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Changes in International Business Cycle Affiliations

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Abstract

We investigate changes in international business cycle affiliations using an iterative procedure for detecting system-wide structural breaks. We analyze GDP growth rates in two systems, one with the US, Euro-area, UK and Canada and the other for the Euro-area countries of France, Germany and Italy. We discover that international dynamic interactions change in both the mid-1980s and early 1990s, with such changes being particularly important for studying influences on the aggregate Euro-area. However, contemporaneous (conditional) correlations between these Euro-area countries increase in 1984 and 1998, with a large increase in correlations also evident across the international system during the 1990s.

Keywords: International business cycle, structural breaks, spillovers, business cycle synchronization

JEL codes: C32, E32, F43

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1. Introduction

There is now a substantial body of empirical evidence relating to the nature of international business cycle linkages and whether these have changed over the recent so-called globalization era¹. There is no doubt that some recessions are essentially global events, with those of the 1970s and the latest episode following the financial crisis being most notable in this sense. However, in line with much of this literature, the issue investigated in this paper is not the causes of these specific events, but the more general one of whether international business cycle affiliations, measured using cross-country linkages in output growth, have altered over the last 40 years. If cross-country affiliations have increased, as implicitly assumed in much of the general discussion of globalization, then purely domestic models become less relevant for explaining economic growth, even in the large G-7 countries.

Nevertheless, a somewhat surprising consensus appears to be emerging from many recent studies, namely that the era of globalization has not been associated with a general increase in the strength of international business cycle affiliations. For example, Heathcote and Perri (2004) find the business cycle correlation of US output with the rest of the world to be lower after 1986 compared with the earlier post-Bretton Woods period (1972-1986). Similar conclusions are drawn by Kose, Otrok and Whiteman (2008) in relation to output and consumption for the G-7 countries and by Del Negro and Otrok (2008) for output across 19 developed countries. Also, both Stock and Watson (2005) and Doyle and Faust (2005) find relatively little evidence of increased synchronization of business cycles across the G-7 countries since the 1960s.

However, other studies draw a different conclusion. Both Helbling and Bayoumi (2003) and Perez, Osborn and Artis (2006) find evidence of time-varying affiliations, with the early 1990s being distinctive as a period of relative disconnection between the US and major

¹ See Doyle and Faust (2005), Kose, Otrok and Prasad (2008) and de Haan, Inklaar and Jong-a-Pin (2008).

countries of mainland Europe and a restoration of strong trans-Atlantic links in the latter part of that decade. In contrast to other studies, Bordo and Helbling (2003) take a long-run perspective by using data over one and a quarter centuries, and find an increasing role for globalization over time. Kose, Otrok and Prasad (2008) find increased business cycle convergence within groups of industrial economies and groups of emerging market economies but between them a decoupling has been observed with divergence of these business cycle fluctuations and a decline in the importance of the global factor.

In the light of the huge changes in the formal structures linking European economies since the 1960s, which culminated in the establishment of the Euro area in January 1999, an important strand of analysis has focussed on changes in the business cycle affiliations of these countries; see the review in de Haan, Inklaar and Jong-a-Pin (2008). There is a general finding of stronger business cycle linkages over time between countries that are now Euro area members, especially for previously ‘peripheral countries’, such as Italy and Spain becoming more closely integrated with Germany and the Euro area more generally (for example, Artis and Zhang, 1997, 1999, Koopman and Azevedo, 2008). Similarly, increases are documented for the strength of the business cycle linkages of new European Union member countries with the Euro area; see Darvas and Szapáry (2008). Nevertheless there is also some evidence of regimes in affiliations rather than monotonic movement towards a common cycle (Inklaar and de Haan, 2001, Massmann and Mitchell, 2004). Interestingly, however, the literature in an international context often finds no evidence of changing affiliations for the three major economies of the Euro area, that is France, Germany and Italy (Canova *et al.*, 2007, Del Negro and Otrok, 2008, Kose, Otrok and Whiteman, 2003).

Although this literature employs a variety of econometric techniques, relatively little use has been made of formal tests for structural change at unknown dates. Rather, most studies either use essentially descriptive techniques (for example, output growth correlations based on rolling windows) or assume known dates of change. Although major international events

(such as the breakdown of the Bretton Woods system in the early 1970s or the ‘Maastricht Treaty’ that firmly committed a number of European countries to the formation of the Euro area in 1992) may lead to changes in business cycle linkages, such changes are not necessarily synchronous with the legal dates of such events and, indeed, could pre-date changes to formal structures when the latter are pre-announced or otherwise anticipated. Further, events in individual countries may affect these affiliations. For example, Helbling and Bayoumi (2003) associate low trans-Atlantic linkages in the early 1990s with a sequence of country-specific shocks over that period. Therefore, an appropriate econometric methodology for examining changes in international business cycle affiliations should allow for changes that are both unknown in number and occur at unknown dates. This is exactly what we do in this paper.

A further econometric complication is that many countries have experienced substantial changes in output volatility over the last four decades. This is best documented for the US (see Kim and Nelson, 1999, and McConnell and Perez-Quiros, 2000, Sensier and van Dijk, 2004, among others), but has also been established for other G-7 countries (van Dijk, Osborn and Sensier, 2002; Doyle and Faust, 2005), while Del Negro and Otrok (2008) refer to business cycle volatility as converging across countries. Consequently, results based on an explicit or implicit assumption of a constant cross-country disturbance covariance matrix may not be valid. To our knowledge, Doyle and Faust (2005) is the only previous study to employ formal tests for breaks in both the co-movement and volatility of international business cycle linkages.

