coefficients.
2. To obtain the finite sample bootstrap distribution of the statistic under the null hypothesis, the residual vectors $\hat{\mathbf{u}}_t$ for $t = T_{j-1}^{(Vol)} + 1, ..., T_j^{(Vol)}, ..., T_{j+1}^{(Vol)}$ are randomly i.i.d. re-sampled, with a wild bootstrap employed in other regimes, to create the bootstrap residuals $\hat{\mathbf{u}}_t^*$. Since we resample a vector of residuals, the contemporaneous correlation structure is kept

intact. Using these bootstrap residuals, together with the associated VAR coefficient estimates, a pseudo dataset $\mathbf{y}_t^*$ is generated recursively from a set of randomly chosen starting values. The VAR coefficients are re-estimated and associated individual coefficient tests applied using $\mathbf{y}_t^*$, but for computational feasibility the coefficient break dates are assumed known at $T_1^{(B)}, ..., T_\ell^{(B)}$. The pseudo residuals and the T^2 statistic as in (A.5) are calculated for $\boldsymbol{\mu}_j = \boldsymbol{\mu}_{j+1}$, assuming volatility break dates are known, yielding the value T^2*. Replicate this procedure N times and obtain the empirical p-value as $p = \frac{\#I(T^2* \geq T^2)}{N}$.

3. If one or more volatility breaks examined in step 2 is not significant, the number of breaks is reduced to $m* = m - 1$ and the least significant break is removed. Then set $m = m*$ and return to step 1. Iterate until the empirical p-values for all m volatility breaks are individually significant.

When testing for breaks in the contemporaneous correlation matrix **P** at a candidate break date, allowing for possible changes in volatility, we use the Jennrich (1970) statistic. Here we wish to test H_0: $\mathbf{P}_1 = \mathbf{P}_2$ versus H_A: $\mathbf{P}_1 \neq \mathbf{P}_2$ where $\mathbf{P}_i$, $i = 1, 2$, are correlation matrices before and after a candidate break date. First let $\overline{\mathbf{P}} = \frac{T_1\hat{\mathbf{P}}_1 + T_2\hat{\mathbf{P}}_2}{T_1 + T_2}$, $\mathbf{Z} = \sqrt{\frac{T_1T_2}{T_1 + T_2}}\,\overline{\mathbf{P}}^{-1}(\hat{\mathbf{P}}_q - \hat{\mathbf{P}}_{q+1})$, and $\boldsymbol{\Delta}$ **is** a matrix with typical element $\{\delta_{x,y} + r_{xy}r^{xy}\}$, where $\delta_{x,y}$ is a Kronecker delta, $\{r_{xy}\}$ and $\{r^{xy}\}$ are typical elements of $\overline{\mathbf{P}}$ and $\overline{\mathbf{P}}^{-1}$, respectively. Then Jennrich (1970) shows the statistic

$$J = \frac{1}{2}tr(\mathbf{Z}'\mathbf{Z}) - dg(\mathbf{Z})\Delta^{-1}dg(\mathbf{Z}), \tag{A.6}$$

where tr and dg denoting the trace and the diagonal, respectively, is asymptotically distributed as $\chi^2(n(n-1)/2)$ under the null hypothesis that the correlation matrix is constant. However, our inference again relies on a bootstrap procedure, which proceeds as follows.

First, we obtain the VAR residuals, allowing for VAR coefficient breaks and imposing zero restrictions (if appropriate), and standardize them as $\hat{\mathbf{u}}_t'^{(Vol)} = \hat{\mathbf{u}}_t'\hat{\mathbf{D}}_h^{-1}$, h = 1, …, v + 1, where v is the number of significant volatility breaks uncovered. Then:

