

STRUCTURAL BREAKS IN THE INTERNATIONAL DYNAMICS OF INFLATION

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Abstract—This paper proposes an iterative procedure to discriminate between structural breaks in the coefficients and the disturbance covariance matrix of a system of equations, with recursive procedures then identifying individual coefficient shifts and separating volatility from correlation breaks. Structural breaks in short-term cross-country inflation relations are then examined for major G-7 economies and within the euro area. There is evidence that the euro area leads inflation in North America, while changing short-term interactions apply within the euro area. Covariability generally increases from the late 1990s, while euro-area countries move from essentially idiosyncratic contemporaneous variation to comovement in the 1980s.

I. Introduction

THE principal aim of monetary policy is to keep inflation at low and stable levels. Many central banks around the globe have formally adopted such inflation targeting as the main goal of their monetary policy since the early 1990s. Although monetary policy decisions are (generally) taken individually by central banks in each country, recent research has uncovered evidence of a strong international dimension to the inflation experience of developed countries (Ciccarelli & Mojon 2010; Monacelli & Sala 2009; Mumtaz & Surico, 2012; Neely & Rapach, 2011). 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, of course, one important contemporary instance where monetary policy is taken on a cross-country basis, which is that of the euro area, for which the common monetary policy regime came into operation in January 1999. To

our knowledge, it has not yet been studied whether this has changed the links between inflation in the euro-area countries.¹ More generally, although previous studies (including Ciccarelli & Mojon, 2010; Mumtaz & Surico, 2012; and Neely & Rapach, 2011) allow some temporal variation in cross-country relations, a systematic examination of the dates and nature of structural breaks in international inflation dynamics has not yet been conducted.

This paper fills this gap by examining changes in cross-country short-run inflation dynamics among major international economies over the period 1973 to 2007. Our analysis employs a conventional vector autoregressive (VAR) specification but separates the roles of dynamic coefficients, variances, and correlations in that we apply structural break tests to examine the presence of changes in each of these three components separately. Although our methodology builds on the multivariate procedure of Qu and Perron (2007), it does not restrict the structural breaks in all VAR parameters to occur at common dates. Indeed, an important methodological contribution of our paper is that we develop a new iterative procedure that separates breaks in the dynamic coefficients from those in the covariance matrix, providing a flexible algorithm that can work well even with moderate sample sizes. We further propose an algorithm for decomposing covariance matrix breaks into volatility and correlation changes. Taken together, these procedures enable us to pinpoint the nature of changes in the dynamic variability and covariability of international inflation.

In contrast to the relative lack of previous research on changes in international inflation linkages, an important focus of univariate analyses has been the nature and timing of structural breaks, with changes in the level of inflation or its persistence often linked with monetary policy changes (see, for example, Altissimo, et al., 2006; Benati, forthcoming; Cecchetti & Debelle, 2006; Halunga, Osborn, & Sensier, 2009; O'Reilly & Whelan, 2005). The clustering of break dates across countries documented by Bataa et al. (forthcoming), Corvoisier and Mojon (2005), Levin and Piger (2004), and others suggests that a multivariate analysis may be worthwhile to exploit the cross-sectional dimension for obtaining more precise inference on these structural breaks (see Bai, Lumsdaine & Stock, 1998). In addition to changes in the marginal characteristics, such as the persistence of inflation, a multivariate analysis allows us to discover if changes have occurred in cross-country inflation linkages.

We examine links between inflation in Canada, the euro area, the United Kingdom, and the United States, on the

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¹ Although Wang and Wen (2007) conduct a robustness analysis comparing euro-area countries with other countries in their sample, this constant-parameter analysis starts in the 1970s.

one hand, and between France, Germany, and Italy on the other.² 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. First, when the euro area is treated as a single economy, short-run dynamic interactions are largely absent in international inflation. The most notable exception is that the euro area appears to lead inflation in the United States and Canada (but not the United Kingdom). Second, contemporaneous cross-correlations are generally relatively small until 1983, such that contemporaneous conditional inflation movements during the first part of the sample period can be characterized as largely country-centric, despite the conventional wisdom that the large oil price shocks of the 1970s substantially affected general price levels globally. Euro-area countries undergo substantial correlation increases around 1983, but no further change is detected that can be associated with the introduction of the common currency. However, inter-continental correlations jump to levels between 0.3 and 0.6 in 1999; indeed, the post-1999 U.S./euro area correlation is even larger than that between France and Germany. Third, findings from previous univariate analyses that inflation persistence falls over this period while volatility is unstable are confirmed in our multivariate setting.

The predominant econometric method in recent analyses of inflation linkages is based on dynamic factor models, including the contributions of Ciccarelli and Mojon (2010), Mumtaz and Surico (2012), and Neely and Rapach (2011). A factor-based approach is attractive in that it allows for disentangling common and country-specific shocks to inflation. It then becomes possible to analyze, for example, how inflation in individual countries reacts to a common shock and how country-specific shocks affect inflation internationally. The extraction of factors, however, requires quite strong assumptions about the dynamics and covariances linking the inflation series; typically the underlying parameters are assumed constant, although the analysis of Mumtaz and Surico (2012) permits them to evolve as random walk processes. While our VAR-based methodology is not designed to distinguish between common and idiosyncratic shocks, it is better suited for examining dynamic interactions between inflation across countries. Specifically, the VAR framework makes it possible to disentangle breaks in level, persistence, dynamic covariability, and contemporaneous correlations of inflation, while it is also able to account for the presence of structural breaks in volatility.

The paper is organized as follows. Section II discusses our data and some of its cross-country features. The methodology we employ is outlined in section III, with further details concerning the tests for identifying breaks in volati-

lities and contemporaneous correlations provided in the online appendix. Our principal empirical results concerning changes in international inflation linkages are presented in section IV. Finally, section V concludes.

II. Data and Preliminary Analysis

We analyze consumer price inflation in individual G-7 countries and the euro area, with inflation computed as one hundred times the monthly difference of the log consumer price index (CPI) over the period April 1973 to December 2007. A detailed description of the data set is provided in Bataa et al. (forthcoming), where a univariate analysis is undertaken. A salient feature of that analysis is the substantial communality in the timing of breaks in persistence (during the late 1980s and early 1990s) and in volatility (in the early 1980s and in the mid- to late 1990s).³ This suggests that a multivariate treatment may deliver gains in the precision of break date estimates (see Bai et al., 1998), in addition to revealing changes in international dynamic comovements. It is useful to note that while previous studies on international inflation and structural changes in inflation generally use quarterly data, our analysis is conducted at the monthly frequency in order to more accurately separate contemporaneous and dynamic effects. The monthly frequency is also relevant because it reflects the typical frequency of monetary policy decisions in the countries under study.

The analysis of this paper is based on series cleaned of seasonality and outliers through an iterative (univariate) procedure that allows for time-varying unconditional mean, seasonality, persistence, and (conditional) volatility (see Bataa et al., forthcoming, for details).

As noted in Section I, our analysis considers two systems—one consisting of Canada, the euro area, the United Kingdom, and the United States and the other of the individual euro-area countries of France, Germany, and Italy. The former captures short-run dynamic interactions between the G7 economies, with the euro area treated as a single entity in order to abstract (as far as possible) from changes in inflation linkages between G7 countries that are now members of the European monetary union.⁴ These intra-euro area changes are the focus of the system for France, Germany, and Italy. Japan is excluded because, as documented in previous research, it was largely disconnected from other G7 economies during our sample period.

