UNDERSTANDING CHANGES IN INTERNATIONAL BUSINESS CYCLE DYNAMICS

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Abstract
The volatility of economic activity in most G7 economies has moderated over the past 40 years. Also, despite large increases in trade and openness, G7 business cycles have not become more synchronized. After documenting these facts, we interpret G7 output data using a structural VAR that separately identifies common international shocks, the domestic effects of spillovers from foreign idiosyncratic shocks, and the effects of domestic idiosyncratic shocks. This analysis suggests that, with the exception of Japan, a significant portion of the widespread reduction in volatility is associated with a reduction in the magnitude of the common international shocks. Had the common international shocks in the 1980s and 1990s been as large as they were in the 1960s and 1970s, G7 business cycles would have been substantially more volatile and more highly synchronized than they actually were. (JEL: C3, E5)

1. Introduction

During the past two decades, most of the G7 economies have experienced a reduction in the volatility of output growth and a concomitant moderation of business cycle fluctuations. Table 1 presents standard deviations of four-quarter growth rates of per capita GDP in the G7 countries during each of the past four decades. Germany, Italy, Japan, the UK, and the US all experienced large reductions in volatility. Over this period, international trade flows have increased substantially, financial markets in developed economies have become increasingly integrated, and continental European countries moved to a single currency. These developments raise the possibility of changes not only in the severity of international business cycles, but also in their synchronization.

There already is a large body of research on these changes, and there is agreement on many of the basic facts. As initially pointed out by Kim and Nelson (1999) and McConnell and Perez-Quiros (2000), there has been a substantial

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TABLE 1. Standard deviation of four-quarter percentage growth of per capita GDP in the G7 by decade.

	1960–1969	1970–1979	1980–1989	1990–2002
Canada	1.83	1.82	2.67	2.24
France	1.24	1.66	1.27	1.43
Germany	2.56	2.13	1.67	1.53
Italy	2.34	3.14	1.33	1.30
Japan	2.19	3.16	1.57	2.08
UK	1.84	2.48	2.51	1.60
US	2.09	2.74	2.66	1.47

Notes: Entries are the standard deviation of 100 $\ln(GDP_t/GDP_{t-4})$.

moderation in output fluctuations in the US, with these and most other authors suggesting that this moderation is well modeled as a single break in the mid-1980s. Some of the proposed explanations of the U.S. moderation, such as changes in monetary policy and adoption of new inventory management methods, are domestic in origin, while others, such as smaller international shocks or stabilizing effects of trade, have international roots; for further discussion and references see Blanchard and Simon (2001), Stock and Watson (2002a), and Ahmed, Levin, and Wilson (2004).

Although the moderation of volatility is also evident in international data, when modeled as a single break the reductions generally are neither concurrent nor of similar magnitudes (e.g., Mills and Wang 2000; Simon 2001; Dalsgaard, Elmeskov, and Park 2002; van Dijk, Osborn, and Sensier 2002; Doyle and Faust 2002; Del Negro and Otrok 2003; Fritsche and Kouzine 2003). Moreover, existing research suggests little tendency towards increasing international synchronization of cyclical fluctuations (Doyle and Faust 2002, 2005; Kose, Prasad, and Terrones 2003; Heathcote and Perri 2004). Instead, there appears to have been an emergence of at least one cyclically coherent group, the major countries in the Euro-zone (Artis, Kontelemis, and Osborn 1997; Artis and Zhang 1997, 1999; Carvalho and Harvey 2002; Dalsgaard, Elmeskov and Park 2002; Helbling and Bayoumi 2003; Del Negro and Otrok 2003; Luginbuhl and Koopman 2003), and possibly a second, English-speaking group, consisting of Canada, the UK, and the US (Helbling and Bayoumi 2003).

This paper has two specific objectives. The first is to provide a concise summary of the empirical facts about the moderation in output volatility, changes in persistence, and changes in cyclical comovements for the G7 countries. One conclusion is that the single- break model of variance reduction, which fits the US well, does not adequately describe the international patterns of moderation. In addition, we provide further evidence of the emergence of two cyclically coherent groups, the Euro-zone and English-speaking countries.

Our second objective is to provide quantitative estimates of the sources of these changes. Are they domestic or international in origin? Do they reflect changes in the magnitudes of structural shocks or, rather, changes in the response of the economies to those shocks? To obtain these estimates, we use a so-called factor-structural vector autoregression (FSVAR), specified in terms of the growth rates of quarterly GDP in the G7 countries. This FSVAR is a conventional structural VAR, where the identifying restrictions come from imposing an unobserved-component factor structure on the VAR innovations. The idiosyncratic shocks are allowed to affect future output in other countries, so this FSVAR makes it possible to quantify both the direct effect of common international shocks and the indirect effect of spillovers from the domestic shocks in one country to its trading partners. The FSVAR is overidentified, and tests of the overidentifying restrictions suggest that the G7 output data are well described as being driven by two common international shocks, plus seven country-specific shocks. This FSVAR makes it possible to address various counterfactual questions, and (for example) facilitates estimating the extent to which the moderation in volatility is a result of smaller common international shocks, is domestic in origin, or is the result of a moderation in the US that spills over into the other G7 countries.

