Available online at www.sciencedirect.com
ScienceDirect
Journal of International Money and Finance 26 (2007) 149–165
Journal of International Money and Finance
www.elsevier.com/locate/jimf

A cross-country financial accelerator: Evidence from North America and Europe

Ashoka Mody [a], Lucio Sarno [b,c,*], Mark P. Taylor [c,d]

[a] *Research Department, International Monetary Fund, 700 19th St. NW, Washington, DC 20431, USA*
[b] *Finance Group, Warwick Business School, University of Warwick, Coventry CV4 7AL, UK*
[c] *Centre for Economic Policy Research (CEPR), 90-98 Goswell Road, London EC1V 7RR, UK*
[d] *Department of Economics, University of Warwick, Coventry CV4 7AL, UK*

Abstract

A growing literature has examined the importance of credit-market imperfections for macroeconomic fluctuations, the so-called 'financial accelerator.' A related literature has provided evidence of international and regional comovements in macroeconomic fluctuations. We tie together these strands of the literature in that we investigate the importance of both cross-country and country-specific credit cycles in explaining output fluctuations. Using data for four major economies and two world regions from 1973 to 2001, we find that both regional and country-specific components of indicators of credit availability are powerful in explaining output movements. This research provides the first empirical evidence of a cross-country financial accelerator.
Published by Elsevier Ltd.

JEL classification: E32; E51; F36

Keywords: Financial accelerator; Credit cycle; International business cycles; Kalman filter

1. Introduction

It is well known that credit-market conditions may have important effects on an economy's business cycle (see Bernanke et al., 1996, 1999; Hubbard, 1998, and the references therein). A

* Corresponding author. Finance Group, Warwick Business School, University of Warwick, Coventry CV4 7AL, UK. Tel.: +44 2476 528219; fax: +44 2476 572871.
E-mail address: lucio.sarno@warwick.ac.uk (L. Sarno).

0261-5606/$ - see front matter Published by Elsevier Ltd.
doi:10.1016/j.jimonfin.2006.10.009

number of authors have argued that these effects may be amplified at the macroeconomic level so that, for example, adverse shocks to the economy are exacerbated by worsening credit-market conditions which generate increasing credit rationing which, in turn, adversely affects economic activity, and so on in a vicious circle. The generic term for this effect, which arises in a number of models involving credit-market imperfections, is the 'financial accelerator.' Spurred by these theories, a growing empirical literature has provided evidence supporting the existence of a link between indicators of credit availability and macroeconomic fluctuations at the country level, suggesting that credit-market conditions tend to impact significantly on measures of real activity over the business cycle (e.g. Bernanke et al., 1999; Gertler and Lown, 1999; Mody and Taylor, 2003).

A separate strand of research has investigated the existence of cross-country linkages in macroeconomic fluctuations, generally providing convincing evidence that business cycles of major industrialized economies are highly correlated (Backus et al., 1995; Baxter, 1995; Sarno, 2001). For example, using dynamic factor analysis, Gregory et al. (1997) decompose aggregate output, consumption, and investment for the G7 countries into factors that are common across all countries and aggregates, common across aggregates within a country, and specific to each individual aggregate. In quarterly data for the period from 1970 through 1993, fluctuations in all aggregates are found to contain world and country-specific common components that are both statistically significant and quantitatively important.[1]

Other studies argue that business cycle comovements are stronger for subsets of countries, so that it is possible to identify a world component in macroeconomic fluctuations as well as regional and country-specific components. In particular, several studies provide evidence in favor of the existence of a world business cycle or a European business cycle (e.g. Artis et al., 1997; Artis and Zhang, 1997; Bergman et al., 1998; Imbs, 2003; Lumsdaine and Prasad, 2003). This evidence seems to suggest that, in addition to a world component that characterizes output comovements across countries, one might also expect a regional component, common to a subset of countries, and a country-specific component.[2]

To date, however, the literature on the financial accelerator and the literature on international business cycles have moved in parallel. In particular, researchers have not investigated the extent to which international and regional components of credit cycles may be important for the international and regional transmission of business cycles. This is the purpose of the research reported in this paper. Using Kalman filtering maximum likelihood techniques and vector autoregressive (VAR) modeling, we examine the dynamic interactions linking real activity to international and regional components of an indicator of credit availability. The international and regional components of real activity and credit variables are first extracted and analyzed using multivariate Kalman filtering techniques applied to dynamic multiple indicator-multiple cause (DYMIMIC) models of the relevant time series for four G7 countries over the 1973–2001 sample period.[3] The four countries we examine are the US and Canada, which we view as comprising the North American region, and Germany and France, which we view as representative of the European region. The dynamic interactions linking these cross-country

[1] In a study of postwar economic growth across the G7 countries, using nonlinear equilibrium correction models Sarno (2001) also provides evidence supporting the existence of significant spillover effects in economic growth dynamics across the G7, which appear to be largely driven by the US business cycle.

[2] Using a world dynamic factor model, Kose et al. (2003) argue that the finding of a European business cycle commonly reported by this literature may be an artifact of limited samples.

[3] See Engle and Watson (1981) for an early application of DYMIMIC modeling.

components are then examined using multivariate vector autoregressive modeling and impulse response analysis.

We find that real GDP for all four countries does indeed contain an important international component—which appears to be well characterized by a stochastic trend—as well as stationary regional and country-specific components. Therefore, one could interpret the international component as driven by supply-side (e.g. technology) shocks with important permanent components that are largely common to the four major economies examined. The stationary regional and country-specific components may be interpreted as driven by nominal regional and country-specific shocks. The time series of bank credit is well described by the same model used for real GDP, i.e. as the sum of an international permanent component and regional and country-specific temporary components. Using the components extracted from our DYMIMIC model, estimation of a trivariate VAR model for GDP growth, the regional component of bank credit and the country-specific component of bank credit, reveals that both regional and country-specific components of bank credit are empirically important determinants of GDP growth for each of the four countries examined. In particular, we find unidirectional causality from regional and country-specific bank credits to GDP growth for each country investigated over the sample. The impulse response functions of GDP growth in response to bank credit shocks implied by our VAR models appear to be consistent with the propagation mechanism implied by the financial accelerator and with previous empirical studies which have ignored international and regional effects. Overall, this research provides the first empirical evidence of the presence of a cross-country financial accelerator.