In common with many previous analyses of the international business cycle, this paper examines quarterly GDP growth for G-7 countries within a vector autoregressive (VAR) framework. Our sample period of 1970 to 2008 allows us to focus on changes in business cycle affiliations over the post-Bretton Woods period, which allows us to examine the impact of changes relevant to the international economy, including globalization and the establishment of the Euro area. Following Doyle and Faust (2005), and also based on

evidence in many other recent studies (including Canova *et al.*, 2007, Del Negro and Otrok, 2008, Kose, Otrok and Whiteman, 2008), Japan is excluded from our analysis as it has not been closely linked to other G-7 economies since the 1970s.

Our analysis seeks to examine changes in the business cycle for both the G-7 as a whole (excluding Japan) and also between the countries that are now members of the Euro area. Reflecting this, we employ two VAR models, in order that the two types of cross-country changes can be clearly distinguished. One specification (which we term the ‘international VAR’) comprises the US, the Euro area, Canada and the UK, while the second (the ‘Euro area VAR’) examines the three Euro area countries that are members of the G-7, namely Germany, France and Italy. We consciously study the Euro area as an aggregate in the former, in order to recognise the international importance of this economic region, with aggregate output comparable to the US. While Canada and the UK are smaller, both of these countries have close trading links with the US and the Euro area. Further, the role of the UK in terms of links to the US and the Euro area has been a subject of much interest in the literature to date (for example, Artis and Zhang, 1997). Although a major world economy, Italy is of particular interest in a European context, since it is not historically part of the ‘core’ group of countries that now comprise the Euro area (Artis and Zhang, 2001), and hence changes in its business cycle affiliations with the key ‘core’ countries of Germany and France is of particular interest.

Our econometric methodology² is based on the system multiple break tests of Qu and Perron (2007), but we develop this further by separating mean and covariance breaks through an iterative approach. Further, covariance breaks are decomposed into variance and correlation breaks. While the broad approach is similar to that employed by Doyle and Faust

² The methodology is identical to that in Bataa, Osborn, Sensier and van Dijk (2009), where we examine international inflation linkages.

(2005), ours is more flexible in that we do neither specify *a priori* the number of breaks and nor are coefficient and covariance breaks required to be contemporaneous.

Our results contrast with much of the earlier literature, in that we find correlation breaks to be an important feature of international business cycle affiliations. More specifically, our results demonstrate that the Euro area is strongly linked (in terms of contemporaneous correlations) with the US, Canada and the UK from 1992 onwards, while correlations among the three large Euro area countries are very high since the launch of the euro. While the former points to international business cycles being a feature of the recent period of globalization, the latter accords with business cycle synchronization being endogenous with monetary union (Frankel and Rose, 1998).

The structure of this paper is as follows. Section 2 outlines our methodology, while Section 3 discusses our data and the results of a univariate analysis for each series. Our principal results on business cycle affiliations are then presented in Section 4, while Section 5 concludes.

2. Methodology

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

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

where $\mathbf{y}_t = [y_{1,t}, \dots, y_{n,t}]'$ is a vector of quarterly output growth rates. No intercept is included in (1), since all series are mean-corrected through our univariate analysis, discussed in subsection 2.5 below. 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 \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) $\mathbf{\Sigma} = \mathbf{D} \mathbf{P} \mathbf{D}$. Our methodology seeks to date structural breaks in each of the three components

of (1), namely output spillovers as captured by the VAR coefficients \mathbf{A}_i ($i = 1, \dots, p$), output volatility measured by \mathbf{D} , and contemporaneous output growth correlations in \mathbf{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 \mathbf{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. 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 the results of Doyle and Faust (2005) suggest the possibility of breaks in both components for international output growth, these need not occur at the same dates (and, consequently, the numbers of breaks need not even be the same). Indeed, the previous literature concerning the univariate properties of output growth implies volatility declines might be anticipated in the early 1980s (see, e.g., Sensier and Dijk, 2004), whereas globalization may affect dynamic linkages from the latter part of the century (Kose, Otrok and Whiteman, 2008).

For those reasons, we implement 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 2.1. Subsection 2.2 then describes the 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 hypothesis tests concerning business cycle linkages are then discussed in subsections 2.3 and 2.4. Finally, subsection 2.5 outlines our preliminary univariate analysis of structural breaks in mean, autocorrelations and volatility. Further details of these procedures can be found Bataa, Osborn, Sensier and van Dijk (2008, 2009).

2.1 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 $H_0 : \mathbf{A}_{i,j} = \mathbf{A}_{i,0}$ ($j = 1, \dots, m+1; i = 1, \dots, p$) in

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

for $t = T_{j-1} + 1, \dots, T_j$, $j = 1, \dots, m+1$, 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

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

where λ_j ($j = 1, \dots, m$) indicate possible break dates as fractions of the sample size, with $0 < \lambda_1 < \dots < \lambda_m < 1$ 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

$\sup_{(\lambda_1, \dots, \lambda_m \in \Lambda_\varepsilon)} F_T(m, q, \varepsilon)$ against a specific number of m breaks. For a total of q VAR coefficients

in (1), all of which are allowed to change,

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

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

$\hat{\mathbf{V}}(\hat{\boldsymbol{\beta}})$ ³, and \mathbf{R} is the non-stochastic matrix such that $(\mathbf{R}\boldsymbol{\beta})' = (\boldsymbol{\beta}'_1 - \boldsymbol{\beta}'_2, \dots, \boldsymbol{\beta}'_m - \boldsymbol{\beta}'_{m+1})$ 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

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

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 (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 $(\hat{T}_1, \hat{T}_2, \dots, \hat{T}_l)$ obtained from estimating the model with l breaks. The test is applied sequentially for $l = 0, 1, \dots$ until it fails to reject the null hypothesis of no additional break.