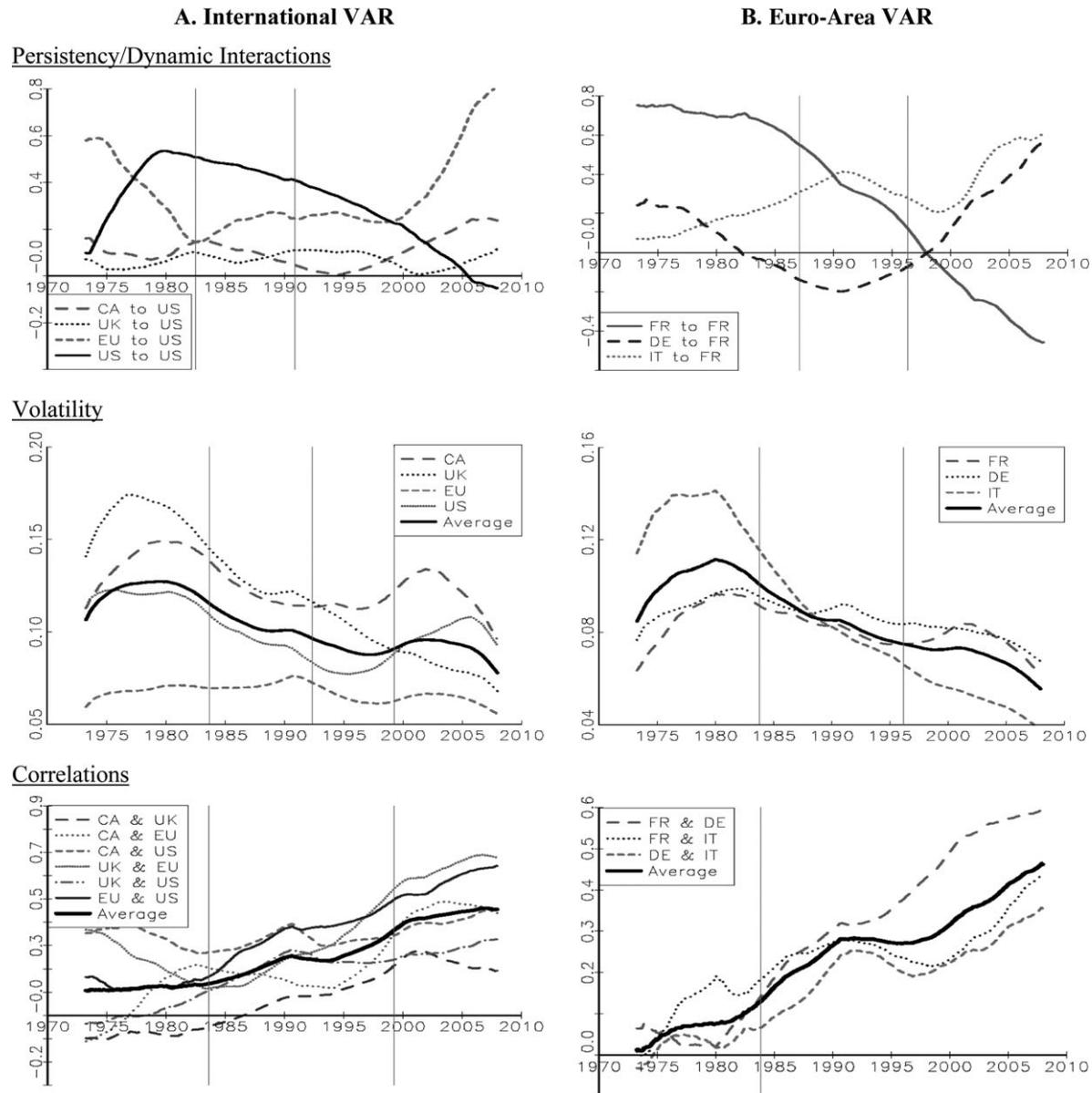
For both sets of countries, we analyze their inflation dynamics and possible changes therein using VAR models. As a prelude to this formal structural breaks analysis, figure 1 provides a descriptive analysis of possible time variation.

³ For most countries (with the exception of Italy), inflation persistence falls to 0 in the latter part of the sample. Inflation volatility decreases substantially in the early 1980s but increases again for the United States, Canada, and France around 1999.

⁴ While recognizing that it is a misrepresentation of the nature of economic and monetary union, nevertheless for convenience we later often refer to the euro area as a “country.”

² 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.

FIGURE 1.—TIME-VARYING VAR MODELS, APRIL 1973–DECEMBER 2007



The international VAR refers to Canada (CA), the euro area (EU), United Kingdom, and United States with $p = 1$ lag, while the euro-Area VAR includes France (FR), Germany (DE), and Italy (IT) with $p = 2$. In the latter case, the persistence/dynamic interaction results are the sum of the relevant VAR coefficients in the equation for France. The results are obtained by estimating VAR models with weights centered at each month τ such that an observation at time $\tau + k$ has weight $\lambda^{|k|}$, for $k = \dots, -2, -1, 0, 1, 2, \dots$ and $\lambda = 0.99$, with the results then plotted for each τ .

These time-varying specifications are obtained by estimating VAR models using the full sample but with weighted least squares, such that the observation at time $\tau + k$ is given weight $\lambda^{|k|}$, for $k = \dots, -2, -1, 0, 1, 2, \dots$, with $0 < \lambda < 1$, resulting in VAR estimates centered at $t = \tau$. Repeating this for $\tau = 1, 2, \dots, T$ (where T denotes the sample size) yields a sequence of VARs with smoothly time-varying parameter estimates. Although by design, this weighting will not reproduce any abrupt change, it nevertheless provides information about the presence, or otherwise, of temporal variation. We employ $\lambda = 0.99$ for figure 1 to ensure that each estimate reflects information in a sample of reasonable effective size. The lag orders employed are selected

using the BIC criterion, leading to one and two lags for the international and euro-area VARs, respectively. The vertical lines in the figure indicate break dates found to be significant for the specific component (coefficients, volatility, or correlation) in the analysis of section IV.

Although, to conserve space, not all VAR coefficients are shown, those presented in the top panel of figure 1 (relating to the equations for the United States and France) are indicative of the general patterns found. The coefficient values shown, here and throughout, are the sum of the respective coefficients for each country when the VAR order $p > 1$. Thus, in the equation for France, the coefficients on inflation for each of Germany, France, and Italy are summed

over the two lags in the euro-area VAR. We observe that persistence, as measured by the own-coefficients, declines substantially over time (after an initial increase in the U.S. case). Dynamic interactions, represented by the (sum of) coefficients for other countries, are small and do not exhibit clear temporal patterns for most combinations of countries (as for the United Kingdom and Canada in the equation for the United States). Exceptions to this general finding do occur. For example, the coefficient of lagged euro-area inflation in the equation for the United States declines substantially from 0.6 to 0.2 in the first decade of the sample period, then remains stable at this level until the year 2000, and increases sharply to about 0.8 at the end of the sample period. A similar pattern of change is observed for Germany in the equation for France.

Broadly speaking, after an initial increase during the second half of the 1970s, (conditional) volatility also declines over time in figure 1, with this feature especially clear for the United Kingdom and Italy. For some countries, in particular the United States, it seems that volatility increases (perhaps temporarily) again after 2000. The clearest patterns occur in the contemporaneous disturbance correlations, which increase over time for all country pairs except Canada and the United States. For the euro-area economies, this increase appears to start around 1980 and might be attributed to progress toward monetary integration. However, similarly marked increases also occur in the international VAR, with these being especially notable in the figure from the second half of the 1990s.

Although informative that temporal change occurs and of its broad characteristics, such a descriptive analysis does not provide firm statistical evidence for the presence and dates of structural breaks in various elements of the VAR systems. The next section details our testing methodology, with empirical results presented in section IV. These results back up figure 1 by indicating that temporal change is a feature of international inflation, with substantial increases in the disturbance correlations being the most important feature uncovered.

III. Testing Methodology

Our multivariate analysis adopts a systematic approach in order to identify structural breaks in short-run dynamics, volatility, and correlations of inflation. As already noted, the framework starts from a conventional VAR system for n countries,

$$\mathbf{y}_t = \boldsymbol{\delta} + \sum_{i=1}^p \mathbf{A}_i \mathbf{y}_{t-i} + \mathbf{u}_t \quad (1)$$

where $\mathbf{y}_t = [y_{1,t}, \dots, y_{n,t}]'$ is the $n \times 1$ vector of inflation in month t , \mathbf{A}_i ($i = 1, \dots, p$) are $n \times n$ coefficient matrices, and $\boldsymbol{\delta}$ is an $n \times 1$ vector of intercepts. The error term \mathbf{u}_t in equation (1) has mean zero and covariance matrix $E(\mathbf{u}_t \mathbf{u}_t') = \boldsymbol{\Sigma}$ and is temporally uncorrelated. Further, defining \mathbf{D} to be

the diagonal matrix containing the standard deviations of \mathbf{u}_t and \mathbf{P} to be the corresponding correlation matrix, then by definition, of $\boldsymbol{\Sigma} = \mathbf{D} \mathbf{P} \mathbf{D}$. Our methodology dates structural breaks in each of the three components of equation (1): first, the VAR coefficients \mathbf{A}_i ($i = 1, \dots, p$) and $\boldsymbol{\delta}$, which together capture mean effects as well as dynamics;⁵ second, (conditional) volatility measured by \mathbf{D} ; and finally, contemporaneous (conditional) correlations in \mathbf{P} . In addition to dating any breaks found to occur, we also examine the statistical significance of international relations by conducting inference on \mathbf{A}_i and \mathbf{P} .