The remainder of the paper is set out as follows. Section 2 offers a brief review of the literature on the financial accelerator. In Section 3, we describe our Kalman filtering DYMIMIC approach, which is designed to extract international, regional and country-specific components from the economic time series of interest. In Section 4, we describe our data set and report the empirical results from applying Kalman filtering techniques to extract international, regional and country-specific components from each of real GDP and bank credit to the private sector as well as the results from estimating VAR models designed to examine the dynamic interactions linking cross-country components of bank credit and GDP growth. A final section concludes.

2. The financial accelerator: a brief review

Loosely speaking, the ‘financial accelerator’ is the mechanism that much theoretical macroeconomics has used to characterize how financial factors may amplify and propagate business cycles. Although the various models inspired by the financial accelerator may differ in a number of ways, they have as a common factor the assumption that frictions in financial markets generate a wedge between the cost of external funds and the opportunity cost of internal funds—the ‘premium for external funds.’ These frictions may be associated, for example, with asymmetric information problems or the costs of enforcing contracts. The premium for external funds is an endogenous variable, which depends inversely on the balance sheet strength of the borrower, since the balance sheet is the key signal through which the creditworthiness of the firm is evaluated. However, balance sheet strength is itself a positive function of aggregate real economic activity, so that borrowers’ financial positions are procyclical and hence movements in the premium for external funds are countercyclical. Thus, as real activity expands, the premium on external funds declines, which, in turn, leads to an amplification of borrower spending, which further accelerates the expansion of real activity. This is the basic mechanism of the financial accelerator: frictions in financial markets amplify fluctuations in the spending of borrowers and hence exacerbate fluctuations in aggregate economic activity, relative to a world with frictionless financial markets (see e.g. Bernanke et al., 1999; Hubbard, 1998).

The empirical evidence on the existence of a financial accelerator mechanism is dominated by panel data studies which, in general, report that financing constraints affect the behavior of both consumers and firms using individual-level data (e.g. Bernanke et al., 1999). Little evidence exists, however, on the statistical and economic relevance of the financial accelerator at the aggregate level. First attempts in this direction include the studies of Gertler and Lown (1999) and Mody and Taylor (2003). Gertler and Lown (1999) start from noting that the 'high-yield spread'—the spread between the yield on below-investment-grade or 'junk' corporate bonds and that on investment-grade bonds such as government or corporate AAA-rated debt—may be interpreted as closely approximating the 'premium for external funds' which is a central variable in theories of the financial accelerator. In particular, financial accelerator theories imply that a rise in the high-yield spread would indicate a tightening of credit-market conditions and therefore a future decline in real economic activity. Gertler and Lown do indeed find that, since the inception of the 'junk bond' market in the mid-1980s, the high-yield spread outperforms several other leading indicators of economic activity. Mody and Taylor (2003) also focus on the ability of the high-yield spread to predict real economic activity, using long-horizon regressions, and find that the high-yield spread does indeed predict movements in real US industrial output well during the 1990s and, in particular, is superior to the term spread as a leading indicator of real economic activity over this period. This is an interesting finding since there is a large literature supporting the view that the term spread has predictive content for real economic activity for the US and other industrialized countries (see e.g. Dotsey, 1998; Mody and Taylor, 2003, and the references therein).[4]

In general, however, empirical investigation of the financial accelerator has tended to focus on a single country and in particular on the US. Given that business cycles appear to have important international components, it seems logical to investigate the extent to which international and regional components of credit cycles may be important for the international and regional transmission of business cycles. In an attempt to examine this issue empirically, we employ a VAR model which allows us to shed light on the dynamic interactions linking real activity to international and regional components of an indicator of credit availability. Prior to the VAR analysis, however, the international and regional components of real activity and credit variables are extracted using DYMIMIC models, which we describe in detail in the next section.

3. Extracting international, regional and country-specific components from economic time series

The model we employ for extracting international, regional and country-specific components from the economic time series of interest in this paper—namely, real GDP and bank credit to the private sector—is based on an extension of the unobserved components' model (e.g. Harvey, 1989; Gregory et al., 1997). The essential idea is to break the time series of interest down into unobserved permanent and temporary components that are international

[4] In addition, Mody and Taylor (2003) find some evidence of nonlinearity in that abnormally high levels of the high-yield spread have significant additional short-term predictive power.

(common to all time series or countries in the system), regional (common to a subset of time series or countries in the system) or country-specific, using maximum likelihood estimation.

Consider a panel of N countries, each belonging to one of K regions ($j=1,\ldots,K$), with the time series of interest for country $i=1,\ldots,N$ at time t denoted p_{it}. The model may be written as follows:

$$p_{it} = i_{\pi t} + i_{\tau t} + r_{jt} + c_{it} + \theta_{it}, \tag{1}$$

where $i_{\pi t}$ is an international permanent component, $i_{\tau t}$ is an international temporary component, r_{jt} is a regional component, c_{it} represents a country-specific autoregressive component, and the country-specific irregular component θ_{it} is approximately normally independently distributed (NID) with zero mean and constant variance, i.e. $\theta_{it} \sim \mathrm{NID}(0, \sigma^2_{\theta_i})$.

The international permanent component, which represents the common trend for all time series in the system, is modeled as:

$$i_{\pi t} = i_{\pi t-1} + \varepsilon_{\pi t}, \tag{2}$$

while the international temporary component is modeled as a stationary first-order autoregressive, or AR(1) process

$$i_{\tau t} = \rho i_{\tau t-1} + \varepsilon_{\tau t}, \tag{3}$$

where $|\rho| < 1$ determines the degree of persistence of the component; $\varepsilon_{\pi t} \sim \mathrm{NID}(0, \sigma^2_{\varepsilon_\pi})$ and $\varepsilon_{\tau t} \sim \mathrm{NID}(0, \sigma^2_{\varepsilon_\tau})$.

The regional component, common to a subset of economic time series in the system, is modeled as

$$r_{jt} = \beta_j r_{jt-1} + \eta_{jt} \quad j = 1, \ldots, K \tag{4}$$

where β_j is unrestricted and determines the degree of persistence of the regional component, and $\eta_{jt} \sim \mathrm{NID}(0, \sigma^2_{\eta_j})$.

The country-specific component, idiosyncratic to a particular country i, is assumed to follow the process

$$c_{it} = \gamma_i c_{it-1} + u_{it} \quad i = 1, \ldots, N \tag{5}$$

where γ_i is unrestricted and determines the degree of persistence of the country-specific component, and $u_{it} \sim \mathrm{NID}(0, \sigma^2_{u_i})$.