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

³ As there are potential breaks in the variance-covariance matrix in the residuals of (2), we use the Heteroskedasticity Consistent (HC) version when testing for the breaks in the conditional mean dynamics.

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

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

with $\hat{\mathbf{u}}_t$ ($t = 1, \dots, T$) the residual vector from (2), while \sim 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

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

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 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 the sample size for our quarterly data, we set $\varepsilon = 0.20$ with $M = 3$. Finally, we use a significance level of 5 percent for all tests.

2.2 Disentangling dynamic, volatility and correlation breaks

We adopt an iterative procedure to disentangle breaks in the VAR coefficients and in the conditional covariance matrix, which 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. In practice, the VAR order p in (1) is specified using the Hannan-Quinn criterion applied over the whole sample, using a maximum value of four, with the adequacy of this order checked for each subsample identified for the VAR coefficients using a heteroskedasticity robust serial correlation test, as in Godfrey and Tremayne (2005).

The approach to break detection initially examines 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 ($i = 1, \dots, p$) and in $\mathbf{\Sigma}$ until convergence, with the existence of identified breaks verified using finite sample inference; see below.

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 business cycle linkages, identifying volatility or correlation as the source of a covariance break is of crucial importance. Indeed, correlation changes are a key focus of interest for measuring the strength of international business cycles.

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 Σ . 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.

2.3 Individual coefficients breaks, spillover and contemporaneous correlation tests

The coefficient breaks resulting from the analysis outlined in subsection 2.2 apply to the VAR system as a whole. However, it is also of interest to identify whether these relate to changes in persistence of individual output growth series or to changes 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 \mathbf{R} defined as $(\mathbf{R}\hat{\boldsymbol{\beta}})' = (0, \dots, \hat{a}_{(h,k),1}^{(j)} - \hat{a}_{(h,k),1}^{(j-1)}, \hat{a}_{(h,k),2}^{(j)} - \hat{a}_{(h,k),2}^{(j-1)}, \dots, \hat{a}_{(h,k),p}^{(j)} - \hat{a}_{(h,k),p}^{(j-1)}, \dots, 0)$, where $\hat{a}_{(h,k),i}^{(j)}$ is the $(h,k)^{\text{th}}$ 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 output growth in country k on that of country h at all lags $i = 1, \dots, p$, 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, we employ the F -test 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 coefficient breaks are individually significant at the five percent level⁴.

In addition, spillovers (or Granger causality) between the output growth 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 business cycle 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 output 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 \mathbf{P} as identified by the correlation break dates. The test employed is the instantaneous causality test of Lütkepohl (2005).

2.4 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

⁴ Note that when the lag order $p > 1$, reference to an individual coefficient in this procedure should be understood as referring to a joint test on all coefficients $\hat{a}_{(h,k),i}^{(j)}$ at lags $i = 1, \dots, p$.

finite sample bootstrap analysis. In particular, 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 Σ provides some assurance that the asymptotic procedure does not lead to spurious breaks.

Finite sample inference is also conducted for all hypotheses tests. This includes tests of constancy of individual coefficients and spillovers, volatility constancy and correlation tests. 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, tests applied to specific VAR coefficients (including constancy tests applied jointly over lags $i = 1, \dots, p$) are based on 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). Further details of the bootstrap algorithms can be found in Bataa *et al.* (2009). All bootstrap inference is based on 5,000 replications.

2.5 Univariate analysis

A preliminary univariate analysis is employed in order to correct the data for outliers and mean breaks. By eliminating mean breaks a priori, the analysis of breaks in the VAR coefficients of (1) is able to focus more clearly on dynamic interactions. The univariate procedure is again based on testing for breaks using the methodology of Qu and Perron, but now applied to distinguish between shifts in the level (mean), persistence and volatility. Further, outlier detection is also undertaken in this iterative framework. The methodology is identical to that employed to analyze inflation in Bataa *et al.* (2008), except that seasonality is

not relevant in the present case as seasonally adjusted data are employed for our business cycle analysis.

The univariate decomposition for an observed series x_t can be written as:

$$x_t = L_t + O_t + y_t \quad (8)$$

$$L_t = \mu_j \quad t = T_{k_1-1} + 1, \dots, T_{k_1}; \quad k_1 = 1, \dots, m_1 + 1 \quad (9)$$

$$y_t = \sum_{i=1}^p \phi_{k_3, i} y_{t-i} + u_t \quad t = T_{k_2-1} + 1, \dots, T_{k_2}; \quad k_2 = 1, \dots, m_2 + 1 \quad (10)$$

$$\text{var}(u_t) = \sigma_{k_3}^2 \quad t = T_{k_3-1} + 1, \dots, T_{k_3}; \quad k_3 = 1, \dots, m_3 + 1 \quad (11)$$

where $t_0 = 0$; $T_{m_i} = T$ ($i = 1, 2, 3$) and T denotes the total sample size. Note that the number and timing of structural breaks in the level (L_t) and dynamic (y_t) components in (9) and (10), and also the volatility in (11) are not constrained to be equal.