Our approach builds on the recent methodology of Qu and Perron (2007), developed to test for mean and covariance breaks in a VAR system.⁶ Their methodology provides tools to deal with three scenarios: (a) breaks occurring only in the VAR coefficients, $\boldsymbol{\delta}$ and \mathbf{A}_i ; (b) breaks occurring only in the covariance matrix $\boldsymbol{\Sigma}$; or (c) breaks occurring in both the VAR coefficients and the covariance matrix. In the last case, the two types of breaks are restricted to occur simultaneously, but this may be inappropriate in practice. Indeed, although there is widespread evidence suggesting the possibility of breaks in both components for international inflation, it does not seem that these occurred at the same time (and, consequently, the numbers of breaks need not even be the same). Previous literature concerning the univariate properties of inflation suggests that volatility declines occurred in the early 1980s (see Sensier & van Dijk, 2004) while structural changes in dynamics happened mostly in the early 1990s (see Cogley & Sargent, 2002, 2005). A further motivation for our extension of the Qu and Perron (2007) methodology is the potential problem of misspecified breaks in dynamics affecting inferences on structural stability in volatility and vice versa, as shown in a univariate context by Pitarakis (2004; see also our univariate analysis of the same inflation series in Bataa et al., forthcoming).

For the above reasons, we develop a new iterative procedure to test for (possibly distinct) 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 section IIIA. Section IIIB then describes our iterative coefficient and covariance procedure, with the small sample properties of this procedure examined through a Monte Carlo experiment. Section IIIC provides details on hypothesis tests concerning own- and cross-country inflation dynamics, and section IIID discusses the testing procedure for disentangling breaks in volatilities and correlations. Further details of the latter procedure can be found in the online appendix.

⁵ We also conduct an analysis that separates mean effects from dynamics by prior removal of mean breaks; see section IVB.

⁶ This procedure improves on the commonly used single-equation approach of Bai and Perron (1998, 2003) and the approach of Bai et al. (1998) for breaks in VAR systems. The main advantages of the Qu and Perron (2007) methodology are 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 heteroskedasticity autocorrelation, and it allows multiple breaks.

A. Tests for Coefficient and Covariance Breaks

Prior to testing, the order p of the VAR in equation (1) is selected using the BIC criterion over the entire sample period. Then, using the procedure of Qu and Perron (2007), the stability of the VAR coefficients in equation (1) is checked against the possibility of $\ell \leq M$ breaks, where ℓ is unknown and the maximum number of breaks M is prespecified. This is implemented as a test of the null hypothesis $H_0 : \delta_j = \delta; \mathbf{A}_{i,j} = \mathbf{A}_i (j = 1, \dots, \ell + 1; i = 1, \dots, p)$ in

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

for $t = T_{j-1} + 1, \dots, T_j, j = 1, \dots, \ell + 1$, where T_j are the break dates marking the $\ell + 1$ subsamples, with $T_0 = 0$ and $T_{\ell+1} = T$, and where \mathbf{u}_t can be heteroskedastic.

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

$$WD\max F_T(M) = \max_{1 \leq \ell \leq M} a_\ell \left[\sup_{(\lambda_1, \dots, \lambda_\ell \in \Lambda_\varepsilon)} F_T(\ell, q, \varepsilon) \right], \quad (3)$$

where $\lambda_j (j = 1, \dots, \ell)$ indicate possible break dates as fractions of the sample size, with $0 < \lambda_1 < \dots < \lambda_\ell < 1$ and $T_j = [T\lambda_j]$, and Λ_ε denotes the set of all permissible sample partitions satisfying the requirement that at least a fraction ε of the sample is contained in each segment for some $0 < \varepsilon < 1$. The parameter $a_\ell = c(\alpha, 1)/c(\alpha, \ell)$, with $c(\alpha, \ell)$ the asymptotic critical value (at a significance level of 100 $\alpha\%$) of the supremum statistic $\sup_{(\lambda_1, \dots, \lambda_\ell \in \Lambda_\varepsilon)} F_T(\ell, q, \varepsilon)$

against a specific number of ℓ breaks. For a total of q VAR coefficients (including intercepts) in equation (1), all of which are allowed to change,

$$F_T(\ell, q, \varepsilon) = (T - (\ell + 1)q) \hat{\boldsymbol{\beta}}' \mathbf{R}' [\mathbf{R} \hat{\mathbf{V}}(\hat{\boldsymbol{\beta}}) \mathbf{R}']^{-1} \mathbf{R} \hat{\boldsymbol{\beta}} \quad (4)$$

is a Wald-type test statistic for structural change at ℓ fixed dates, $\hat{\boldsymbol{\beta}}$ is the stacked vector of estimated coefficients given the ℓ breaks with estimated robust covariance matrix $\hat{\mathbf{V}}(\hat{\boldsymbol{\beta}})$, and \mathbf{R} is the nonstochastic matrix such that $(\mathbf{R} \hat{\boldsymbol{\beta}})' = (\hat{\boldsymbol{\beta}}_1' - \hat{\boldsymbol{\beta}}_1', \dots, \hat{\boldsymbol{\beta}}_{\ell+1}' - \hat{\boldsymbol{\beta}}_\ell')$, where $\boldsymbol{\beta}_j$ is the vector of coefficients in the j th segment.

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

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

where $\Lambda_{j,\varepsilon} = \{\tau; \hat{T}_{j-1} + (\hat{T}_j - \hat{T}_{j-1})\varepsilon \leq \tau \leq \hat{T}_j - (\hat{T}_j - \hat{T}_{j-1})\varepsilon\}$, and F_T is defined as in equation (4). The statistic in equation (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 $(\hat{T}_1, \hat{T}_2, \dots, \hat{T}_l)$ obtained from estimating the model with l breaks. The test is applied sequentially for $l = 1, 2, \dots$ until it fails to reject the null hypothesis of no additional break.

Having determined the number of structural breaks using equation (5), we estimate the break dates and VAR coefficients 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 $\boldsymbol{\Sigma}$ proceeds along similar lines as just described. The null hypothesis of no covariance breaks, that is, $H_0 : \boldsymbol{\Sigma}_j = \boldsymbol{\Sigma} (j = 1, 2, \dots, m+1)$ for an unknown $m \leq M$ number of breaks, is tested using a double maximum likelihood ratio-type test statistic. In particular, the $SupF$ statistic in equation (3) is replaced by the $SupLR$ statistic defined as

$$\sup LR_T(m, q, \varepsilon) = \sup_{(\lambda_1, \dots, \lambda_m \in \Lambda_\varepsilon)} 2 \ln \left(\frac{\hat{L}_T(T_1, \dots, T_r)}{\tilde{L}_T} \right), \quad (6)$$

where

$$\ln \hat{L}_T(T_1, \dots, T_m) = -\frac{T}{2} (\ln 2\pi + 1) - \sum_{j=1}^{m+1} \frac{T_j - T_{j-1}}{2} \ln |\hat{\boldsymbol{\Sigma}}_j|$$

$$\boldsymbol{\Sigma}_j = \frac{1}{T_j - T_{j-1}} \sum_{t=T_{j-1}+1}^{T_j} \hat{\mathbf{u}}_t \hat{\mathbf{u}}_t'$$

$\hat{\mathbf{u}}_t (t = 1, \dots, T)$ are the VAR residual series obtained from equation (2) and \tilde{L}_T is the (quasi) likelihood computed under the null hypothesis of no breaks.

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 equation (5) replaced by

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

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 either type (coefficient or covariance matrix) of analysis, 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 on 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.

B. Disentangling Coefficient and Covariance Breaks

We employ the tests outlined in the previous section within an iterative procedure to disentangle breaks in the VAR coefficients and in the conditional covariance matrix, analogous to that in Bataa et al. (forthcoming) in a single-equation context. This allows the numbers of breaks in δ and A_i , on the one hand, and Σ , on the other, to be different and for breaks in these components to occur at different dates. After outlining our iterative methodology, we present a Monte Carlo analysis of its performance.

Iterative methodology. The procedure starts from VAR coefficient breaks identified using a heteroskedasticity robust covariance matrix in equation (4). However, the Monte Carlo analysis of Pitarakis (2004), in a univariate context, shows that such robust tests can be very badly oversized for the detection of breaks in the coefficients of a dynamic process. Therefore, within each subsequent iteration, our procedure first tests for breaks in the covariance matrix conditioning on the estimated break dates for δ and A_i and then, retests for breaks in the VAR coefficients conditioning on the estimated break dates for Σ . In the latter case, a feasible generalized least squares (GLS) procedure is employed to exploit the covariance breaks information. The iterations continue until convergence. Although initialized from consistent estimates of the coefficient break dates, it is anticipated that the iterative procedure, and particularly exploitation of the covariance break information, will result in improved finite sample performance for the detection of coefficient breaks.