Finally, the disturbance terms driving each of the components—namely the disturbances of the international permanent and temporary components $\varepsilon_{\pi t}$ and $\varepsilon_{\tau t}$, the regional disturbances η_{jt} for $j=1,\ldots,K$, the country-specific disturbances u_{it} and θ_{it} for $i=1,\ldots,N$—are assumed to be mutually uncorrelated.

Intuitively, Eq. (1) expresses the economic time series in question as the sum of an international permanent component (IPC, denoted $i_{\pi t}$), an international temporary component (ITC, denoted $i_{\tau t}$), a regional component (RC, denoted r_{jt}), a first-order autoregressive, temporary country-specific component (CSC1, denoted c_{it}) and a purely temporary, zero-persistence country-specific component (CSC2, denoted θ_{it}). Thus, the dynamic multiple indicator-multiple cause or DYMIMIC model separates out the international, regional and country-specific components of the data in a fairly general, comprehensive fashion, where the stochastic process for p_{it} is allowed to be stationary or integrated.[5] The statistical treatment of the model outlined above is conveniently handled by writing it in state-space form (SSF), involving a set of measurement equations relating the unobserved components (the state vector) to observed series, together with a set of transition equations governing the evolution of the state vector. In the context of this paper, where we model each of real GDP and bank credit for $N = 4$ countries and $K = 2$ regions, the SSF corresponding to the model outlined in Eqs. (1)–(5) may be written as

$$\begin{bmatrix} p_{1t} \\ p_{2t} \\ p_{3t} \\ p_{4t} \end{bmatrix} = \begin{bmatrix} 1 & 1 & 1 & 0 & 1 & 0 & 0 & 0 \\ 1 & 1 & 1 & 0 & 0 & 1 & 0 & 0 \\ 1 & 1 & 0 & 1 & 0 & 0 & 1 & 0 \\ 1 & 1 & 0 & 1 & 0 & 0 & 0 & 1 \end{bmatrix} \begin{bmatrix} i_{\pi t} \\ i_{\tau t} \\ r_{1t} \\ r_{2t} \\ c_{1t} \\ c_{2t} \\ c_{3t} \\ c_{4t} \end{bmatrix} + \begin{bmatrix} \theta_{1t} \\ \theta_{2t} \\ \theta_{3t} \\ \theta_{4t} \end{bmatrix} \tag{6}$$

and

$$\begin{bmatrix} i_{\pi t} \\ i_{\tau t} \\ r_{1t} \\ r_{2t} \\ c_{1t} \\ c_{2t} \\ c_{3t} \\ c_{4t} \end{bmatrix} = \begin{bmatrix} 1 & 0 & \dots & \dots & \dots & \dots & 0 & 0 \\ 0 & \rho & 0 & \dots & \dots & \dots & \dots & 0 \\ 0 & 0 & \beta_1 & 0 & \dots & \dots & \dots & 0 \\ 0 & \dots & 0 & \beta_2 & 0 & \dots & \dots & 0 \\ 0 & \dots & \dots & 0 & \gamma_1 & 0 & \dots & 0 \\ 0 & \dots & \dots & \dots & 0 & \gamma_2 & 0 & 0 \\ 0 & \dots & \dots & \dots & \dots & 0 & \gamma_3 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & \gamma_4 \end{bmatrix} \begin{bmatrix} i_{\pi t-1} \\ i_{\tau t-1} \\ r_{1t-1} \\ r_{2t-1} \\ c_{1t-1} \\ c_{2t-1} \\ c_{3t-1} \\ c_{4t-1} \end{bmatrix} + \begin{bmatrix} \varepsilon_{\pi t} \\ \varepsilon_{\tau t} \\ \eta_{1t} \\ \eta_{2t} \\ u_{1t} \\ u_{2t} \\ u_{3t} \\ u_{4t} \end{bmatrix} \tag{7}$$

where Eq. (6) represents the measurement equations, which show how the observed series are related to the state vector, whereas Eq. (7) represents the transition equations, describing the dynamic evolution of the state vector. The covariance matrix of the vector of disturbances in Eq. (7), say Q, is a diagonal matrix with diagonal elements given by the hyperparameters $\{\sigma^2_{\varepsilon_\pi}, \sigma^2_{\varepsilon_\tau}, \sigma^2_{\eta_1}, \sigma^2_{\eta_2}, \sigma^2_{u_1}, \sigma^2_{u_2}, \sigma^2_{u_3}, \sigma^2_{u_4}\}$.

The state-space parameters can be estimated by maximum likelihood Kalman filtering methods (see Harvey, 1989; Gregory et al., 1997). The estimated hyperparameters or variance parameters indicate the relative contribution of each component in the state vector to explain the total variation in the time series under consideration. In some sense, therefore, the estimated variances allow us—by providing information on the relative sizes of the different (international, regional and country-specific) components in the series—to quantify the relative importance of each component in the model in explaining the variation in the economic time series under investigation. If the estimated variance of the international permanent component of real GDP is found to be relatively large and statistically significant, for example, then one may argue that global permanent shocks are the key driving shocks of real GDP. In contrast, if

[5] For simplicity and to minimize the number of parameters to be estimated, we employ autoregressive components of order one. However, we test the validity of this parsimonious autoregressive specification in our empirical work using a battery of test statistics for absence of residual serial correlation (see Section 4.2).

the estimated variance of the country-specific component is found to be relatively large and statistically significant, then real GDP would be driven largely by domestic shocks.

The modeling procedure is essentially a general-to-specific procedure where we start from the most general model (Eqs. (6) and (7)) and test down by imposing exclusion restrictions on the parameters found to be statistically insignificant at conventional nominal levels of significance. When the maximum likelihood estimate of the variance of an element of the state vector is zero, the model can be re-estimated making the corresponding component deterministic. Also, standard tests of the significance of the component itself can be carried out: if the component concerned is not found to be statistically significantly different from zero, the model may be simplified by eliminating the component from the SSF altogether. In choosing the most appropriate model for each country and sample period examined, we relied on standard goodness-of-fit measures such as the coefficient of determination, as well as on the Akaike information criterion.

We now turn to our empirical analysis, where we first apply the DYMIMIC model (Eqs. (6) and (7)) to each of real GDP and bank credit to the private sector in order to extract their international, regional and country-specific components. We then estimate VAR models designed to examine the dynamic interactions between cross-country components of bank credit and GDP growth.