In an analogous way to the multivariate analysis outlined in subsection 2.1, the univariate procedure iterates between testing for structural breaks in the level, dynamics and volatility components as well as testing for the presence of outliers. In these iterations, all components except the one under study are removed using the latest estimates. Thus, for example, level and dynamic components are removed when outliers are examined. To account for possible interaction between dynamic and volatility breaks, an additional ‘inner loop’ iterates between testing for breaks in the autoregressive coefficients of the dynamic component y_t and its conditional volatility. To be precise, after removing outliers and mean components, the sub-loop tests for breaks in dynamics; in the first iteration this employs heteroskedasticity robust inference, but subsequently volatility-regime estimates are used⁵. If any break is detected, the AR model is estimated allowing for these breaks, with variance

⁵ Pitarakis (2004) uses Monte Carlo simulations to assess the properties of mean break tests in the presence of volatility break, uncovering an extreme size distortion that actually increases with the sample size. He then provides evidence on improvements offered by a GLS transformation in that context.

breaks then investigated using the resulting residuals. If volatility breaks are detected, the variances are estimated over the implied segments. Once this ‘inner loop’ has converged, we return to the main loop.

Break detection for each of L_t , y_t and σ^2 uses the Qu and Perron (2007) methodology, while outliers are defined as observations more than a given distance (measured in terms of the interquartile range) from the median, using the procedure described in Stock and Watson (2003). By embedding the outlier analysis within the iterative procedure, changes in other characteristics are taken into account when testing and correcting for outliers.

Since fewer parameters are estimated in the univariate analysis compared with the multivariate models, the maximum number of breaks allowed in each of (9) to (11) is $M = 5$, while the minimum sample proportion in a regime set to $\varepsilon = 0.15$. Outliers are detected as observations lying beyond four times the interquartile range from the “local” (regime-dependent) median, while the order p in (10) is specified in an analogous way to the multivariate case (namely, using the Hannan-Quinn criterion for the whole sample and checking the adequacy via subsample heteroskedasticity robust tests for serial correlation). The dynamic component y_t obtained from this univariate analysis is then the data series input to the multivariate modelling. Consequently, this series is cleaned of mean breaks and outliers, but dynamic and volatility changes are retained.

3. Data and Univariate Results

3.1 Data issues

Our analysis employs quarterly real GDP growth (measured as 100 times the difference of the log values) over the period 1970Q1 to 2008Q4⁶. All series for individual countries are

⁶ Although data are available from 1960 onwards, this earlier data is not reliable for some European countries and hence we base our analysis on data from 1970.

obtained from the OECD database and are seasonally adjusted, with that for Germany taking account of the reunification in 1990.

Of course, the Euro area came into existence only in 1999 and its membership has expanded since that date. Although Greece was not formally a member until 2001, most Euro area analyses include Greece as its membership was anticipated. This is, however, less evident for members who have joined since the beginning of 2007 (Cyprus, Malta, Slovenia and Slovakia). Therefore, our Euro area series is confined to the “Euro 12” as of January 2001 and is constructed as a weighted average of the GDP growth rates for these 12 countries, using the weights employed in the historical data of the Area Wide Model (AWM) of the European Central Bank (Fagan, Henry and Mestre, 2001)⁷.

Many business cycle analyses filter GDP growth rate data in order to remove very short run fluctuations and hence concentrate on the so-called business cycle frequencies. However, such filtering has substantial consequences for the dynamics of the process and hence we prefer to analyze unfiltered data.

3.2 Univariate decompositions

The results of the univariate decomposition for each series (Canada, UK, Euro area, US, France, Germany and Italy) are shown in Table 1. In addition to the estimated break dates, 90% confidence intervals for these dates, computed using the methodology of Qu and Perron (2007) are reported. Although only the mean breaks and outliers are immediately relevant because these are removed prior to the cross-country analyses, nevertheless results for changes in dynamics and volatility provide benchmarks for the interpretation of the results in Section 4.

⁷ Updated AWM data to the end of 2007 is available at www.eabcn.org. Although we prefer to use our constructed data because this updated AWM data includes the new Euro area countries, the correlation between the growth rates for our Euro area GDP series and that of the AWM is 0.935 over the common data period.

– Table 1 about here –

Mean breaks in GDP growth are relatively rare, with Table 1 evidencing this only for the Euro area aggregate at the end of 1979 and for Italy in the mid-1980s. However, the wide confidence interval for Italy indicates that the break date is not precisely estimated. In both cases the mean growth rate approximately halves at the break date. Outliers are confined to the UK and the individual Euro area countries of France and Italy, with all of those identified being in the 1970s or the very beginning of the 1980s. In contrast to many inflation analyses, persistence for growth shows little evidence of change in Table 1, with only Germany experiencing such a change. Indeed, it is interesting that this change, which occurs around 2000, is such that strong persistence (with an autoregressive coefficient of 0.6) applies after that date but is effectively zero prior to that.

In contrast to the relative lack of mean and persistence breaks in growth, all countries except France have experienced volatility breaks. This finding of widespread volatility changes in the G7 countries is previously documented, for a range of series, in van Dijk *et al.* (2002). A relevant point for our subsequent multivariate analysis is that breaks in output growth volatility appear to cluster across countries. For example, the US volatility break dated in 1984 (Kim and Nelson, 1999, and McConnell and Perez-Quiros, 2000, and others) is shared (in the sense of overlapping confidence intervals) by the UK, Italy and the Euro area. The communality is even stronger for two further volatility breaks, namely the early 1990s for Canada, the UK, the Euro area and Germany, and early in the new century (the UK and the Euro area). The proximity of these break dates in this univariate analysis points to potential benefits from a multivariate approach, where the cross-sectional dimension may be especially useful for improving the precision of break date estimates as shown by Bai, Lumsdaine and Stock (1998).