A further finite sample refinement implemented in our procedure is to use bootstrap inference to verify the existence of each (coefficient or covariance) break identified through the asymptotic Qu and Perron (2007) procedure. For this purpose, the possible break dates are treated as known and given by the asymptotic procedure. Building this into the iterative methodology that identifies (separate) breaks in (δ, A_i) and Σ provides some assurance against the detection of spurious breaks.

In more detail, the iterative algorithm consists of the following steps:

0. (a) Initialize the coefficient breaks by using the asymptotic tests of equations (3) and (5), employing heteroskedasticity (HC) robust inference. If ℓ breaks in the VAR coefficients are detected at dates $\hat{T}_1^{(B)}, \dots, \hat{T}_\ell^{(B)}$, use observations $t = \hat{T}_{k-1}^{(B)} + 1, \dots, \hat{T}_k^{(B)}$ to obtain $\hat{\beta}_k$ and corresponding residual series \hat{u}_t , for each regime $k = 1, \dots, \ell + 1$.
- (b) Verify the significance of each coefficient break $k = 1, \dots, \ell$ through a finite sample test of the null $\beta_k = \beta_{k+1}$. Inference is conducted conditional on all other $\ell - 1$ estimated breaks, employing the F -statistic of equation (4) with HC robust covariance matrix and $(R\hat{\beta})' = (0', \dots, \hat{\beta}'_{k+1} - \hat{\beta}'_k, \dots, 0')$. The computed value F is compared to the corresponding empirical distribution obtained from a bootstrap data-generating process (DGP) that employs estimated VAR parameters but restricted through $\beta_k = \beta_{k+1}$, and a wild bootstrap process for u_t in equation (2).⁷
- (c) If not all coefficient breaks $k = 1, \dots, \ell$ are individually significant, reduce the number of coefficient breaks to $\ell - 1$. Then estimate new break dates with corresponding VAR coefficients and residuals and return to step 0b. This is repeated until all coefficient breaks are individually significant.
1. (a) Based on the residuals \hat{u}_t obtained from step 0 (or 2 below), apply the asymptotic tests of equations (6) and (7) for breaks in the covariance matrix Σ . If m breaks are detected at dates $\hat{T}_1^{(C)}, \dots, \hat{T}_m^{(C)}$, obtain $\hat{\Sigma}_j$ for each regime $j = 1, \dots, m + 1$, estimated from observations $t = \hat{T}_{j-1}^{(C)} + 1, \dots, \hat{T}_j^{(C)}$.
- (b) For each covariance break $j = 1, \dots, m$ identified in 1a, and conditioning on all other $m - 1$ breaks, compute the usual quasi-likelihood ratio test statistic, LR , for the null $\Sigma_j = \Sigma_{j+1}$. For inference on break j , the residual vectors \hat{u}_t , for $t = \hat{T}_{j-1}^{(C)} + 1, \dots, \hat{T}_j^{(C)}, \dots, \hat{T}_{j+1}^{(C)}$ are randomly i.i.d. resampled, with a wild bootstrap employed in other regimes to create the bootstrap residuals \hat{u}_t^* .⁸ Then $\hat{u}_t^*(t)$, together with the $(\ell + 1)$ sets of VAR coefficient estimates found in step 0 (or 2), form the bootstrap DGP that is used to obtain the empirical null distribution, and hence the empirical p -value, for the LR statistic.⁹
- (c) If not all covariance matrix breaks are individually significant, reduce the number of coefficient breaks to $m - 1$. Then estimate new covariance break dates and corresponding covariance matrices. Return to step 1b until all covariance breaks are individually significant.
2. Reestimate the VAR coefficient breaks using a feasible generalized least squares (GLS) approach, which is achieved by premultiplying all variables entering the VAR for $t = \hat{T}_{j-1}^{(C)} + 1, \dots, \hat{T}_j^{(C)}$ by $\hat{\Sigma}_j^{-1/2}$, $j = 1, \dots, m + 1$,

⁷ The wild bootstrap sets $\hat{u}_t^* = \gamma \hat{u}_t$ where γ is randomly chosen as $+1$ or -1 with equal probabilities. The wild bootstrap allows heteroskedasticity of unknown form (Gonçalves & Kilian, 2004), and hence is preferable to an i.i.d. bootstrap in this case.

⁸ The i.i.d. bootstrap within the regimes under test enforces the null hypothesis of unchanged variance, whereas the wild bootstrap for the remaining observations allows variances to differ across the other regimes.

⁹ The VAR coefficients and residuals are reestimated for the bootstrap DGP, but for computational feasibility, the coefficient break dates are assumed known at $\hat{T}_1^{(B)}, \dots, \hat{T}_\ell^{(B)}$.

TABLE 1.—MONTE CARLO ANALYSIS OF ITERATIVE BREAK TESTING PROCEDURE

	Proportion Correct Number of Breaks		Average Number of Breaks Detected		Break Fraction Estimation				Nonconvergence (%)
					Coefficient		Covariance		
	Coefficient	Covariance	Coefficient	Covariance	Mean error	MAE	Mean error	MAE	
Case I. No Breaks									
No iteration, asymptotic	0.91	0.95	0.09	0.05	N.A.	N.A.	N.A.	N.A.	N.A.
With iteration, asymptotic	0.91	0.95	0.09	0.06	N.A.	N.A.	N.A.	N.A.	0.1
With iteration, bootstrap	0.93	0.96	0.08	0.05	N.A.	N.A.	N.A.	N.A.	0.0
Case II. One Coefficient Break Only									
No iteration, asymptotic	0.96	0.95	1.03	0.06	0.002	0.015	N.A.	N.A.	N.A.
With iteration, asymptotic	0.96	0.95	1.03	0.06	0.002	0.015	N.A.	N.A.	0.4
With iteration, bootstrap	0.97	0.94	1.03	0.06	0.003	0.015	N.A.	N.A.	0.2
Case III. One Covariance Break Only									
No iteration, asymptotic	0.97	0.96	0.04	1.04	N.A.	N.A.	-0.001	0.003	N.A.
With iteration, asymptotic	0.91	0.96	0.10	1.04	N.A.	N.A.	-0.001	0.003	0.5
With iteration, bootstrap	0.94	0.96	0.07	1.04	N.A.	N.A.	-0.001	0.003	0.2
Case IV. One Coefficient and One Covariance Break									
No iteration, asymptotic	0.03	0.96	1.99	1.05	0.299	0.299	-0.002	0.003	N.A.
With iteration, asymptotic	0.95	0.96	1.05	1.05	-0.001	0.014	-0.001	0.003	0.4
With iteration, bootstrap	0.98	0.97	1.02	1.04	0.000	0.013	-0.001	0.004	0.4

Results are presented for the coefficient and covariance test procedure of section IIIB, both without and with the bootstrap confirmation of breaks, denoted asymptotic and bootstrap, respectively; the former case is considered without and with iteration. The DGP is a three-equation VAR(1) with $T = 420$ observations and parameters as given in equations (8) and (9). The coefficient and covariance breaks occur at $0.25T$ and $0.55T$, respectively. The mean error and mean absolute error (MAE) for break fraction estimation are computed over cases that detect the correct number of breaks in the component indicated. All results are based on 2,000 replications. All tests allow a maximum of three breaks in each component and apply 20% trimming. When iteration is applied, the final column shows the percentage of replications that do not converge in eight iterations

where $\hat{\Sigma}_j^{-1/2}$ is the inverse square root of the corresponding estimated covariance matrix. Follow the coefficient test procedure as in step 0, but now apply the multiple breaks test procedure to the VAR coefficient vector β using equation (4) with constant disturbance covariance matrix in the GLS transformed system.

- Iterate between steps 1 and 2 until the numbers and dates of breaks in the VAR coefficients and the conditional covariance matrix do not change.

By omitting parts b and c of these steps, an iterative procedure based on asymptotic tests is obtained, while the application of step 0a followed by 1a, with no iteration results in a two-step asymptotic test procedure for distinct coefficient and covariance breaks.