1. For each correlation break j = 1, …, m calculate the Jennrich (1970) statistic J_j as in (A.6).
2. Using an i.i.d. bootstrap for the residual vector at each time period in the adjacent correlation regimes $t = T_{j-1}^{(Cor)} + 1,...,T_{j+1}^{(Cor)}$, and a vector wild bootstrap in other regimes $t \neq T_{j-1}^{(Cor)} + 1,...,T_{j+1}^{(Cor)}$, obtain pseudo residuals, $\hat{\mathbf{u}}_t'^{*}$. Then re-apply the priori-removed volatilities, i.e. $\hat{\mathbf{u}}_t'^{**} = \hat{\mathbf{u}}_t'^{*}\hat{\mathbf{D}}_h$, h = 1, …, v+1. Generate recursively a pseudo dataset $\mathbf{y}_t^*$ using randomly chosen initial values, together with the coefficients $\hat{\boldsymbol{\beta}}_i$, i = 1, …, ℓ and the residuals $\hat{\mathbf{u}}_t^{**}$. Allowing for ℓ coefficient breaks at dates $T_1^{(B)},...,T_\ell^{(B)}$, estimate the VAR coefficients and apply the associated individual coefficient tests using $\mathbf{y}_t^*$ to obtain the bootstrap regime residuals $\hat{\mathbf{u}}_t^{***} = \mathbf{y}_t^* - (\mathbf{I} \otimes \mathbf{z}_t'^{*})\mathbf{S}\hat{\boldsymbol{\beta}}_i^*$, i = 1, …, ℓ. Clear the residuals of volatility changes using the given volatility regime dates, then obtain an empirical correlation break statistic using (A.6) and denote the value obtained as J_j^*. Repeat this process N times and obtain empirical p-value corresponding to the null of no volatility break as $p = \dfrac{\#I(J_j^* \geq J_j)}{N}$.
3. If one or more of the volatility breaks examined in the previous step is not significant, the number of breaks is reduced to m^* = m-1 and the least significant break is removed. Set m = m^* and return to step 1 and iterate until the empirical p-values for all m correlation breaks are individually significant.

A.3 Test for No Contemporaneous Correlation

The test statistic we apply is described in Lütkepohl (2005) as a test for no-instantaneous causality for Gaussian variables. The test statistic employs the vectors of standardized VAR residuals $\hat{\mathbf{u}}_t'^{(Vol)} = \hat{\mathbf{u}}_t'\hat{\mathbf{D}}_h^{-1}$, h = 1, …, v+1, as above. Our algorithm for finite sample inference proceeds as follows:

1. For a correlation regime at $t = T_{q-1}^{(Cor)} + 1,...,T_q^{(Cor)}$ (q = 1, …, k+1), calculate a test statistic for the null hypothesis that the residuals in a particular equation, say c, are not

contemporaneously correlated with residuals in any other equation. In particular, if $\boldsymbol{\Sigma}_u = E(\mathbf{u}_t^{(vol)}\mathbf{u}_t'^{(vol)})$ within this regime, the hypothesis is tested as $H_o : \mathbf{C\sigma} = 0$ against $H_A : \mathbf{C\sigma} \neq 0$, where $\boldsymbol{\sigma} = vech(\boldsymbol{\Sigma}_u)$ and $\mathbf{C}$ is a selection matrix of zeros and unity. The Wald test described by Lütkepohl (2005, p93) for the null has form

$$\lambda = T\,\hat{\sigma}'\mathbf{C}'[2\,\mathbf{C}\,\mathbf{D}_n^+(\hat{\boldsymbol{\Sigma}}_u \otimes \hat{\boldsymbol{\Sigma}}_u)\mathbf{D}_n^{+}{}'\mathbf{C}']\mathbf{C}\hat{\sigma} \qquad \text{(A.7)}$$

where $\mathbf{D}_n^+$ is a $n^2 \times \frac{1}{2}n(n+1)$ duplication matrix . Although Lütkepohl (2005) discusses asymptotic inference based on this statistic, we conduct finite sample inference.

2. For each correlation break q = 1, …, k+1, subdivide the matrix of residuals for a regime into two sub-matrices, one containing the vector of all residuals for equation c, denoted $\hat{u}_t^{(c,Vol)}$ and the other containing the residuals for the remaining n - 1 equations, denoted $\hat{\mathbf{u}}_t^{(\bar{c},Vol)}$, where $\bar{c}$ indicates the complement of c.

   (a) For each $t = T_{q-1}^{(Cor)} + 1, ..., T_q^{(Cor)}$, independently re-sample the residuals from $\hat{u}_t^{(c,Vol)}$ and $\hat{\mathbf{u}}_t^{(\bar{c},Vol)}$ to generate two independent pseudo sub-matrices of residuals $\hat{\mathbf{U}}_q^{*(c,Vol)}$ and $\hat{\mathbf{U}}_q^{*(\bar{c},Vol)}$. Combine these to generate a pseudo residual matrix for regime q, $\hat{\mathbf{U}}_q^{*(Vol)}$.