Most breaks evidence a substantial decline of output volatility, but not all. In particular, the 2001 breaks identified for the UK and the Euro area indicate an end to the low volatility

era of the 1980s and 1990s. Nevertheless, volatility (in terms of the standard deviation) in the latter part of the period is typically around half of that during the 1970s.

4. International Business Cycle Changes

We now turn to the principal interest of this paper, namely changes in international business cycle affiliations. As already discussed, the series used in this analysis are de-measured and corrected for outliers, using the results of Table 1.

We analyze a four-country ‘International VAR’, consisting of Canada, the Euro area, the UK and the US, and also a ‘Euro Area VAR’ comprising the individual countries of France, Germany and Italy. A lag order of one is selected by the Hannan-Quinn criterion in both cases, with the residuals in the subsamples indicated by coefficient breaks all being free from first-order serial correlation⁸.

The iterative procedure (outlined in subsection 2.2) used to identify any dynamic and covariance matrix breaks yields the results as shown in Table 2, for both the International and Euro Area VARs. In addition to point estimates of the break dates, 90 percent confidence intervals (computed using the method of Qu and Perron, 2007) are presented. It may be noted that, in general, these confidence intervals are reasonably tight, covering two to five years. Also, these confidence intervals are considerably tighter than those obtained with the univariate analysis as shown in Table 1, which illustrates the usefulness of exploiting the cross-sectional dimension for dating (common) structural breaks.

– Table 2 about here –

Although our procedure places no restrictions across the coefficient and covariance matrices in relation to either the timing or the number of breaks, two breaks are identified for

⁸ The maximum lag considered in each case is eight. Using the selected lag order of one, the lowest p -value obtained from the system bootstrap test for first order serial correlation in any coefficient subsample is 18.4%.

both components in the International VAR, with the covariance matrix breaks in both cases preceding those for the dynamics. This feature of relative timing is also evident for the Euro Area VAR for the second covariance matrix break, but the 2000 break is the only one identified in that case for the coefficients. The covariance matrix break of 1984 is common to both systems, whereas the dating of a second break differs. Breaks in the international G-7 system are identified by Doyle and Faust (2005) in 1981Q1 and 1992Q2, which are more closely aligned with the covariance than the coefficient break dates for our International VAR. However, these authors do not separately study the Euro area countries, for which the coefficient and covariance breaks we uncover around the turn of the century suggest changes in business cycle affiliations between France, Germany and Italy in the run-up to, and the initial period of, full monetary integration⁹.

More detailed results are discussed for the coefficient and covariance matrices in subsection 4.1 and 4.2 respectively.

4.1 VAR coefficients

When considering the VAR coefficient breaks revealed in Table 2 for the International VAR, it should be recalled that Table 1 indicated no persistence breaks for any of these four individual economies. This is indicative of changes in the international transmission of growth, which is confirmed by Table 3. Panel A of Table 3 shows the outcome of the recursive general to specific procedure for identifying the sources of the coefficient breaks, and shows a significant (at 5 percent) change in 1986 in the equations for all countries except the US. In 1994, however, the changes are confined to impacts on the Euro Area. None of the

⁹ As indicated in Table 2, the iterative procedure to decompose coefficient and covariance matrix breaks did not converge in the case of the Euro area VAR. However, this lack of convergence relates only to the precise dating of the coefficient break, which may be difficult to date as it lies close to the limit implied by the restriction of each regime containing at least 20% of the total sample observations.

own lagged coefficients are found to change significantly in Table 3, in line with the univariate findings above.

– Table 3 about here –

Interpretation of these changes in coefficients is facilitated by the persistence/spillover results provided for the International VAR in Panel B of Table 3. One particular case of interest is the role of the US as the potential leader of world business cycle movements and, in particular, in relation to the Euro area. In this context, the role of the US apparently declines after 1986 (Panel A) and, according to Panel B, has a negative lagged effect on Euro area growth until 1994. Interestingly, over this period of the late 1980s to early 1990s, Canada apparently replaces the US in leading the Euro area cycle. On the other hand, growth in Canada is consistently led by the US, where the relevant coefficient changes remarkably little over subsamples (Panel A) and is highly statistically significant (Panel B). There is, however, no significant feedback from Canada to the US, whereas such a feedback applies from the UK to the US.

It is interesting that there is little evidence that the impact of the Euro area, as an aggregate, has changing spillover effects on other countries (Panel A), although it is notable that it consistently leads both Canada and, at the 10 percent significance level, the UK (Panel B). On the other hand, the UK has a feedback effect on the Euro area only from the end of 1994.

Table 4 provides comparable information to Table 3, but now for the individual Euro area G7 countries of France, Germany and Italy. Here the system test indicates only one significant coefficient break, in 2000. However, the individual coefficient test fails to detect a significant change (at 5 percent) in any individual coefficient, although the increase in the own coefficient for Germany is significant at 10 percent, which is in line with the change uncovered in the univariate analysis of the previous section. Consequently, the 2000 break is dropped for the analysis in Panel B.

– Table 4 about here –

Aside from significant own lagged effects for France and Italy, the significant interactions revealed in Panel B of Table 4 relate to Italy. In particular, there are highly significant spillovers from Italy to both France and Germany, but no dynamic effects originating in either of these two countries. While this result may appear surprising *a priori* in the light of previous analyses of European integration (such as Artis and Zhang, 1997, 1999, Inklaar and de Haan, 2001, Koopman and Azevedo, 2008), two points are worthy of note. Firstly, a number of coefficient changes in Panel A appear relatively large but are not significant at 5 percent; see, for example, that relating to France in the equation for Germany or Italy in the France equation, which may indicate a lack of power for the individual coefficient tests. Secondly, a lack of change in the VAR coefficients does not preclude changes in the contemporaneous correlations, which is analyzed in the next subsection.