4. Empirical results

4.1. Data

We employ quarterly time series for real GDP and real bank credit to the private sector for four major industrialized countries over the sample period from 1973:1 to 2001:4. The four countries examined are the US and Canada, comprising the North American region, and Germany and France, as representative of the European region. While the sample period coincidentally corresponds to the recent floating exchange rate regime following the collapse of the Bretton Woods system, it should be noted that this sample period was in fact dictated by data availability. The data were taken from the International Monetary Fund's *International Financial Statistics* (IFS). The two time series of interest in the empirical work below are real GDP (y) and real bank credit to the private sector (bc). Both y and bc were expressed in natural logarithms, demeaned and deseasonalized using standard dummy variables prior to beginning of the empirical analysis. This transformation is useful since it allows us to avoid introducing unobserved components specifically designed to capture deterministic seasonal variation in our DYMIMIC model, which would increase the number of parameters to estimate without adding much to our economic interpretation of the empirical results.[6]

4.2. Kalman filter results

In Tables 1 and 2 we report the results of estimation of the most appropriate DYMIMIC time series model in state-space form by the Kalman filter maximum likelihood method for real GDP

[6] Using augmented Dickey–Fuller tests to investigate the integration properties of the data, for each country and sample period considered we were unable to reject the null hypothesis of a unit root for each of y and bc but were able to do so for their first differences. The results from these unit root tests (not reported but available from the authors upon request) are consistent with the DYMIMIC models reported in the next sub-section, in that they suggest that each of y and bc are $I(1)$, for each country.

Table 1
Kalman filter results for real GDP

	US	Canada	Germany	France
Model	IPC, RC, CSC2	IPC, RC, CSC2	IPC, RC, CSC2	IPC, RC, CSC2
Q-ratios				
IPC	1.000	1.000	1.000	1.000
ITC	–	–	–	–
RC	0.627	0.627	0.685	0.685
CSC1	–	–	–	–
CSC2	0.283	0.244	0.286	0.339
Coefficients				
$\widehat{\rho}$	–	–	–	–
$\widehat{\beta}_1$	0.478 (0.052)	0.478 (0.052)	–	–
$\widehat{\beta}_2$	–	–	0.544 (0.091)	0.544 (0.091)
$\widehat{\gamma}_i \quad (i = 1, \ldots, 4)$	–	–	–	–
$\overline{R}^2$	0.830	0.826	0.784	0.802
LB(4)	{0.674}	{0.540}	{0.437}	{0.412}
LB(8)	{0.839}	{0.596}	{0.530}	{0.501}
LB(12)	{0.850}	{0.621}	{0.678}	{0.773}

Notes: we report the components included in the final estimation of the four-country DYMIMIC model for real GDP in the first row, for each country; IPC, RC and CSC2 denote the international permanent component, the regional component and the country-specific irregular component, respectively—namely $i_{\pi t}$, r_{jt} and θ_{it} in Eqs. (1)–(5). The Q-ratios are the ratios of the estimated standard deviation of each component to the largest estimated standard deviation across components for each country (real GDP series). $\widehat{\rho}$, $\widehat{\beta}_1$, $\widehat{\beta}_2$ and $\widehat{\gamma}_i \ (i = 1, \ldots, 4)$ denote the estimated coefficients for the AR(1) international temporary component, each of the two AR(1) regional components, and each of the four AR(1) country-specific components, obtained in the final estimation of the model. $\overline{R}^2$ and LB(m) denote the adjusted coefficient of determination and the p-value (in braces) for the Ljung–Box test statistic for absence of residual serial correlation up to order $m = 4$, 8, 12, respectively.

and bank credit, respectively, for each of the four countries examined. In the second row of these tables we report details of the unobserved components included in the estimated model, selected on the basis of the goodness-of-fit criteria discussed in Section 3. In the following six rows we report the estimated Q-ratios implied by the estimated standard deviations of the disturbances of the stochastic components included in the state vector; the Q-ratios are calculated as ratios of each estimated standard deviation to the largest standard deviation across components for each model, and indicate the relative statistical importance of the components (Harvey, 1989).[7] In the following five rows of Tables 1 and 2 we report the estimated coefficients of the components found to be statistically significant in the final estimation, which include the AR(1) coefficient (the damping factor) for each of the regional and country-specific components. These coefficients provide evidence on the degree of persistence of the AR(1) components of the model. In the last four rows, we report the adjusted coefficient of determination, $\overline{R}^2$ and the p-values from a battery of Ljung–Box test statistics of no residual serial correlation, LB.

[7] In terms of the SSF Eqs. (6) and (7) given in Section 3, the standard deviations are the square roots of the estimated diagonal elements of Q; the largest variance (standard deviation) is concentrated out of the likelihood function and, therefore, the Q-ratios are the ratios of each variance (standard deviation) to this variance (standard deviation).

Table 2
Kalman filter results for real bank credit

	US	Canada	Germany	France
Model	IPC, RC, CSC2	IPC, RC, CSC2	IPC, RC, CSC2	IPC, RC, CSC2
Q-ratios				
IPC	1.000	1.000	1.000	1.000
ITC	–	–	–	–
RC	0.728	0.728	0.581	0.581
CSC1	–	–	–	–
CSC2	0.465	0.410	0.110	0.139
Coefficients				
$\widehat{\rho}$	–	–	–	–
$\widehat{\beta}_1$	0.396 (0.041)	0.396 (0.041)	–	–
$\widehat{\beta}_2$	–	–	0.489 (0.094)	0.489 (0.094)
$\widehat{\gamma}_i \quad (i = 1, \ldots, 4)$	–	–	–	–
$\overline{R}^2$	0.802	0.773	0.850	0.831
LB(4)	{0.459}	{0.593}	{0.581}	{0.672}
LB(8)	{0.431}	{0.604}	{0.590}	{0.547}
LB(12)	{0.692}	{0.782}	{0.829}	{0.865}

Notes: we report the components included in the final estimation of the four-country DYMIMIC model for real GDP in the first row, for each country; IPC, RC and CSC2 denote the international permanent component, the regional component and the country-specific irregular component, respectively—namely $i_{\pi t}$, r_{jt} and θ_{it} in Eqs. (1)–(5). The Q-ratios are the ratios of the estimated standard deviation of each component to the largest estimated standard deviation across components for each country (real GDP series). $\widehat{\rho}$, $\widehat{\beta}_1$, $\widehat{\beta}_2$ and $\widehat{\gamma}_i \ (i = 1, \ldots, 4)$ denote the estimated coefficients for the AR(1) international temporary component, each of the two AR(1) regional components, and each of the four AR(1) country-specific components, obtained in the final estimation of the model. $\overline{R}^2$ and LB(m) denote the adjusted coefficient of determination and the p-value (in braces) for the Ljung–Box test statistic for absence of residual serial correlation up to order $m = 4$, 8, 12, respectively.