   (b) For all other regimes, namely for $t \neq T_{q-1}^{(Cor)} + 1, ..., T_q^{(Cor)}$, form $[\hat{\mathrm{u}}_t^{*(Vol)}]' = \delta_3[\hat{\mathrm{u}}_t^{(Vol)}]'$, where δ_3 has the Radamacher distribution (that is, takes plus and minus unity with equal probabilities) and put the results in $\hat{\mathrm{U}}_{\bar{q}}^{*(Vol)}$, where $\bar{q}$ indicates the complement of q. Appropriately concatenate $\hat{\mathbf{U}}_q^{*(Vol)}$ and $\hat{\mathrm{U}}_{\bar{q}}^{*(Vol)}$ to form $\hat{\mathbf{U}}^{*(Vol)}$, which has typical row $[\hat{\mathbf{u}}_t^{*(Vol)}]'$.

   (c) Reapply the volatility effects to form $[\hat{\mathbf{u}}_t^*]' = [\hat{\mathbf{u}}_t^{*(Vol)}]'\hat{\mathbf{D}}_h$ t = 1, …, T and for volatility regimes h = 1, …, v+1 (where v is the number of volatility breaks).

   (d) Generate recursively the pseudo dataset $\mathbf{y}_t^*$ using randomly chosen initial values, together with $\hat{\mathbf{u}}_t^*$ and the VAR coefficients $\hat{\boldsymbol{\beta}}_i$, i = 1, …, ℓ . Use this to estimate the VAR coefficients subject to the restrictions, as above, and obtain the residuals.

Then compute the no instantaneous causality test statistic λ^{*} as in step 1, but with $\hat{\mathbf{D}}_h^{-1}$, h = 1, …, v+1 calculated from the residuals of the VAR estimated from the pseudo dataset.

(e) Repeat steps (a) – (d) N times and obtain the empirical p-value as $p = \frac{\# I(\lambda^* \geq \lambda)}{N}$.

Table 1. Structural Break Test Results

A. International VAR		B. Euro Area VAR	
VAR Coefficients	**Covariance Matrix**	**VAR Coefficients**	**Covariance Matrix**
		1980.02	
		1979.08	
		1980.11	
	1983.07	**1985.05**	**1984.04**
	1980.11	1984.09	1983.02
	1983.10	1985.08	1984.05
1990.12	**1992.04**	**1990.10**	
1989.11	1989.09	1989.11	
1992.02	1992.10	1991.11	
			1996.11
			1995.07
			1998.04
	1999.03	**2001.03**	
	1999.02	2000.01	
	2001.12	2001.11	

Notes: The table shows estimated break dates (first value, in bold) followed by lower and upper bounds of the 90 percent confidence interval for this date, obtained by iteratively applying the Qu and Perron (2007) procedure to the VAR coefficients and disturbance covariance matrix at a 5 percent significance level, as explained in the text. Panel A relates to the International VAR (Canada, the UK, Euro area, and the US) and Panel B to the Euro area VAR (France, Germany and Italy).

Table 2. Individual Coefficient Break and Spillover/Persistence Results: International VAR

		A. Coefficient Breaks				B. Inflation Persistence and Spillovers			
Subsample	Explan. Var.	**Dependent Variable**				**Dependent Variable**			
		Canada	**UK**	**Euro Area**	**US**	**Canada**	**UK**	**Euro Area**	**US**
1973.03 - 1990.10	**Canada**					**-0.16**** 4.8	**0.07** 17.0	**0.05** 27.1	**0.06** 15.4
1990.11 - 2007.12		**0.28**** 1.4	**0.00** 97.5	**0.02** 72.9	**0.11** 25.6	**0.12** 16.9			
1973.03 - 1990.10	**UK**					**-0.02** 74.9	**0.49**** 0.0	**0.03** 42.0	**0.01** 67.6
1990.11 - 2007.12		**0.09** 42.8	**-0.54**** 0.0	**0.05** 43.6	**0.07** 52.3		**-0.05** 65.9		
1973.03 - 1990.10	**Euro Area**					**0.35**** 0.9	**0.03** 59.7	**0.46**** 0.0	**0.28**** 2.6
1990.11 - 2007.12		**0.05** 78.4	**0.13** 52.7	**-0.41**** 0.0	**-0.10** 58.1			**0.05** 57.0	
1973.03 - 1990.10	**US**					**0.12** 8.8	**0.07** 27.5	**0.03** 67.9	**0.37**** 0.0
1990.11 - 2007.12		**-0.34**** 0.4	**-0.20** 13.0	**-0.04** 65.7	**-0.29**** 3.3	**-0.22**** 3.0			**0.12** 29.0