Monte Carlo results. A Monte Carlo experiment is conducted to shed light on the performance of the procedures just outlined. The data-generating process (DGP) is a zero-mean trivariate VAR(1) without and with breaks in both the coefficient and covariance matrices. The (regime-dependent) coefficient matrices are selected to replicate the broad features of short-run international inflation dynamics, with

$$\mathbf{A}_{1,1} = \begin{bmatrix} 0.7 & 0.1 & 0.5 \\ 0.2 & 0.3 & 0.4 \\ 0 & 0 & 0.5 \end{bmatrix}, \quad \mathbf{A}_{1,2} = \begin{bmatrix} 0.1 & 0.1 & 0.5 \\ 0.2 & 0.1 & 0.4 \\ 0 & 0 & 0.1 \end{bmatrix} \quad (8)$$

capturing the reduction in persistence (as represented by the diagonal elements) that we find to be a feature of the later empirical analysis. The covariance matrices shed light on the detection of changes in correlations (matrix \mathbf{P}) and stan-

dard deviations (vector \mathbf{D}^* , which contains the diagonal elements of \mathbf{D}). The regime-specific values employed are

$$\mathbf{P}_1 = \begin{bmatrix} 1 & 0 & 0 \\ 0 & 1 & 0 \\ 0 & 0 & 1 \end{bmatrix}, \quad \mathbf{P}_2 = \begin{bmatrix} 1 & 0.5 & 0.2 \\ 0.5 & 1 & 0.2 \\ 0.2 & 0.2 & 1 \end{bmatrix}, \quad (9)$$

$$\mathbf{D}_1^* = [0.2, 0.2, 0.35], \quad \mathbf{D}_2^* = [0.15, 0.15, 0.15].$$

This covariance break introduces both cross-equation disturbance correlations (representing global inflation influences) and a general decline in volatility, features observed in the later empirical analysis.¹⁰ All disturbance processes are normally distributed with $T = 420$ observations.¹¹ In common with the empirical analysis, a nominal level of significance of 5% is employed for all tests. The maximum number of breaks considered in the break testing procedure is $M = 3$, with $\varepsilon = 0.20$, which apply for each (coefficient or covariance) component. Results are based on 2,000 replications.

Table 1 summarizes the results, with our proposed iterative asymptotic and bootstrap procedures compared to the more conventional two-step asymptotic approach. All procedures yield good size for the covariance break test results in all cases, with iteration and the bootstrap having little effect for this component. For case I, with no breaks (the coefficient matrix $\mathbf{A}_{1,1}$ and covariance matrix Σ_1 apply throughout), and case II, with one coefficient break only (namely, a change from $\mathbf{A}_{1,1}$ to $\mathbf{A}_{1,2}$ occurring at $T^{(B)} = 0.25T$ with con-

¹⁰ Indeed, the parameter values employed in equations (8) and (9) for the coefficients and covariance matrix are specified particularly in the light of breaks found for the euro-area VAR in section IV.

¹¹ These and all subsequent results are obtained using Gauss. Zero initial starting values are employed for all variables in these simulations, with the first 100 observations discarded to remove their effect.

stant covariance matrix), the three methods also deliver very similar results for the coefficient breaks test, though with some indication of better size using the bootstrap for case I (the proportion correct should be 0.95 for the nominal 5% significance level).

In the presence of a covariance break and no coefficient break (case III, with the correlation and volatility break implied by a change from \mathbf{P}_1 to \mathbf{P}_2 and \mathbf{D}_1 to \mathbf{D}_2 in equation (9) occurring at $T^{(C)} = 0.55T$), the coefficient break test is undersized when using the asymptotic procedure without iteration; this is in line with the simulation results of Pitarakis (2004) concerning the heteroskedasticity robust testing approach in a univariate context. However, using iterations in the asymptotic procedure results in oversizing, with the rejection frequency increasing to 9% at the nominal 5% significance level. The best size is delivered by the iterative bootstrap procedure; again, this in line with Pitarakis (2004), who proposes using a bootstrap in the GLS transformed model (but without iteration). The bootstrap may also help reduce nonconvergence of the iterative procedure, although this applies here only 0.5% of the time.

The most revealing case is the one with coefficient and covariance breaks: case IV with breaks occurring at $T^{(B)} = 0.25T$ and $T^{(C)} = 0.55T$, with coefficients as in equations (8) and (9). The asymptotic procedure without iteration performs very poorly here in the sense that it almost always identifies two coefficient breaks when only one actually occurs; the correct number of coefficient breaks is detected in only 3% of the replications. Iteration clearly improves on this, but the best performance is achieved with the bootstrap procedure, which correctly identifies one coefficient break for 98% of the replications.

When the correct number of breaks is identified, there is little difference between the procedures in terms of their accuracy of estimation of the true covariance break date (as a fraction of the sample size). All procedures give very accurate estimates, with the average absolute error being approximately 0.003, or an average error of a little more than one month with a sample of 35 years (420 monthly observations).

Coefficient break dates are also quite well estimated by all procedures, with the exception of the asymptotic procedure without iteration for the coefficient and covariance break case. Here the single identified coefficient break is wrongly estimated to occur at the true covariance break date. Otherwise an average absolute break fraction error of 0.015 is equivalent to about six months in a sample of 35 years of monthly data. With iteration, the coefficient break fraction estimation is as accurate in the presence of a covariance break as when no such break occurs (case IV compared with case II).

Overall, table 1 verifies that the iterative procedure with bootstrap inference copes well in delivering tests of appropriate size when one component changes and the other does not, hence guarding against the problems that Pitarakis (2004) identified. The benefit of bootstrap inference is seen particularly in relation to the detection of coefficient breaks,

although the asymptotic iterative procedure is also reliable for the dating of breaks.

C. Individual Coefficient Tests

The coefficient breaks resulting from the analysis outlined in section IIIB apply to the VAR system as a whole. However, it is also of interest to identify whether these relate to intercept changes, persistence changes in individual inflation series, or changes in cross-country dynamics. To shed light on the source of change, we employ a general-to-specific approach to test the equality of individual VAR coefficients across subsamples. This is based on the conventional test statistic of equation (4), but with the restriction matrix \mathbf{R} defined such that either $(\mathbf{R}\hat{\boldsymbol{\beta}})' = (0, \dots, \hat{\delta}_h^{(j)} - \hat{\delta}_h^{(j-1)}, \dots, 0)$ where $\hat{\delta}_h^{(j)}$ is the h th element of δ , or

$$(\mathbf{R}\hat{\boldsymbol{\beta}})' = (0, \dots, \hat{a}_{(h,k),1}^{(j)} - \hat{a}_{(h,k),1}^{(j-1)}, \hat{a}_{(h,k),2}^{(j)} - \hat{a}_{(h,k),2}^{(j-1)}, \dots, \hat{a}_{(h,k),p}^{(j)} - \hat{a}_{(h,k),p}^{(j-1)}, \dots, 0),$$

where $\hat{a}_{(h,k),i}^{(j)}$ is the (h, k) th element of \mathbf{A}_i in the j th regime, and the degrees of freedom are correspondingly adjusted. Note that this test applies to either an individual intercept or the set of VAR coefficients for the dynamic impact of inflation in country k on that of country h at all lags $i = 1, \dots, p$, with regime j compared to regime $j - 1$. For this purpose, the analysis is conditional on the estimated VAR break dates obtained from the system, with all coefficients not under study initially allowed to change at these dates (as in Doyle & Faust, 2005). However, since testing only adjacent regimes may lead to relatively low power, these tests are employed recursively. Specifically, we initially compute F -tests for all intercept and coefficient breaks $j = 1, \dots, \ell$. The intercepts for each equation, δ_h for $h = 1, \dots, n$, and each set of dynamic coefficients for countries (h, k) , $h, k = 1, \dots, n$, are individually considered, with the specific break that renders the highest finite sample p -value removed. The other F -tests are then recomputed and the process is repeated until all remaining (intercept and dynamic) breaks are significant.

In addition, dynamic interactions (or Granger causality) between the inflation series are examined by means of the heteroskedasticity-robust F -statistic. Such an analysis could be applied to the subperiods identified by the breaks in the autoregressive dynamics of the system, identified as in section IIIB. 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, the interaction (causality) analysis conditions on the significant breaks for individual coefficients using the procedure described in the preceding paragraph.

The significance of the resulting test statistics for both individual coefficient constancy and Granger causality is established through 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 subsamples, we employ a wild bootstrap, 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 & Kilian, 2004).