The estimation results are encouraging in that the coefficient of determination may be regarded as rather high in each case and the residuals appear to be uncorrelated.[8,9] As one might expect in the context of modeling real GDP (Table 1), a stochastic trend is important in characterizing the data. This stochastic trend is the international permanent component, common to all four countries under examination. If one is willing to interpret this component as driven largely by technology shocks, then it seems plausible to view this trend as common to all countries examined as an effect of technology diffusion. Hence, the differences in real GDP behavior across countries are due to differences in the adjustment process toward the stochastic trend, driven by the statistically significant regional and country-specific components. The latter two components are, in fact, found to be stationary, as implied by the values of the autoregressive parameters within the unit circle. While the international permanent component is always found to be the dominant component in the state vector, as clearly indicated by the Q-ratios, the regional and country-specific components do explain a sizeable proportion of real GDP fluctuations.

[8] We also carry out tests for no heteroskedasticity and the Jarque–Bera test statistic for normality (not reported to conserve space). In no case these statistics imply rejection of the hypotheses of no heteroskedasticity and normality.

[9] Also, note that the convergence achieved by the Broydon–Fletcher–Goldfarb–Shanno numerical optimization method is always very strong in the sense that it is achieved with a convergence criterion satisfied using a tolerance level of 1.0E–7.

This evidence seems consistent with the literature supporting the existence of cross-country links in macroeconomic fluctuations (Backus et al., 1995; Baxter, 1995; Gregory et al., 1997; Sarno, 2001; Kose et al., 2003) as well as with the studies that argue that business cycle comovements are stronger for subsets of countries or regions (e.g. Artis et al., 1997; Artis and Zhang, 1997; Bergman et al., 1998; Imbs, 2003; Lumsdaine and Prasad, 2003). Our results imply that, in addition to an international component that characterizes the long-run stochastic trend in real GDP for all four countries examined (Bernard and Durlauf, 1995), GDP fluctuations are also characterized by significant regional components (one for North America and one for Europe), in addition to country-specific components. The existence of significant regional components for each of North America and Europe is not surprising. On the one hand, it seems plausible that the US and Canada have business cycles that are highly synchronized. On the other hand, Germany and France, and presumably several other countries within Europe, have displayed increasing synchronization of their business cycles during the sample period under examination as an effect of their coordination of macroeconomic policy during the European Monetary System—e.g. see the relevant discussion in Artis et al. (1997) and Artis and Zhang (1997).[10] The AR(1) coefficients suggest that the regional cycles are stationary, but fairly persistent.

These results are well summarized in Fig. 1, which graphs, for each of the four countries examined, the actual real GDP data over the sample period together with the permanent component (the international permanent component) and the stationary component (the regional component plus the country-specific component). Fig. 1 makes clear both the slight dominance of the international permanent component over the stationary component in explaining the variation in real GDP as well as the persistence of the stationary component.

It is worth noting that all of the studies decomposing real activity measures into its international, regional and country-specific components cited in this paper have modeled GDP *growth* (which is stationary) rather than the *level* of GDP. Hence, the finding of an international permanent component in the level of real GDP is novel in this context. However, also note that we do not find a significant international transitory component in our model; this is the component that these studies tend to interpret as the world business cycle. In some sense, therefore, our results suggest that when one models real GDP in levels and allows for international permanent and temporary components, in addition to regional and country-specific components, regional business cycles tend to swamp the world business cycle.[11] Turning to our results for bank credit to the private sector (Table 2), we find

[10] The finding of a statistically and economically important European region is not trivial in this context, given that the German unification, which occurred in 1991, inevitably affected the behavior of the German business cycle. Indeed, simple correlation calculations suggest that the comovement between German and French output appears to be weaker in the 1990s than the output comovements between Italy and France, for example. Nevertheless, given Germany's leading economic role in Europe during the sample period under examination, we chose to use Germany in our model along with France.

[11] Since the primary focus of this paper is to assess empirically the existence of a financial accelerator across countries, rather than establishing the relative importance of world versus regional output fluctuations, we do not investigate this finding further and leave it for future research. However, we did re-estimate our DYMIMIC model without the international permanent component to check whether our finding of a negligible world business cycle could be due to the particular specification of the model; in this case, the autoregressive parameter in the international temporary component was always estimated to be very close to unity so that effectively the international component was picking up the permanent common component across the GDP series in the system. This simple robustness check made us more confident in the result that the most important international component is in fact permanent, rather than temporary.

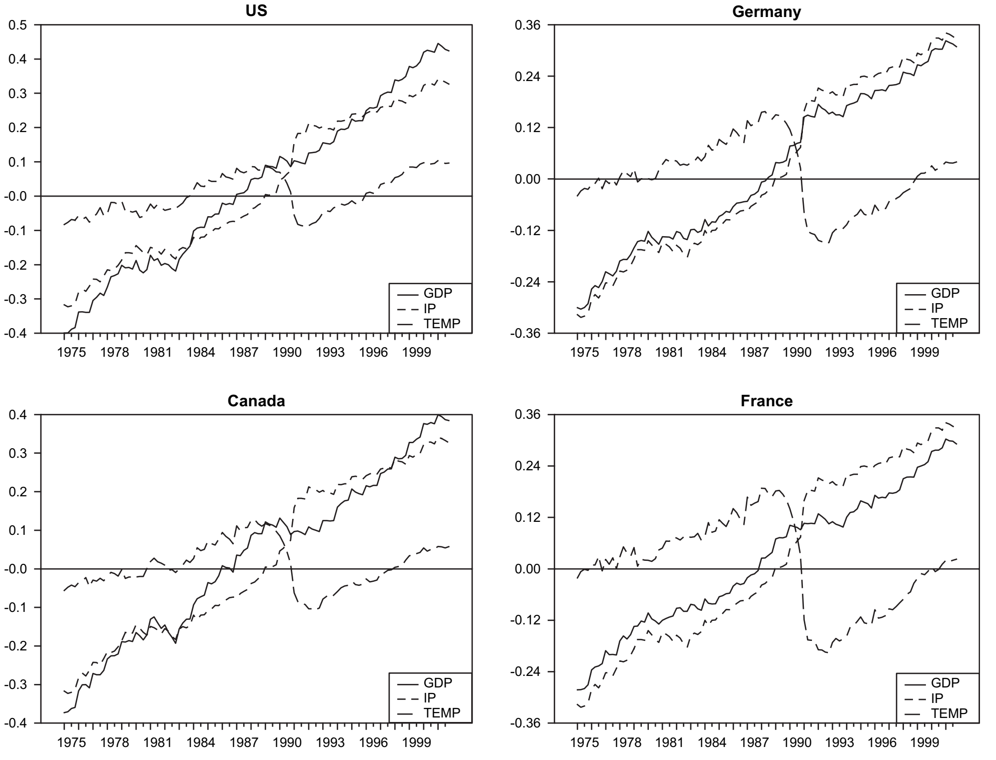

Fig. 1. Structural decomposition of real GDP.