Notes: Columns represent equations. The first value (in bold) of each cell in Panel A reports the difference between the sum of the relevant coefficients after and before the break date, with this placed against the dates of the second subsample used in the comparison. The value reported is the final one computed for the effect of country k on country h over adjacent subsamples in the recursive general to specific break test procedure (see text). The second value of each cell in Panel A is the bootstrap p-value (expressed as percentage) for the null hypothesis that the coefficients do not change. The first value (in bold) in each cell in Panel B reports the estimated coefficient sum (persistence or spillover) over the indicated subsample, while the second value in each cell in Panel B is the bootstrap p-value (expressed as percentage) for the null hypothesis that the corresponding true value is zero. If an individual break is not significant at 5% in Panel A, the corresponding subsample coefficients are restricted to be equal in Panel B, and are presented under the dates of the earlier subsample. Subsamples are conditional on the estimated VAR coefficient structural break dates of Table 1. ** indicates significance at the 5% level and * significance at the 10% level, both using the bootstrap p-value.

Table 3. Individual Coefficient Break and Spillover/Persistence Results: Euro Area VAR

		A. Coefficient Breaks			B. Inflation Persistence and Spillovers		
Subsample	Explan. Var.	Dependent Variable			Dependent Variable		
		France	Germany	Italy	France	Germany	Italy
1973.03 - 1980.02	**France**				**0.67**** 0.0	**0.30*** 7.8	**0.04** 25.7
1980.03 – 1985.05		**0.21** 20.5	**-0.31**** 2.3	**-0.39*** 7.2		**-0.07** 14.8	
1985.06 - 1990.10		**-0.89**** 0.0	**0.05** 80.6	**-0.23** 49.9	**-0.21**** 5.0		
1990.11 - 2001.03		**0.14** 14.0	**-0.23** 12.2	**0.38*** 9.3			
2001.04 - 2007.12		**-0.74*** 6.6	**-0.27** 12.1	**0.07** 36.5			
1973.03 - 1980.02	**Germany**				**0.13**** 1.9	**0.25**** 2.8	**0.22** 12.0
1980.03 – 1985.05		**-0.43*** 5.7	**-0.09** 34.9	**0.02** 14.6			
1985.06 - 1990.10		**0.36** 30.6	**0.51**** 1.7	**0.12** 48.6		**0.77**** 0.1	
1990.11 - 2001.03		**-0.23** 18.0	**-0.64**** 2.0	**-0.45** 25.3		**0.13** 15.5	
2001.04 - 2007.12		**0.72** 12.6	**0.05** 45.2	**0.10** 73.6			
1973.03 - 1980.02	**Italy**				**0.02** 61.0	**-0.05** 47.8	**0.61**** 0.0
1980.03 – 1985.05		**0.08** 62.7	**0.06*** 11.3	**0.04** 41.8			
1985.06 - 1990.10		**0.05** 83.1	**0.12** 74.3	**-0.24** 37.3			
1990.11 - 2001.03		**-0.08** 32.9	**-0.17** 33.4	**-0.48**** 0.0			**0.13** 45.9
2001.04 - 2007.12		**0.52** 22.4	**0.30** 70.1	**-0.21** 45.5			

Notes: see Table 2.