4.2 Volatility and correlations

After imposing the subsample and causality restrictions implied by the results of Tables 3 and 4, 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 5 and 6 for the international and Euro area systems, respectively. The break dates considered are those detected for Σ in Table 2.

– Table 5 about here –

The first covariance break for the International VAR is dated at 1984. Although the first Euro area volatility break is dated earlier in Table 1, at 1978, the confidence interval for that break is extremely wide, while the UK volatility break of 1981 in that table falls within the lower bound of the confidence interval for the system break in Table 2. Of course, 1984 coincides with the well established break for US growth volatility (see, for example,

McConnell and Perez-Quiros, 2000). Although less well documented in the literature, a further volatility break in the early 1990s is evident for a number of countries (though not the US) in Table 1.

In contrast to conclusions from previous studies (including Doyle and Faust, 2005, Stock and Watson, 2005), however, we find that these dates mark not only volatility changes, but also changes in the extent of common business cycle movement, as captured by the contemporaneous disturbance correlation matrix \mathbf{P} . These correlation breaks are similarly highly significant. Before examining the individual correlations in Panel C of Table 5, note that the tests of zero correlation which consider, for each economy in turn, the null hypothesis that it is isolated from contemporaneous business cycle movements in all other countries¹⁰ reveal the presence of significant correlations (or common movements) across the four economies in the period 1970-1984, with even more significant linkages since 1992¹¹. However, except for Canada and (to a lesser extent), the US, they are contemporaneously disconnected during much of the 1980s.

The correlations themselves confirm this picture. All correlations in the period to 1984 are positive, with those between the US and Canada, and also between the UK and the Euro area being relatively strong, at 0.38 and 0.46 respectively. Although the UK economy is often considered to be strongly linked to the US, the correlation is only around half that with the Euro area aggregate over this period. The general pattern is of substantially lower correlations between 1984 and 1992, except for that between the US and Canada. Indeed, the correlation

¹⁰ It tests the null hypothesis that the off-diagonal elements in the corresponding row (or column) of \mathbf{P} are zero.

¹¹ This result is not a consequence of the sample period covering the onset of the 2008 recession, which is an international phenomenon due to the impact of the financial crisis. Employing a sample to 2007Q4 leads to substantively unchanged results, with significant correlation breaks dated in 1984Q2 and 1992Q2. However, the correlations in the final subsample are generally lower than those reported in Table 5, with those between the US and the UK and the Euro area, and also between the Euro area and the UK, being in the range 0.31 to 0.33.

between the UK and the Euro area is strongly negative over this period, while that between the Euro area and both the US and UK is effectively zero. Hence, except for the enduring US/Canada linkage, other contemporaneous business cycle movements are essentially idiosyncratic.

This idiosyncratic period is, however, reversed in 1992, with correlations subsequently returning to their levels of the 1970s and early 1980s or even exceeding, those. The weakest correlation in this final period is between the US and Canada, and the strongest between the UK and Euro area, indicating that the cycles between these latter two economies are now largely synchronized. Further, the large economic blocks of the Euro area and the US are strongly correlated at around 0.45, pointing to the existence of an important common international business cycle. Although the different recessionary periods across a number of major international economies at the beginning of the 1990s gave rise to a view that the Euro area business cycle may have become disconnected from that in the US (see, for example, International Monetary Fund, 2001), our results (in line with those of Perez *et al.*, 2006) imply that any such phenomenon was temporary.

Table 6 presents a comparable analysis to Table 5, but now for the three Euro area countries. An important feature of Panel A is that, although the 1984 covariance matrix break is a both a volatility and a correlation break, that of 1999 is found to be a correlation break only. In line with the pattern of the univariate volatility results of Table 1, volatility in Table 6 appears effectively unchanged for France, while it declines in 1984 for both Germany and Italy.

– Table 6 about here –

Panel C of Table 6 clearly shows the impact of economic integration on business cycle affiliations across these countries. Initially, in the period to 1984, Italy is isolated from contemporaneous movements in other countries, with the joint test for zero correlation having a high p -value and the individual cross-country correlations with Italy being small (although

positive). Although Germany and France exhibit a substantial positive correlation, around 0.4-0.5 throughout, Italy becomes strongly correlated with France from 1984 and with Germany from 1998. Indeed, with correlations around 0.7 with both of these countries from 1998, business cycles are more highly synchronized between Italy and these countries than between France and Germany. These results closely resemble those in Bataa *et al.* (2009) for inflation linkages between these countries, for which contemporaneous correlation breaks are uncovered in mid-1984 and late 1996. In particular, strong contemporaneous correlations apply for inflation across all three countries from the latter part of the 1990s, as for growth in Table 7.

5. Conclusions

The principal finding of the present analysis is that business cycles, as captured by the growth rate of quarterly GDP, have been more closely aligned in the main developed economies since 1992 than at any previous period since 1970. This is particularly important in relation to the two major economic blocks of the developed world, namely the US and the Euro area, which we find to experience strong synchronous business cycle movements from 1992 onwards. Indeed, since we find no (positive or negative) growth spillovers from US to the Euro area or vice versa from the early 1990s, it appears that the business cycles in these economies are now largely aligned.