that the same DYMIMIC model used for real GDP—including an international permanent component, a regional component and a country-specific component—is again our preferred specification. Also, the stationary (regional plus country-specific) component is about as important in explaining the variation of bank credit as the corresponding component was in explaining real GDP, with the international permanent component again slightly dominating the stationary component. The persistence of the stationary regional component also seems close to the persistence of the regional component for GDP, as evidenced by the estimated AR(1) coefficients in Table 2 and the graphs plotted in Fig. 2. However, it is interesting to note that, just like for real GDP, the Q-ratios of the regional component are generally larger than the Q-ratios of the country-specific components, and this evidence is particularly strong for the European region (Germany and France), where the contribution of the country-specific component explaining variation in real bank credit is indeed quite small. The latter result is not surprising if one considers the high degree of coordination of monetary policy and interest rate setting that has characterized the second part of our sample in these economies. Indeed, in the last part of the sample, the low values of the regional component for the European region may be seen as suggesting that the birth of the euro may have induced a credit crunch in this region. Finally, the existence of an international permanent component in bank credit suggests that the stock of bank credit may be moved by common permanent shocks so that the stock of bank credit of the four countries examined essentially follows the same stochastic trend.

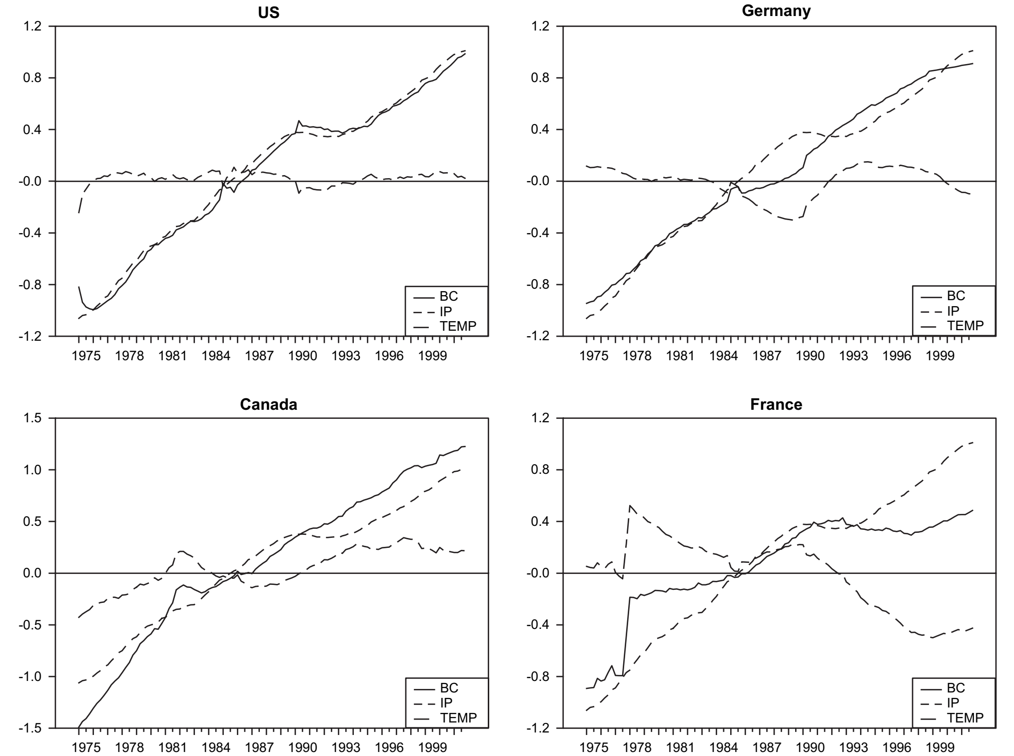

Fig. 2. Structural decomposition of bank credit.

4.3. VAR estimation results

Having analyzed international, regional and country-specific components of both real GDP and bank credit, a logical question arising in light of our results in the previous sub-section concerns whether bank credit affects an economy's real GDP growth, consistent with the financial accelerator mechanism. In particular, given our finding that bank credit displays a statistically important regional component, it is tempting to conjecture that this regional component may be important in driving GDP growth, independently from and in addition to the country-specific component.

To this end, we estimated a trivariate VAR for real GDP and each of the regional and country-specific components of bank credit, for each country examined.[12] On the basis of this VAR, we then carry out causality tests and impulse response analysis to investigate the statistical and economic importance of the regional and country-specific components of bank credit for GDP growth. We use four lags of each variable and order GDP growth first, region-specific growth in bank credit next, and country-specific growth in bank credit last.[13] In Table 3, we report Granger-causality tests for each equation of the VAR and for each country. These results

[12] We chose to focus on GDP growth rather than other measures of business cycles for consistency with the previous literature in this context (e.g. Bernanke et al., 1999).

[13] Hence we assume that movements in GDP may have a contemporaneous effect on bank credit but not vice versa.

Table 3
Causality tests

	Test on	US	Canada	Germany	France
Equation for Δy					
	Δy	0.043	0.034	0.041	0.002
	Δbc^r	0.005	0.001	0.043	0.033
	Δbc^{cs}	0.001	0.003	0.045	0.036
Equation for Δbc^r					
	Δy	0.653	0.084	0.689	0.433
	Δbc^r	0.000	0.000	0.000	0.000
	Δbc^{cs}	0.246	0.735	0.124	0.237
Equation for Δbc^{cs}					
	Δy	0.929	0.399	0.614	0.312
	Δbc^r	0.115	0.301	0.081	0.943
	Δbc^{cs}	0.000	0.000	0.000	0.000

Notes: we report the test results from executing Granger-causality tests in a trivariate VAR comprising real GDP growth (Δy), the regional component of bank credit (Δbc^r) and the country-specific component of bank credit (Δbc^{cs}); the VAR is estimated separately for each of the US, Canada, Germany and France. For each country and for each of the three equations in the VAR, we report the causality results for the null hypothesis that each of the three variables in the VAR Granger causes the left hand side variable, hence testing for causality in all possible directions. All figures reported are p-values for the null hypothesis of no Granger causality.

suggest a very clear pattern: in the equations for regional and country-specific bank credit, only the lagged dependent variables have explanatory power, implying that GDP growth does not Granger-cause bank credit. On the other hand, for each country, the null hypothesis of non-causality is rejected at conventional nominal significance levels for each of regional and country-specific bank credit, suggesting unidirectional Granger causality from bank credit to GDP growth.