The data and methods we use to remove trends and to isolate business cycle components are briefly described in Section 2. Section 3 summarizes the empirical facts about changes in volatility and persistence for the individual G7 output data, and Section 4 summarizes the changes in international correlations. The FSVAR is described, and its overidentifying restrictions are tested, in Section 5. Empirical results and counterfactual calculations based on the FSVAR are presented in Section 6. Section 7 concludes.

2. Data and Filters

The data are quarterly values of the logarithm of per capita real GDP for the G7 countries (Canada, France, Germany, Italy, Japan, UK, and US), covering 1960:1–2002:4. The data are described in detail in the Appendix.

Our focus is on economic fluctuations over the horizons relevant for medium-term macroeconomic policy and over business cycle horizons. Accordingly, we consider transformations of the data that filter out the highest frequency, quarter-to-quarter fluctuations. One way to do this is to use band-pass-filtered log GDP, with a pass band that focuses on business cycle frequencies (periods of 6 to 32 quarters). An alternative is to consider four-quarter growth rates, which use differencing to eliminate the linear growth rate in the series and four-quarter averaging to eliminate high-frequency noise. Finally, as has proved useful in VAR analysis, forecast errors at different forecasts horizons can be used to study behavior at different frequencies. These methods are complementary and all three will be used in this paper.

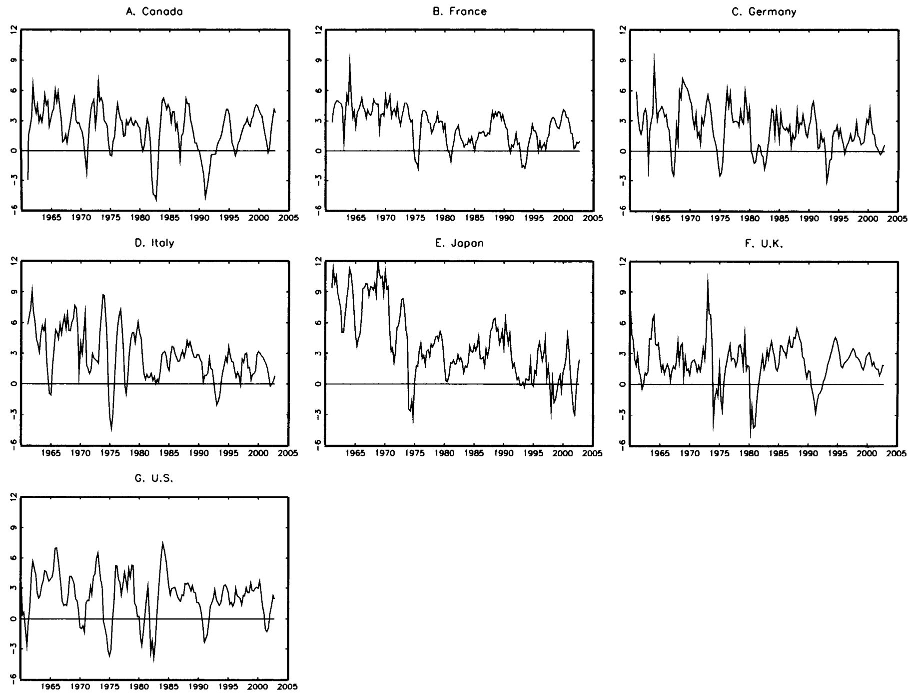

FIGURE 1. Four-quarter growth rates of GDP.

Figure 1 plots the four-quarter growth rate of per capita GDP for each country. The long- term growth rate of GDP is not constant for some of these countries, especially Germany, Japan, and Italy. The focus of this paper is fluctuations at yearly through business cycle horizons, not the determinants of early postwar trend growth in Germany, Japan, and Italy. Because a low-frequency drift can introduce bias into certain statistics, such as cross-country correlations computed over long subsamples, in some of our analysis we will use detrended versions of growth rates, where we use a flexible detrending method based on a model with a stochastic drift. Let $y_t = 400\Delta \ln(GDP_t)$ be the quarterly growth of GDP at an annual rate. We adopt an unobserved components specification that represents y_t as the sum of two terms, a slowly evolving mean growth rate and a stationary component

$$y_t = \mu_t + u_t, \text{ where } \mu_t = \mu_{t-1} + \eta_t \tag{1}$$

and $a(L)u_t = \varepsilon_t$, where L is the lag operator and ε_t and η_t are serially and mutually uncorrelated mean zero disturbances. The Kalman smoother can be used to estimate the local mean, μ_t, and the residual. The detrended GDP growth rate is the residual, that is, the Kalman smoother estimate of u_t.