Alongside these results, Canada is strongly connected to the US throughout the period from the 1970s, with spillovers from the US being an important feature of the Canadian business cycle. Perhaps surprisingly, the only spillovers to significantly impact on the US are from the UK, with the UK also having bidirectional spillover effects with the Euro area (although the UK to Euro area one applies only from 1994). Indeed, although the alignment of the UK business cycle with that of the Euro area has been one issue in terms of the potential

for the UK joining the Euro area, our correlation results indicate that they were seriously out of phase only between 1984 and 1992.

For the three key Euro area countries of France, Germany and Italy, any changes to the dynamic linkages between the countries are relatively unimportant, in contrast to a pattern of increasing contemporaneous correlations. In particular, whereas France and Germany maintain consistently strong linkages, the business cycle in Italy is affiliated strongly first with France (in 1984) and then Germany (1998).

Methodologically, our analysis employs the same iterative decomposition of coefficient and covariance breaks as in Bataa *et al.* (2009), and we find a similar pattern of results for the international business cycle in this study as for international inflation in our earlier analysis. In particular, the period of full monetary union leading up to the launch of the euro currency in 1999 is associated with substantial increases in contemporaneous correlations between the participating economies of France, Germany and Italy, but little or no change in their dynamic interactions. However, whereas a strong increase in contemporaneous inflation correlations is detected between the US and the Euro area around this date, we find the change in business cycle correlations to occur in 1992.

Our results about the strength of the international business cycle over the recent period of globalization contrasts with other studies, including Doyle and Faust (2005), despite the similarity of the methodology employed in that paper compared with our study. There are, however, also a number of important differences. From a methodological perspective, the most important difference is that our approach is more flexible in allowing coefficient and covariance breaks to occur at different dates and in recursively separating volatility and correlation breaks, while also treating the latter in a system context rather than examining bivariate correlations. We believe that this flexibility is important in delivering our results. For instance, not only do we find that coefficient and covariance breaks do not necessarily coincide, but the numbers of such breaks may differ. Further, covariance matrix breaks do not

necessarily involve both volatility and correlation breaks, so that an assumption to this effect may adversely affect power.

Another important feature of our analysis is that we treat the Euro area aggregate as an entity in the international analysis and individual Euro area countries in another. This is also important, because the dates at which key historical intra-Euro area changes have occurred do not necessarily coincide with changes in the international business cycle more broadly. Nevertheless, the results we obtain in this paper relating to the business cycle and in Bataa *et al.* (2009) for inflation indicate that the establishment of the Euro area has led to important changes in international economic linkages.

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Table 1. Univariate GDP Growth Decomposition

Canada	UK	Euro Area	US	France	Germany	Italy
A. Mean (Level) Breaks						
		1979Q4 1977Q1 1982Q3				1985Q4 1979Q4 1991Q4
B. Means in Each Regime						
0.73	0.55	0.92 0.47	0.71	0.57	0.55	0.75 0.35
C. Outliers						
	1973Q3 1974Q3 1979Q4 1980Q1			1975Q2		1970Q3
D. Dynamic Component AR Order						
1	3	2	1	2	0	1
E. Dynamic Breaks						
					2000Q1 1996Q2 2003Q4	
F. Persistence in Each Dynamic Regime						
0.38	0.43	0.50	0.27	0.59	-0.08 0.62	0.48
G. Volatility Breaks						
	1981Q2 1971Q2 1982Q1 1991Q3 1984Q4 1992Q1	1978Q3 1914Q4 2025Q2 1992Q1 1988Q3 1992Q3 2001Q2 1995Q1 2004Q2	1984Q3 1972Q3 1985Q1		1993Q3 1980Q1 1994Q2	1984Q3 1971Q3 1985Q1
H. Standard Deviation in Each Volatility Regime						
0.94 0.43	1.09 0.66 0.29 0.47	0.64 0.56 0.27 0.44	1.16 0.53		1.12 0.61	0.95 0.52
I. Number of Iterations (Inner Loop)						
2 (2)	19* (2)	3 (2)	2 (2)	3 (2)	2 (3)	4 (2)

Notes: Decomposition into mean, outlier and dynamic components uses the iterative method outlined in the text. Breaks are detected using the Qu and Perron (2007) structural break test (Mean: trimming 15%, max breaks 5), with confidence intervals computed using HC standard errors. Panel A shows the estimated break dates for the level component in bold, together with lower and upper bounds of asymptotic 90% confidence intervals. Mean quarterly growth rates in the various subsamples determined by the level breaks are given

in Panel B. The dates of detected outliers are given in Panel C, where an outlier is defined as being four times the inter-quartile range from the median. Panel D indicates the autoregressive order of the dynamic component, selected according to the HQ information criterion, and used at entry to the dynamic/volatility sub-loop. If the selected order is zero, this is replaced by an order of one. Panel E reports the estimated break dates in AR coefficients (in Bold) with asymptotic 90% confidence intervals immediately below, while Panel F reports estimated persistence, defined as the sum of autoregressive coefficients, based on sub-samples defined by the break dates. Panel G shows the estimated volatility breaks (in Bold) and their asymptotic 90% confidence intervals, with Panel H reporting the corresponding estimated standard deviations of the errors. Finally, Panel I shows the number of iterations required for convergence of the main loop, with * indicating that the iteration converged to a two cycle oscillation and choice between these is made based on HQ criterion; the number of iterations required for convergence of the volatility/persistence loop is shown in parentheses.