Fig. 3 reports the effect of a one-standard deviation increase in each of the bank credit variables in the VAR on GDP growth. As the figure clearly shows, an unanticipated one standard deviation rise in the growth of regional or country-specific bank credit leads to a significant and persistent increase in GDP growth, both statistically and economically. Our interpretation is that the link between regional and country-specific bank credit is consistent with the financial accelerator mechanism, which operates not only through the conventional country-specific channel highlighted in previous research, but also through the regional channel. In turn, this regional channel may be viewed as one of the mechanisms inducing correlations in output fluctuations across countries within the same region, i.e. North America and Europe.

It is interesting, however, that the impact of a bank credit shock is much more persistent for the European region than for the North American region. This seems intuitively reasonable, since financial market frictions are perhaps more substantial in Europe than in North America. The impulse responses for Germany and France display, in fact, a marked degree of similarity and indicate that the role of the financial accelerator is not only statistically but also economically very significant. On the other hand, the role of the financial accelerator in the US and Canada appears relatively smaller (Allen and Gale, 2000). In particular, for the US and Canada, we observe a large initial response, followed by a steady decay, such that the effect dies out by the end of seven quarters. In contrast, for the two European countries, we see an initial rise that gets stronger and then there is a much more gradual decay so that the effect persists for some 15 quarters. What may explain the differences? It is possible that because relationship banking is

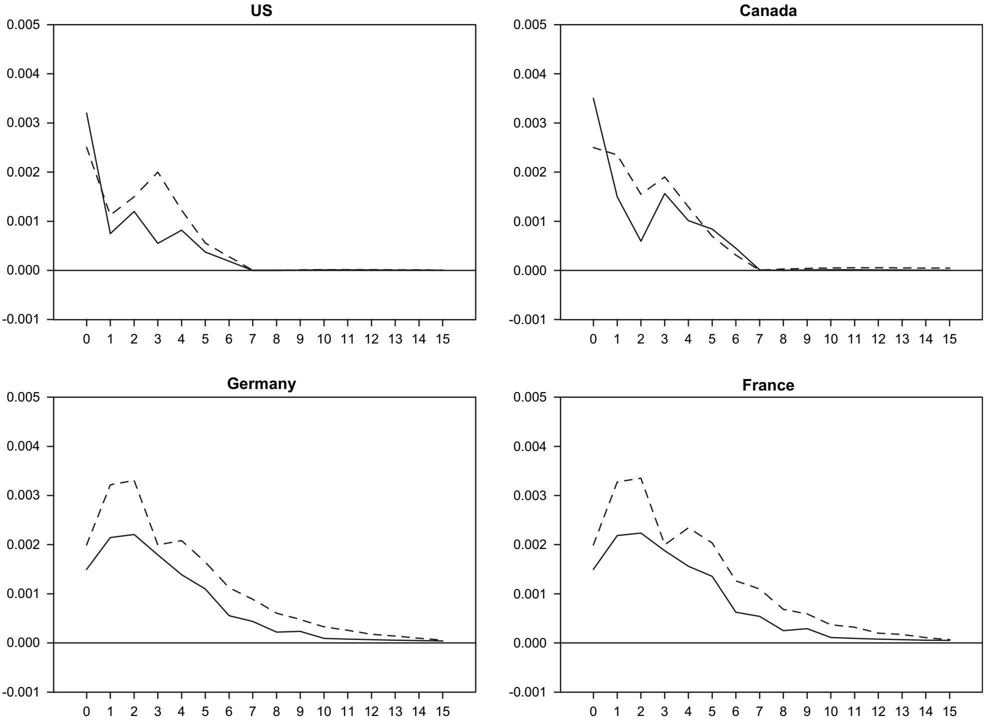

Fig. 3. Impulse response of GDP growth. Solid (broken) line=regional (country-specific) bank credit shock.

more important in Europe, the shocks take time to materialize fully and, hence the ramping up in the accelerator effect after the initial shock. The same persistence may explain the slower decay. In the US and Canada, capital markets mediate more of the funds and so the effect is felt sooner but fades out faster.

4.4. Caveats and extensions

One caveat to our empirical work is that it is based on a specific proxy for credit-market imperfections, namely real bank credit to the private sector. While this proxy makes intuitive sense and has the advantage of being available for a long span of data at the quarterly frequency, it has the shortcoming that bank credit could fall, for example, in response to a drop in the demand for credit from the private sector. In this case, clearly the fall in our proxy would not indicate anything about credit rationing, confounding the interpretation of our empirical results.

A potentially better proxy for credit conditions may be a measure of credit spreads—above government rates. Indeed, we constructed such a measure using Merrill Lynch data.[14] However, these data are only available for some of the countries examined here since January 1998,

[14] Specifically, we employed the US High-Yield Master II Index (ticker H0A0), the Canadian High-Yield Index (ticker HC00), and for the Eurozone countries we employed the Euro High-Yield Index (ticker HE00). The data were taken from the Merril Lynch *Bond Index Database*.

which prevents us from estimating the DYMIMIC model with a sufficient degree of accuracy due to the small number of observations and, perhaps more importantly, short sample span. However, it is worth noting that the correlation between these measures based on credit spreads and real bank credit ranges between 0.82 for the US and 0.87 for Canada over the four years of data that we can compute such correlation. This is comforting since, insofar as measures based on credit spreads move less than bank credit in response to changes in credit demand, the high correlation is indicative of the fact that the variation of real bank credit which is due to these changes is likely to be relatively small.[15]

While we have attempted to establish baseline results on the existence of a cross-country financial accelerator and its relevance for fluctuations in real activity, a number of extensions of this research appear worth investigating. In particular, the effect through which the financial accelerator mechanism operates is likely to be asymmetric in at least two ways: stronger in downturns than in upturns (since a 'credit crunch' will exacerbate the effects of a downturn), and stronger for small firms than for large firms (since small borrowers are more likely to encounter credit constraints). The existence and relevance of these asymmetries might perhaps be investigated by generalizing our time series models to allow for different effects of negative and positive shocks and by building cross-country data sets at the firm level, respectively.