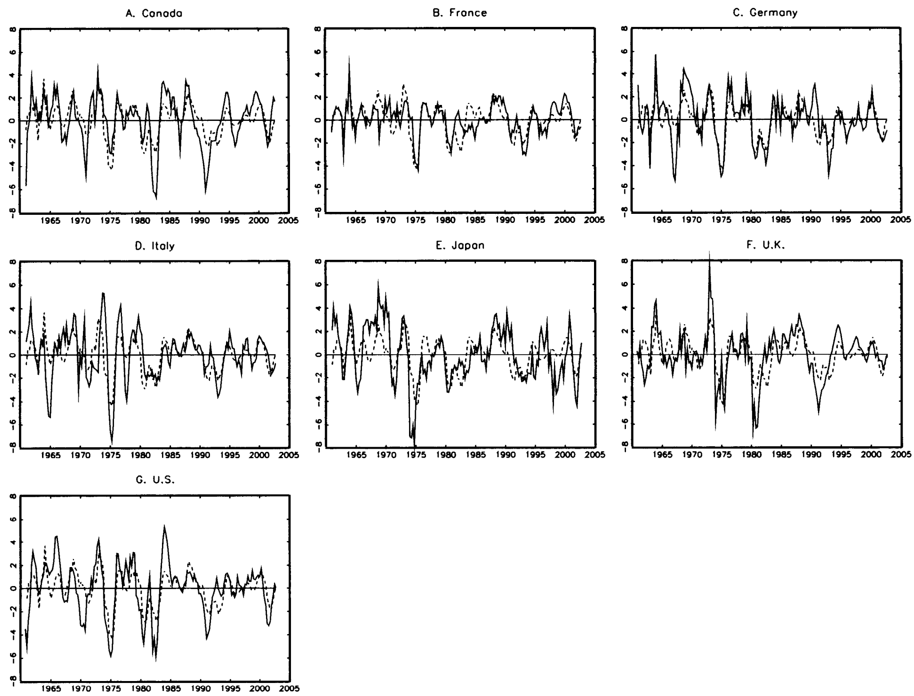

FIGURE 2. Detrended four-quarter GDP growth: Individual countries (solid lines) and G7 average (dashed line).

Implementing this detrending procedure requires a value of the ratio $\sigma_\eta^2/S_{uu}(0)$, where $S_{uu}(0)$ is the spectral density of u_t at frequency zero. When $\sigma_\eta^2/S_{uu}(0)$, is small, as it plausibly is here, the maximum likelihood estimator of $\sigma_\eta^2/S_{uu}(0)$ has the "pileup" problem of having asymptotic point mass at zero even if its true value is nonzero but small, so we estimate $\sigma_\eta^2/S_{uu}(0)$ on a country-by-country basis using the median-unbiased estimator of Stock and Watson (1998), and use this country-specific estimate to detrend GDP growth.[1]

Figure 2 plots the detrended four-quarter growth rates, that is, the rolling four-quarter average of the detrended quarterly growth rates. Comparison of Figures 1 and 2 reveals that the detrending procedure eliminates the local mean of each series, but otherwise leaves the series essentially unchanged. Figure 2 also plots the G7-wide unweighted average detrended four-quarter growth rate. Evidently

1. The median-unbiased estimators of $[T^2\sigma_\eta^2/s_{uu}(0)]^{1/2}$ were computed by inverting the point optimal invariant statistic with local parameter 7; see Stock and Watson (1998) for details. The estimates are: Canada, 6.4; France, 9.3; Germany, 3.3; Italy, 8.9; Japan, 6.2; UK, 0.0; and US, 3.1. For each country, $a(L)$ has degree 4.