Table 4. Volatility and Correlation Results: International VAR

	A. Significance of Breaks		B. Subsample Residual Standard Deviations				C. Subsample Contemporaneous Correlations				
Subsample	**Volatility**	**Correl.**	**Canada**	**UK**	**Euro Area**	**US**		**Canada**	**UK**	**Euro Area**	**Zero Correlation**
1973.03 - 1983.07			0.32	0.40	0.16	0.30	**1973.03 - 1999.03**				
							Canada				5.5
1983.08 - 1992.04	0.0	31.2	0.18	0.22	0.14	0.18	**UK**	0.02			80.3
							Euro Area	-0.02	0.10		0.9
1992.05-1999.03	0.9	21.4	0.20	0.18	0.10	0.09	**US**	0.28	0.06	0.34	0.1
1999.04 - 2007.12	0.0	0.0	0.30	0.19	0.14	0.26	**1999.04-2007.12**				
							Canada				0.0
							UK	0.26			0.1
							Euro Area	0.49	0.43		0.0
							US	0.65	0.29	0.59	0.0

Notes: Panel A reports the significance of structural break tests for the diagonal elements of the covariance matrix of the VAR (Volatility) and for the off-diagonal elements of the correlation matrix (Correl), showing bootstrap *p*-values (expressed as percentages) for the test of no change over adjacent Covariance Matrix subsamples identified in Table 1, with the result placed against the dates of the later subsample. The values reported are the final ones computed in the respective general to specific procedures (see text). The corresponding sub-sample residual standard deviations are reported in Panel B and subsample contemporaneous residual correlations in Panel C. The standard deviations and correlations are computed after merging subsamples based on the respective break test results in Panel A (using 5% significance). The final column of Panel C reports the bootstrap *p*-value for a test of the joint hypothesis test that all contemporaneous correlations relating that country are zero. All results are obtained from a VAR in which the restrictions implied by the results of coefficient breaks and persistence/spillover tests (at 5% significance) are imposed.

Table 5. Volatility and Correlation Results: Euro Area VAR

	A. Significance of Breaks		B. Subsample Residual Standard Deviations			C. Subsample Contemporaneous Correlations			
Subsample	**Volatility**	**Correl.**	**France**	**Germany**	**Italy**		**France**	**Germany**	**Zero Correlation**
1973.03 - 1984.04			0.20	0.23	0.35	**France**			97.5
						Germany	0.00		88.6
						Italy	-0.02	0.04	87.5
1984.05 - 1996.11	0.0	1.2	0.14	0.15	0.13	**France**			0.0
						Germany	0.35		0.0
						Italy	0.05	0.20	4.9
1996.12 - 2007.12	0.1	2.6	0.17	0.18	0.10	**France**			0.0
						Germany	0.53		0.0
						Italy	0.32	0.22	0.0

Notes: See Table 4.

Table 6. Robustness of Contemporaneous Correlation Results

	Significance of Correlation Breaks		Subsample Contemporaneous Correlations						
Subsample	Constant VAR Coefficients	All VAR Coefficients Change	Constant VAR Coefficients				All VAR Coefficients Change		
A. International VAR				**Canada**	**UK**	**Euro Area**	**Canada**	**UK**	**Euro Area**
1973.03 - 1983.07			**UK**	-0.08			0.01		
			Euro Area	0.01	0.41		-0.03	0.10	
1983.08 - 1992.04	23.5	26.5	**US**	0.10	0.27	0.35	0.26	0.08	0.31
1992.05 - 1999.03	3.9	23.9	**UK**	0.06					
			Euro Area	-0.03	-0.03				
			US	0.22	-0.02	0.30			
1999.04 - 2007.12	0.0	0.0	**UK**	0.21			0.26		
			Euro Area	0.47	0.44		0.48	0.42	
			US	0.67	0.28	0.57	0.65	0.28	0.58
B. Euro Area VAR				**France**	**Germany**		**France**	**Germany**	
1973.03 - 1984.04			**Germany**	0.16			0.17		
			Italy	0.05	0.09		0.00	0.08	
1984.05 - 1996.11	0.2	0.0	**Germany**	0.55			0.60		
			Italy	0.26	0.23		0.31	0.26	
1996.12 - 2007.12	7.3	7.9							

Notes: Columns headed 'Significance of Correlation Breaks' show bootstrap *p*-values (expressed as percentages) for the test of no change in correlations over adjacent Covariance Matrix subsamples identified in Table 1, with the result placed against the dates of the later subsample. The values reported are the final ones computed in the general to specific procedure (see text). Correlations are computed after merging subsamples based on the break test results (using 5% significance). The results reported assume all VAR coefficients are constant over the entire sample or all coefficients are allowed to change at the VAR coefficient break dates identified for the respective systems in Table 1, with no further restrictions imposed on these coefficients. All volatility breaks are significant and are taken into account.