Another promising avenue for future research involves examining data for emerging economies, where credit rationing may be quantitatively more important and the impact of the financial accelerator more relevant. In particular, examination of cross-country regional components in East Asia and Latin America might be an interesting case study to validate or reject the role of regional components in credit cycles in explaining real activity. The generality of our results would benefit particularly from such an analysis of a larger set of countries and regions. However, there are at least two problems in moving the research in this direction. First, data gathering for proxies of credit-market imperfections becomes much more cumbersome when one moves away from major industrialized economies. Second, estimation of the resulting DYMIMIC model becomes very difficult, due to the large number of hyperparameters for which joint convergence is required. We would suspect that estimation of larger models would require moving away from classical estimation methods and developing estimation methods based on Gibbs sampling specific for this class of models. While neither of these problems is potentially insuperable, we leave this avenue open for future research.

5. Conclusion

In this paper we have examined the empirical relevance of the financial accelerator in a cross-country context. In particular, we have investigated the extent to which cross-country components of credit cycles may be important for explaining output fluctuations. Using Kalman filtering maximum likelihood techniques applied to data for the US, Canada, Germany and France over the 1973–2001 period, we could identify, for both real GDP and an indicator of

[15] We stress that the latter statement should be taken with caution given that four years of data are indeed a very short span and unlikely to enable us to characterize the time series properties we are interested in, in terms of credit-market imperfections over time. We would argue that having a long span of data is crucial for the analysis of this paper, since it gives the opportunity to observe the relationship between credit-market imperfections and real activity over several business cycles. Indeed, this is the key reason why we rely on real bank credit, regardless of its shortcomings, in our empirical work.

bank credit activity, an international permanent component, which is common to all countries examined, a stationary regional component and a stationary country-specific component.

We then examined the dynamic interactions linking the regional component and the country-specific component of bank credit and movements in real activity using multivariate vector autoregressive modeling and impulse response analysis. Estimation of a trivariate vector autoregression for GDP growth, the regional component of bank credit and the country-specific component of bank credit, reveals that both regional and country-specific components of bank credit are empirically important determinants of GDP growth for each of the four countries examined. In particular, we find strong evidence of unidirectional causality from regional and country-specific bank credit to GDP growth. The impulse response functions of GDP growth in response to bank credit shocks implied by our models may be viewed as consistent with the financial accelerator mechanism, suggesting that the existence of regional credit cycles may be responsible for regional comovements in output fluctuations. Overall, therefore, this research provides the first empirical evidence of the presence of a cross-country financial accelerator.

Acknowledgments

This paper was partly written while Lucio Sarno and Mark Taylor were Visiting Scholars at the International Monetary Fund, whose hospitality is gratefully acknowledged. The authors are grateful for conversations or comments on earlier drafts to Mike Melvin (editor), an anonymous referee, Robin Brooks, Hali Edison, Bob Flood, Ayhan Kose, Eswar Prasad, Alessandro Rebucci, Carmen Reinhart and Antonio Spilimbergo. The authors alone are responsible for the views expressed and for any errors that may remain.

References

Allen, F., Gale, D., 2000. Comparing Financial Systems. MIT Press, Cambridge, MA and London, UK.

Artis, M.J., Kontolemis, Z.G., Osborn, D.R., 1997. Business cycles for G7 and European countries. Journal of Business 70, 249–279.

Artis, M.J., Zhang, W., 1997. International business cycles and the ERM: is there a European business cycle? International Journal of Finance and Economics 2, 1–16.

Backus, D.K., Kehoe, P.J., Kydland, F.E., 1995. International real business cycles. Journal of Political Economy 100, 745–775.

Baxter, M., 1995. International trade and business cycles. In: Grossman, G., Rogoff, K. (Eds.), Handbook of International Economics, vol. III. North-Holland, Amsterdam, pp. 1801–1864.

Bergman, U.M., Bordo, M.D., Jonung, L., 1998. Historical evidence on business cycles: the international experience. In: Fuhrer, J.C., Schuh, S. (Eds.), Beyond Shocks: What Causes Business Cycles? Conference Series No. 42 Federal Reserve Bank of Boston, Boston, pp. 65–113.

Bernanke, B., Gertler, M., Gilchrist, S., 1996. The financial accelerator and the flight to quality. Review of Economics and Statistics 78, 1–15.

Bernanke, B., Gertler, M., Gilchrist, S., 1999. The financial accelerator in a quantitative business cycle framework. In: Taylor, J., Woodford, M. (Eds.), Handbook of Macroeconomics, vol. 1C. Elsevier Science, North-Holland, Amsterdam, New York and Oxford, pp. 1341–1393.

Bernard, A.W., Durlauf, S.N., 1995. Convergence of international output. Journal of Applied Econometrics 10, 97–108.

Dotsey, M., 1998. The predictive content of the interest rate term spread for future economic growth. Federal Reserve Bank of Richmond Economic Quarterly 84, 31–51.

Engle, R.F., Watson, M.F., 1981. A one-factor multivariate time-series model of metropolitan wage rates. Journal of the American Statistical Association 76, 774–781.

Gertler, M., Lown, C.S., 1999. The information in the high-yield bond spread for the business cycle: evidence and some implications. Oxford Review of Economic Policy 15, 132–150.

Gregory, A.W., Head, A.C., Raynauld, J., 1997. Measuring world business cycles. International Economic Review 38, 677–701.

Harvey, A.C., 1989. Forecasting, Structural Time Series Models and the Kalman Filter. Cambridge University Press, Cambridge.

Hubbard, R.G., 1998. Capital market imperfections and investment. Journal of Economic Literature 36, 193–225.

Kose, M.A., Otrock, C., Whiteman, C.H., 2003. International business cycles: world, region and country specific factors. American Economic Review 93, 1216–1239.

Imbs, J., 2003. Co-fluctuations. Centre for Economic Policy Research Discussion Paper, No. 2267; London Business School, mimeo.

Lumsdaine, R.L., Prasad, E.S., 2003. Identifying the common component in international economic fluctuations: a new approach. Economic Journal 113, 101–127.

Mody, A., Taylor, M.P., 2003. The high yield spread as a predictor of real economic activity: evidence of a financial accelerator for the United States. International Monetary Fund Staff Papers 50, 373–402.

Sarno, L., 2001. Nonlinear dynamics, spillovers and growth in the G7 economies: an empirical investigation. Economica 68, 401–426.