D. Volatility and Correlation Tests

As already noted, identified covariance breaks could originate from changes in either volatilities or correlations. For example, an increase in covariance could result from a decline in volatility or an increase in correlation. Since these have quite different implications for the nature of international inflation, it is important to attempt to identify volatility or correlation as the source of a covariance break. Indeed, correlation changes have been the focus of recent interest (see, for example, Neely & Rapach, 2011, Ciccarelli & Mojon, 2010, or Mumtaz & Surico, 2012).

Our analysis conditions on the number and dates identified for the m breaks in the conditional covariance matrix, obtained after convergence of the iterative procedure for dynamic coefficient and covariance breaks. Using the identity $\Sigma_j = \mathbf{D}_j \mathbf{P}_j \mathbf{D}_j$, we aim to distinguish between volatility and correlation changes, represented by \mathbf{D}_j and \mathbf{P}_j , respectively, for each break $j = 1, \dots, m$. This is conducted by first testing for volatility breaks and then, conditional on these, testing for correlation breaks.

The volatility test for a break at $t = T_j^{(C)}$ ($j = 1, \dots, m$) is implemented as a test of $H_0: \boldsymbol{\mu}_j = \boldsymbol{\mu}_{j+1}$ versus $H_A: \boldsymbol{\mu}_j \neq \boldsymbol{\mu}_{j+1}$, where $\boldsymbol{\mu}_j$ and $\boldsymbol{\mu}_{j+1}$ are vectors of the means of squared residuals for the n equations of the system over $t = \hat{T}_{j-1}^{(C)} + 1, \dots, \hat{T}_j^{(C)}$ and $t = \hat{T}_j^{(C)} + 1, \dots, \hat{T}_{j+1}^{(C)}$, respectively, and all residuals are computed allowing for the ℓ identified coefficient breaks. As in section IIIC, finite sample bootstrap inference is employed. When one or more breaks are individually not significant, the number of breaks is reduced by 1, and the tests are reapplied. However, no new search is undertaken for break dates, and hence the identified volatility break dates are a subset of the estimated covariance break dates $\hat{T}_j^{(C)}$ ($j = 1, \dots, m$).

Conditional on the significant volatility breaks, the VAR residuals are standardized, and breaks in the correlation matrix \mathbf{P} are examined employing the test statistic of Jennrich (1970) for $\mathbf{P}_j = \mathbf{P}_{j+1}$ in conjunction with finite sample inference. Once again, our procedure treats the potential correlation break dates as known and given by the covariance break dates $\hat{T}_j^{(C)}$ ($j = 1, \dots, m$). If breaks in \mathbf{P} are not significant at each of these dates, the least significant is dropped and the procedure repeated until all remaining correlation breaks are significant. Details of the volatility and correlation test procedures are provided in online appendix sections A.1 and A.2, respectively.¹²

¹² A Monte Carlo analysis of the procedure (results not included due to space constraints) shows it to perform well in distinguishing volatility and correlation breaks; details can be obtained from the authors.

Finally, international linkages are revealed through the nature of \mathbf{P} , and it is relevant to examine whether a specific country is contemporaneously related to inflation in other countries. Since correlation breaks may result in these changing from zero to nonzero (or vice versa), these tests are conducted for \mathbf{P}_j within each regime identified by the correlation break dates. The test employed is the instantaneous causality test of Lütkepohl (2005).

IV. Empirical Results

This section reports structural break results for both the four-country VAR consisting of Canada, the euro area, the United Kingdom, and the United States, and for the euro-area VAR comprising the individual countries of France, Germany, and Italy. The results obtained from the iterative system procedure of section IIIB are discussed first, before we turn to tests on individual VAR coefficients and the covariance matrix decomposition.

A. Coefficient and Covariance Breaks

In order to balance size and power for the structural change tests in relation to our sample size, the iterative procedure of section IIIB is applied with $\varepsilon = 0.15$ and a maximum number of breaks in each component of $M = 5$. All tests are conducted at the 5% significance level, and 1,000 replications are employed within the bootstrap procedure. As in section III, the VAR orders are set equal to $p = 1$ and 2 for the international and euro-area VARs, respectively, with these values being selected with BIC when a VAR model with constant parameters is estimated for the entire sample.¹³ With these low orders for monthly data, our analysis throughout the paper focuses on short-term international inflation dynamics.

All subsequent analysis builds on table 2, which shows the results of our iterative bootstrap testing procedure, together with the associated coefficient and covariance matrix break dates. For both systems, the procedure converges in three iterations between coefficient and covariance breaks. The WDmax statistics (the values shown in the table are those obtained at convergence of the iterative procedure) provide strong evidence for breaks in both the coefficients and the covariance matrix for each system when compared to the asymptotic critical values.¹⁴

¹³ The maximum lag order considered is $\text{int}[12(T/100)^{0.25}] = 17$. Diagnostic tests applied to the subsamples identified by the coefficient and covariance break dates (respectively) indicate that the residuals are substantially free of serial correlation and heteroskedasticity.

¹⁴ A practical problem in implementing this procedure is that the required asymptotic critical values have been tabulated for only small systems. Therefore, the critical values employed here are obtained by simulation, using the method described in Bai and Perron (1998) and Perron and Qu (2006). The Gauss program we employ for this has been modified from that used by Perron and Qu (2006) and is available from Zhongjun Qu's website (<http://people.bu.edu/qu/code.htm>). The program of Qu and Perron (2007), also obtained from that website, is the basis of our program for testing and estimating coefficient and covariance breaks.

TABLE 2.—ITERATIVE STRUCTURAL BREAK TEST RESULTS

International VAR		Euro-Area VAR	
Coefficients	Covariance Matrix	Coefficients	Covariance Matrix
Asymptotic WDmax test statistics (and critical values)			
174.57 [47.08]	210.39 [28.72]	157.66 [48.57]	289.44 [21.83]
Asymptotic sequential test statistics (and critical values)			
66.14 [41.86]	75.22 [30.29]	77.14 [42.60]	36.13 [22.64]
16.00 [44.79]	65.94 [31.44]	39.77 [45.81]	23.37 [23.56]
	N.C.		
Break dates, confidence intervals (and bootstrap <i>p</i> -values)			
1982.06	1983.07		1983.10
1982.04	1982.07		1982.07
1982.09	1983.09		1983.11
(0.00)	(0.00)		(0.00)
		1987.01	
		1986.11	
		1987.03	
		(0.00)	
1990.10	1992.04		
1990.04	1990.11		
1991.02	1992.07		
(0.00)	(0.01)		
	1999.03	1996.04	1996.02
	1999.02	1995.12	1994.10
	2000.07	1996.07	1996.08
	(0.00)	(0.00)	(0.00)

The table shows asymptotic WDmax and asymptotic sequential test statistics resulting from this procedure, with the latter comparing $l+1$ versus l breaks, beginning with $l=1$, with the respective 5% critical value in brackets. N.C. indicates that the test cannot be computed because no additional break can be inserted while satisfying the minimum regime length requirement. Estimated break dates (in bold) are based on the iterative bootstrap algorithm of section IIIB; this is followed by the 90% confidence interval for this date. The value in parentheses is the bootstrap percentage *p*-value for the specific break at convergence. The VAR lag lengths are $p=1$ for the international VAR and $p=2$ for the euro-area VAR.

For the international VAR, the sequential tests indicate two breaks in the coefficients, with the test for two versus one break giving a clear rejection of the null and that for three breaks versus two giving a clear nonrejection, based on the asymptotic critical values. Further, the bootstrap tests confirm the significance of each of these breaks, with empirical *p*-values that are very small. For the covariance matrix, both sequential tests are significant, providing evidence for at least three breaks. However, given the dates identified for three breaks and the restriction that a minimum of 15% of the sample must lie within each regime, a fourth break cannot be inserted, and hence the procedure stops at three, with these again highly significant according to the bootstrap verification. The test results for the euro-area VAR are straightforward and imply two breaks for each component. However, the final sequential test reported for the covariance matrix (for two versus three breaks) only very marginally fails to achieve significance at 5%.

Turning to the break dates, overall it appears that common breaks across the coefficients and covariance matrix may indeed apply in some cases, specifically in the early 1980s and early 1990s for the international system and the mid-1990s for the euro area, with the latter break date being during the run-up to full monetary union. Nevertheless, the international system also indicates a covariance matrix break at the end of the late 1990s, for which no corresponding coefficient break is detected.

In addition to the break dates, 90% confidence intervals, computed using the method of Qu and Perron (2007), are presented in table 2. Although these confidence intervals are relatively tight in all cases, they should be interpreted with care, since their computation effectively treats each (coefficient or covariance) component in isolation, assuming the break dates for the other to be known.