Table 2. System Structural Break Test Results

A. International VAR		B. Euro Area VAR	
VAR Coefficients	Covariance Matrix	VAR Coefficients	Covariance Matrix
1986Q4	1984Q2		1984Q1
1985Q3	1979Q2		1980Q3
1987Q4	1984Q3		1984Q2
1994Q3	1992Q1		1998Q2
1992Q3	1989Q2		1980Q4
1996Q3	1992Q2		1998Q3
		2000Q3	
		1995Q3	
		2001Q4	
C. Number of Iterations			
2		9*	

Notes: The table shows estimated break dates (first value, in bold) followed by 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 (US, Euro area, Canada and UK) and Panel B to the Euro area VAR (France, Germany and Italy). A maximum of three breaks is permitted, with a minimum of 20% of the observations in each regime. For each VAR, Panel C reports the number of iterations required for convergence, with * indicating that no convergence is obtained and the break dates are selected based on the minimum HQ criterion.

Table 3. Individual Coefficient Test Results: International VAR

Subsample	Explan. Var.	A. Coefficient Breaks				B. Persistence and Spillovers			
		<u>Dependent Variable</u>				<u>Dependent Variable</u>			
		Canada	UK	Euro Area	US	Canada	UK	Euro Area	US
1970Q1 - 1986Q4	Canada					0.12* 6.3	-0.17 18.6	-0.05 58.0	0.11 77.2
1987Q1 - 1994Q3		0.06 86.0	0.52** 0.0	0.62** 0.4	0.08 86.8		0.36** 4.8	0.57** 3.2	
1994Q4 - 2008Q4		0.27 10.7	-0.44* 5.8	-0.51** 2.6	-0.08 82.4			0.10 33.2	
1970Q1 - 1986Q4	UK					-0.09 34.7	0.06 30.7	-0.09 45.7	0.20** 2.2
1987Q1 - 1994Q3		0.58** 0.0	0.19 36.8	-0.09 60.7	0.29 19.5	0.50 17.3			
1994Q4 - 2008Q4		-0.54 14.1	0.26* 5.8	0.58** 0.1	0.43* 5.4			0.49** 5.0	
1970Q1 - 1986Q4	Euro Area					0.13** 2.7	0.24* 6.7	0.40** 0.0	0.01 94.4
1987Q1 - 1994Q3		-0.55* 7.4	-0.05 86.3	0.18 38.4	-0.14 53.7				
1994Q4 - 2008Q4		0.52 21.1	0.15 63.1	-0.22 56.6	-0.05 88.1				
1970Q1 - 1986Q4	US					0.38** 0.2	0.19 22.9	0.22** 1.4	0.17 42.0
1987Q1 - 1994Q3		-0.04 86.6	-0.02 93.4	-0.87** 0.04	-0.21 36.1			-0.66** 1.0	
1994Q4 - 2008Q4		-0.06 53.2	-0.09 46.1	0.77** 0.5	0.09 76.8			0.11 23.2	

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 2. ** indicates significance at the 5% level and * significance at the 10% level, both using the bootstrap p -value.

Table 4. Individual Coefficient Test Results: Euro Area VAR

Subsample	Explan. Var.	A. Coefficient Breaks			B. Persistence and Spillovers		
		<u>Dependent Variable</u>			<u>Dependent Variable</u>		
		France	Germany	Italy	France	Germany	Italy
1970Q1 – 2000Q3	France				0.30** 0.2	0.23 36.0	0.07 68.6
2000Q4 – 2008Q4		-0.32 16.6	0.55 28.2	0.21 54.4			
1970Q1 – 2000Q3	Germany				-0.01 77.6	-0.14 32.5	0.08 27.5
2000Q4 – 2008Q4		-0.06 80.8	0.48* 6.9	-0.12 62.9			
1970Q1 – 2000Q3	Italy				0.23** 0.0	0.34** 0.6	0.43** 0.1
2000Q4 – 2008Q4		0.52 15.8	-0.24 58.0	0.29 23.1			

Notes: see Table 3.

Table 6. Volatility and Correlation Test Results: Euro Area VAR

Subsample	A. Significance of Breaks		B. Subsample Residual Standard Deviations			C. Sub-sample Contemporaneous Correlation		
	Volatility	Correl.	France	Germany	Italy	France	Germany	Zero Correlation
1970Q1 - 1984Q1			0.46	1.01	0.76	France		0.2
						Germany	0.46	0.1
						Italy	0.09	0.15
1984Q2 - 1998Q2	0.1	0.5	0.43	0.63	0.51	France		0.0
						Germany	0.37	0.9
						Italy	0.61	0.15
1998Q3 - 2008Q4	13.6	0.1				France		0.0
						Germany	0.53	0.0
						Italy	0.68	0.74

Notes: See Table 5.

Table 5. Volatility and Correlation Test Results: International VAR

Subsample	A. Significance of Breaks		B. Subsample Residual Standard Deviations				C. Subsample Contemporaneous Correlations			
	Volatility	Correl.	Canada	UK	Euro Area	US	Canada	UK	Euro Area	Zero Correlation
1970Q1 - 1984Q2			0.88	1.01	0.54	1.18	Canada			2.1
							UK	0.13		0.3
							Euro Area	0.20	0.46	0.3
							US	0.38	0.23	0.11
1984Q3 - 1992Q1	0.0	0.4	0.78	0.59	0.46	0.42	Canada			2.3
							UK	0.16		12.1
							Euro Area	-0.14	-0.39	13.2
							US	0.46	0.03	0.06
1992Q1 - 2008Q4	0.0	0.0	0.42	0.41	0.40	0.52	Canada			0.2
							UK	0.24		0.0
							Euro Area	0.38	0.52	0.0
							US	0.42	0.43	0.45

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.