B. Individual VAR Coefficients

To study the nature of the structural changes detected in the VAR coefficients, the restriction of constant intercepts or constant dynamic coefficients over regimes j and $j+1$, as appropriate, is imposed whenever a potential individual break is not significant, using the recursive general to specific procedure described in section IIIC. The results shown for the international VAR in panel A of table 3 are the final estimate of the change for each individual coefficient examined (or sum of coefficients for the lag matrices), together with the corresponding percentage bootstrap *p*-value, either when the respective individual break is dropped during the recursive procedure or when the procedure stops because all remaining individual coefficient breaks are significant.

These results indicate, first, that intercept breaks are relatively unimportant in this system context. More specifically, the only intercept that changes significantly across either the 1982 or 1990 coefficient break date in table 3 relates to Canada in 1990, effectively corresponding to the introduction of inflation targeting in that country in 1991. In the light of previous evidence that breaks exist in univariate mean inflation across these countries (for example, Bataa et al., forthcoming; Cecchetti & Debelle, 2006), this implies that the source of such breaks may be changes in dynamics (whether in own-country dynamics or spillovers from other countries) rather than intercept changes.

Changes in dynamics for the international VAR show that inflation persistence significantly declines for all countries except Canada and all countries experience an own-coefficient change at the 1990 break. Relating these results in panel A to those in panel B, where coefficient estimates are shown after imposing the restrictions implied by the results of panel A, it is notable that whereas persistence is around 0.6 and highly significant for the United Kingdom and the euro area in the early part of the sample, persistence is insignificant for all four countries after 1990. Although in line with our univariate results (Bataa et al., forthcoming), the finding of zero persistence in table 3 differs in that account is taken here of (possibly changing) cross-country dynamic linkages.

Table 3 also provides evidence of changes in short-term international cross-country inflation dynamics, specifically relating to euro-area effects on both the United States and Canada and, conversely, for the United States on the euro area, all in 1982 and for Canada and the United Kingdom on the United States in 1990. Further, while substantial inflation spillover effects are seen in panel B from the euro

TABLE 3.—INDIVIDUAL COEFFICIENT BREAKS AND ESTIMATION RESULTS: INTERNATIONAL VAR

Subsample	Explanatory Variance	A. Coefficient Break Tests				B. Estimated Coefficients			
		Dependent Variable				Dependent Variable			
		Canada	United Kingdom	Euro Area	United States	Canada	United Kingdom	Euro Area	United States
1973.03–1982.06	Intercept					0.33 (0.0)	0.16 (0.0)	0.15 (0.0)	0.15 (0.0)
1982.07–1990.10		-0.20 (24.8)	-0.22 (30.20)	-0.13 (13.9)	-0.07 (63.1)				
1990.11–2007.12		-0.19** (0.1)	-0.06 (42.8)	0.01 (83.2)	0.00 (96.8)	0.14 (0.0)			
1973.03–1982.06	Canada					-0.14 (17.3)	0.11 (49.0)	0.06 (55.3)	0.06 (18.1)
1982.07–1990.10		-0.07 (59.8)	-0.08 (69.3)	-0.04 (52.9)	0.05 (70.7)				
1990.11–2007.12		0.28** (0.8)	0.04 (73.0)	0.01 (85.3)	0.08 (14.7)	0.15 (11.2)			
1973.03–1982.06	United Kingdom					-0.02 (79.7)	0.60** (0.0)	0.04 (21.9)	0.02 (90.6)
1982.07–1990.10		0.08 (30.8)	-0.29** (0.0)	-0.05 (24.9)	-0.01 (92.1)		0.31** (0.6)		
1990.11–2007.12		0.02 (88.7)	-0.24** (2.4)	0.14 (11.4)	0.06 (64.3)		0.06 (60.2)		
1973.03–1982.06	Euro area					0.71** (0.7)	0.07 (16.2)	0.55** (0.0)	0.39* (7.8)
1982.07–1990.10		-0.51** (0.0)	0.38 (56.0)	0.25 (13.7)	-0.26** (0.1)	0.19** (1.6)			0.13** (3.3)
1990.11–2007.12		0.17* (5.7)	0.12 (50.8)	-0.35** (0.0)	0.16 (29.6)		0.20 (16.1)		
1973.03–1982.06	United States					0.12 (43.3)	0.18 (34.5)	0.09 (17.1)	0.34** (1.3)
1982.07–1990.10		0.02 (89.5)	-0.07 (60.9)	-0.17** (0.0)	0.07 (65.2)		-0.07 (52.57)		
1990.11–2007.12		-0.30** (1.6)	-0.31** (0.1)	0.06 (49.6)	-0.37** (3.7)	-0.18** (3.8)	-0.13 (11.2)		0.17 (51.0)

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 first value in each cell in panel B reports the estimated coefficients (or sum of coefficients) over the indicated subsample. In both cases, the values in parentheses are bootstrap p -values (expressed as percentages) for the null hypothesis that the corresponding true value is 0. If an individual coefficient 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 those implied by the estimated structural break dates of table 2. Significant at **5%, *10%, both using the bootstrap p -value.

area to the United States and Canada until 1982 (although only the latter is significant), such effects are more muted in magnitude after this date.

Table 4 shows the euro-area results, based on the two coefficient breaks of table 2. Once again, there is only one significant intercept break in panel A, which relates to Germany in 1987. However, both France and Italy exhibit significant reductions in persistence, in 1987 and 1996, respectively. As for the international VAR in table 3, persistence is small and (except for Germany) insignificant in panel B for the last subsample identified for each country.

Nevertheless, table 4 provides evidence for the existence of dynamic interactions between these major euro-area countries, with these sometimes changing over time. For example, the significant (positive) dynamic effects from France to Germany and from Germany to Italy apparently cease in 1987, although inflation in Germany remains significant for France throughout the period, albeit with a coefficient that is relatively small in magnitude. Perhaps surprisingly, the final stages of the movement toward monetary integration in the euro area during the 1990s do not appear to have altered substantially the short-term dynamic linkages between these countries.

In the context of a univariate analysis of inflation, Cecchetti and Debelle (2006) argue that once changes in mean inflation are taken into account, then changes in persistence (or univariate dynamics) are relatively unimportant. To check whether a failure to take account of mean breaks drives our results in tables 2 and 3, the analysis was repeated after removal of the mean breaks found to be significant in the univariate analysis of this data in Bataa et al. (forthcoming). This mean-corrected VAR was analyzed without an intercept. Only one coefficient break was found in each system (in 1990 and 1987 for the international and euro-area VARs, respectively), but the pattern of declining own-country dynamics and insignificant persistence at the end of the sample period carries through. Indeed, Germany is the only country in that analysis not to experience such a break. However, there is less evidence of change in cross-country dynamics in the mean-corrected systems than in tables 2 and 3. For example, the inflation spillovers from the euro area to the United States and Canada are constant, with coefficients around 0.3 throughout the period.¹⁵

¹⁵ Detailed results are available on request.

TABLE 4.—INDIVIDUAL COEFFICIENT BREAKS AND ESTIMATION RESULTS: EURO-AREA VAR

Subsample	Explanatory Var.	A. Coefficient Break Tests			B. Estimated Coefficients		
		Dependent Variable			Dependent Variable		
		France	Germany	Italy	France	Germany	Italy
1973.03–1987.01	Intercept				0.12 (0.1)	-0.06 (35.6)	0.17 (0.0)
1987.02–1996.04		-0.01 (89.3)	0.21** (3.4)	0.08 (37.4)		0.12 (2.2)	
1996.05–2007.12		-0.02 (76.6)	-0.09 (36.8)	-0.07 (18.7)			
1973.03–1987.01	France				0.59** (0.0)	0.23** (0.3)	0.26 (13.0)
1987.02–1996.04		-0.68** (0.0)	-0.38** (2.4)	-0.19 (44.2)	-0.09 (27.9)	-0.15 (25.0)	
1996.05–2007.12		-0.25 (52.9)	-0.06 (98.1)	0.10 (89.3)			
1973.03–1987.01	Germany				0.15** (0.1)	0.27** (0.3)	0.49** (0.8)
1987.02–1996.04		-0.19* (5.8)	0.08* (6.8)	-0.57** (0.0)			-0.07 (87.6)
1996.05–2007.12		0.41 (34.2)	-0.29 (47.0)	-0.07 (85.0)			
1973.03–1987.01	Italy				0.13** (0.6)	0.10 (58.4)	0.53** (0.0)
1987.02–1996.04		0.22* (6.4)	-0.08 (65.1)	-0.14 (55.4)			
1996.05–2007.12		-0.08* (6.0)	-0.36 (43.1)	-0.56** (0.0)			-0.04 (94.2)

See table 3.

C. Volatility and Correlations

After imposing the restrictions implied by the results of tables 3 and 4, while also setting coefficients that are insignificant at 5% in panel B of those tables to 0 (except that an intercept is always retained in each equation), the results for tests of volatility and correlation breaks are reported in tables 5 and 6 for the international and euro-area systems, respectively. The break dates considered are those detected for the covariance matrix in table 2.

Considering first the international VAR, table 5 provides evidence that volatility changes apply at each of the three identified covariance matrix break dates. The volatility of inflation declines substantially in 1983 for all four countries except the euro area, while the 1992 break is particularly marked by a further substantial volatility reduction for the United States. In contrast, the 1999 break leads to higher volatility in the United States (especially), Canada, and the euro area. The overall pattern of these results is in line with the univariate ones in Bataa et al. (forthcoming); see also figure 1.

One of our primary interests is whether covariability, as captured by the cross-country inflation correlations, conditional on the removal of changing dynamic effects and volatilities, alters over time. For the international VAR (table 5), significant correlation breaks occur with the first and third covariance matrix breaks. On the other hand, the covariance change dated in 1992 appears to be due to volatility shifts. Although the 1983 break gives some evidence of correlation increases, notably for the United States with the United Kingdom and the euro area, all country pairs

experience substantial correlation increases in 1999. The increased exposure of the United States in the final period is particularly notable, having cross-correlations with the euro area and Canada of 0.58 and 0.63, respectively. The correlation test results reinforce the nature of the correlation breaks, with both Canada and the United States apparently being statistically contemporaneously isolated from other countries until 1983. However, this changes for the United States in 1983 and for Canada in 1999. These results in table 5 accord well with the broad patterns evident for volatility and disturbance correlations for these countries in figure 1.

Corresponding results for the three largest euro-area countries are given in table 6. As with the international VAR, conditional inflation volatilities for all three countries, but especially Italy, decline substantially at the first covariance matrix break in 1983. Although the volatility changes at the second euro-area break date of 1996 are statistically significant according to the bootstrap test, the magnitudes of change at this date are relatively small. More important, the contemporaneous correlations change significantly at only the first covariance break date, with the clear pattern being that of increased covariation from this date. In the first subsample (up to 1983), all three countries are contemporaneously immune from inflation arising in the other countries, but this changes from 1983 onward. Although the correlation of 0.47 between France and Germany in this later period is relatively high, figure 1 suggests that it may have increased further from the late 1990s. Consequently,

TABLE 5.—VOLATILITY AND CORRELATION RESULTS: INTERNATIONAL VAR

Subsample	A. Significance of Breaks		B. Subsample Residual Standard Deviations				C. Subsample Contemporaneous Correlations				
	Volatility	Correlation	Canada	United Kingdom	Euro Area	United States	Canada	United Kingdom	Euro Area	Zero Correlation	
1973.03 – 1983.07			0.32	0.41	0.16	0.32	Canada			17.9	
							United Kingdom	–0.16		0.1	
							Euro area	–0.03	0.32	0.3	
							United States	0.13	0.09	0.05	28.4
1983.08 – 1992.04	0.0	1.0	0.19	0.23	0.16	0.19	Canada			18.1	
1992.05 – 1999.03	0.1	17.7	0.20	0.18	0.11	0.10	United Kingdom	–0.01		0.0	
							Euro area	–0.03	0.25	0.0	
							United States	0.15	0.22	0.39	0.0
							Canada				0.0
1999.04 – 2007.12	0.0	0.0	0.30	0.19	0.14	0.27	United Kingdom	0.26		0.1	
							Euro area	0.49	0.43	0.0	
							United States	0.63	0.30	0.58	0.0
							Canada				0.0

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 2, 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 section IIID). The corresponding subsample 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 to that country are 0. All results are obtained from a VAR in which the restrictions implied by the results of coefficient breaks and persistence or dynamic interaction tests (at 5% significance) are imposed.

TABLE 6.—VOLATILITY AND CORRELATION RESULTS: EURO-AREA VAR

Subsample	A. Significance of Breaks		B. Subsample Residual Standard Deviations			C. Subsample Contemporaneous Correlations		
	Volatility	Correlation	France	Germany	Italy	France	Germany	Zero Correlation
1973.03 – 1983.10			0.21	0.23	0.34	France		88.5
						Germany	–0.02	94.0
						Italy	0.04	–0.02
1983.11 – 1996.02	0.0	0.0	0.15	0.17	0.14	France		0.0
1996.03 – 2007.12	0.0	25.4	0.18	0.18	0.10	Germany	0.47	0.0
						Italy	0.28	0.23

See table 5 for notes.

the lack of a significant correlation break in 1996 according to table 6 may be indicative of relatively low power for the correlation break test, especially if the increase is largely confined to a single country pair.

These increased correlations of tables 5 and 6 are not a consequence of the imposition of the restrictions on the VAR coefficients or of not removing (univariate) intercept breaks prior to the analysis. Although detailed results are not shown, when all VAR coefficients are allowed to be non-zero and to change at each of the VAR coefficient break dates indicated in table 2, the international VAR yields p -values of 7.2% and 0.0% for correlation breaks in 1983 and 1999, respectively. Thus, the 1983 correlation break does not survive, but the one in 1999 does. As implied by the results of table 5, the pre-1999 correlations are low, whereas those from 1999 are almost identical to those reported here. Also, prior removal of the univariate mean breaks as detected by Bataa et al. (forthcoming), and applying the same analysis as above to a VAR model with no intercept, yields effectively the same correlation results to the case with an intercept and where all coefficients are allowed to change. When such analyses are repeated for the euro-area system, the results are qualitatively unchanged from those of table 6, with a single correlation break in the 1980s (although the date changes to 1986 for a VAR with prior

removal of mean breaks). Although these analyses yield a covariance matrix break in 1995 or 1996, in both cases the p -value for this being a correlation break is around 20%.

V. Conclusion

This paper provides evidence on the nature of change in cross-country short-term dynamic variability and covariability for monthly consumer price inflation for major industrialized countries since 1973. For this purpose, we embed the recently developed testing procedures of Qu and Perron (2007) for structural breaks in multivariate time series 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 of finite sample bootstrap inference to identify the individual coefficients that change over time, test whether coefficients and contemporaneous correlations differ significantly from 0, and distinguish volatility breaks from changes in (conditional) correlations. Our aim is to develop flexible procedures that are able to provide more reliable structural break inference in samples of moderate size.

Three broad conclusions about international inflation dynamics can be drawn from our results. First, although structural stability tests reject constancy of the VAR coeffi-

cients, dynamic interactions (and changes in these) play a relatively modest role overall. The key cross-country dynamic effects that we find in our international VAR indicate that in the short run, euro-area inflation (at an aggregate level) leads that in both the United States and Canada, although the strength of this effect diminishes from the early 1980s. Within the euro area, Germany and Italy lead France throughout the period, with relatively modest spillovers, but no other dynamic interactions are significant after 1983.

The second general conclusion is that significant persistence and volatility changes occur for inflation, verifying the results from the univariate studies of Altissimo et al. (2006), Benati (2008), Cecchetti and Debelle (2006), O'Reilly and Whelan (2005), and others. Indeed, persistence breaks are found for all countries except Germany, with persistence then insignificant in the latter part of the sample.

Our third conclusion is the most important in terms of the light it sheds on the nature of changes in international inflation linkages. After allowing for dynamic effects and volatility shifts, the globalization of inflationary experiences is evident in the large increases in the contemporaneous correlations of inflation. In the case of the euro-area countries, the pattern of change may be due to the progress toward monetary union. The increased correlations between Germany, France, and Italy from essentially 0 to values up to 0.5 is notable. However, the most marked correlation changes relate to the integration of the United States (and Canada) into world inflationary experiences. The date at which we find this change occurs, 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.

The analysis in this paper also suggests directions for future research. One is to more explicitly examine changing patterns of common versus idiosyncratic shocks in international inflation through a dynamic factor augmented VAR (FAVAR). As noted in the section I, such models typically make (sometimes implicit) constancy assumptions, and hence conducting formal inference for structural breaks in this context would require a new testing procedure to disentangle breaks in the factor loadings, the dynamics and volatility of the common factors, and the volatility of idiosyncratic shocks. A second direction is to consider model specifications that allow longer-term inflation interactions in order to investigate whether different lead-lag relations apply to those uncovered here and whether these change over time.

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