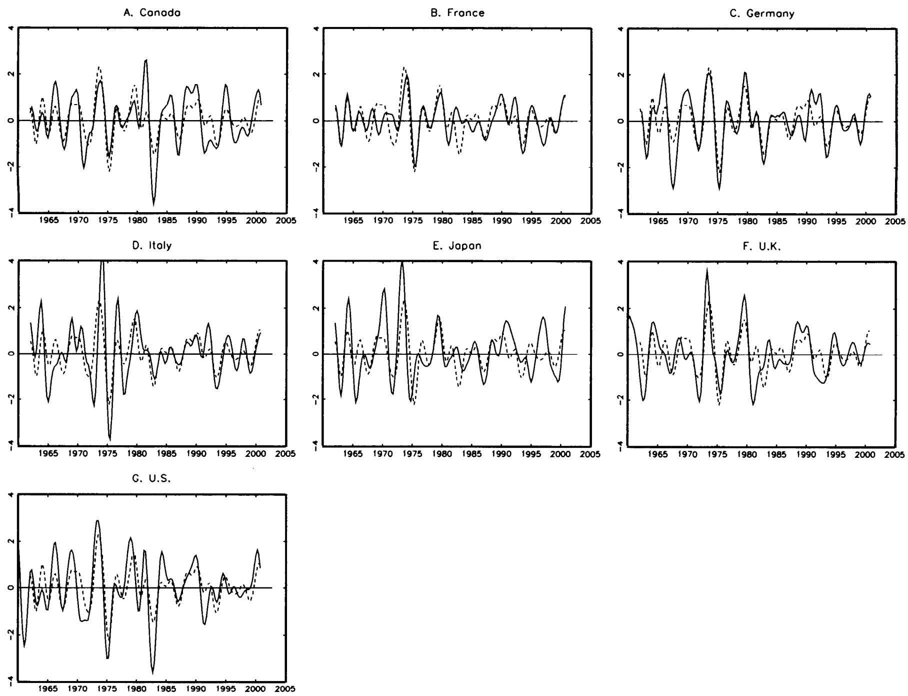

FIGURE 3. Band-pass-filtered GDP growth: Individual countries (solid lines) and G7 average (dashed line).

many of these countries have episodes of considerable comovement, or synchronization, with aggregate G7 fluctuations.

Figure 3 plots the band-pass-filtered logarithm of GDP along with the average of the BP- filtered G7 GDP.[2] Evidently BP-filtered GDP, like four-quarter growth, has periods of considerable international synchronization in business cycles. Notably, at the level of detail of Figures 2 and 3, the period of greatest synchronization appears to be the 1970s, and there is no readily apparent trend towards increased synchronization.

3. Changes in Volatility and Persistence

This section presents statistics summarizing changes in the volatility of GDP and in the persistence of innovations to GDP in the G7 countries.

2. We use the Baxter–King (1999) band-pass (BP) filter, with eight leads and lags and a pass-band of 6 to 32 quarters.

3.1. Volatility

As discussed in the introduction, there has been a substantial moderation in the volatility of economic activity over the past 40 years. To get more detail on this moderation, we estimate the time path of the instantaneous variance of GDP using a non-Gaussian smoother based on a stochastic volatility model with heavy tails and time-varying autoregressive coefficients. Let y_t be the quarterly GDP growth at an annual rate. The stochastic volatility model is

$$y_t = \alpha_{0t} + \sum_{j=1}^{p} \alpha_{jt} y_{t-j} + \sigma_t \varepsilon_t, \text{ where}$$

$$\alpha_{jt} = \alpha_{jt-1} + c\eta_{jt} \text{ and } \ln \sigma_t^2 = \ln \sigma_{t-1}^2 + \zeta_t, \tag{2}$$

where $\varepsilon_t, \eta_{1t}, \ldots, \eta_{pt}$ are i.i.d. $N(0, 1)$ and where ζ_t is distributed independently of the other shocks. To allow for large jumps in the instantaneous innovation variance, ζ_t is drawn from a mixture-of-normals distribution. The time-varying parameters were estimated by Markov Chain Monte Carlo methods. Given $\alpha_{0t}, \ldots, \alpha_{pt}$, and σ_t^2, it is possible to compute the instantaneous standard deviations of GDP growth, of four-quarter GDP growth, and BP-filtered GDP for an idealized BP filter. To facilitate comparisons of the results across countries, the same values of the hyperparameters were used for each country. For details (including the values of the hyperparameters), see Stock and Watson (2002a, Appendix A).

The resulting estimated instantaneous standard deviation of four-quarter GDP growth is plotted in Figure 4. Different countries exhibit quite different paths of instantaneous standard deviations. In the US, there was a sharp moderation in the mid-1980s, while in the UK, volatility declined in the 1970s. Germany experienced a large but gradual decline in volatility, while volatility moderated in Japan but has increased recently.[3]

Formal tests for breaks in the conditional mean (that is, the autoregressive lag coefficients) and the conditional variance (that is, the autoregressive innovation variance) of GDP growth are reported in Table 2.[4] The hypothesis of constant parameters is tested using the Wald version of the Quandt likelihood ratio (QLR) statistic, evaluated over the central 70% of the sample; the test of a constant conditional variance allows for the possibility of a break in the conditional mean at an unknown date that differs from the break date for the conditional variance. The break date and its 67% confidence interval are reported if the QLR statistic

3. Nearly identical patterns emerge when the model (2) is used to estimate the instantaneous variance of bandpass-filtered GDP instead of four-quarter GDP growth (results for BP-filtered GDP are not presented to save space).

4. Raw (i.e., not detrended) GDP growth rates are used in Table 2 to coincide more closely to the distribution theory underlying these statistics (Andrews 1991 and Bai 1997).

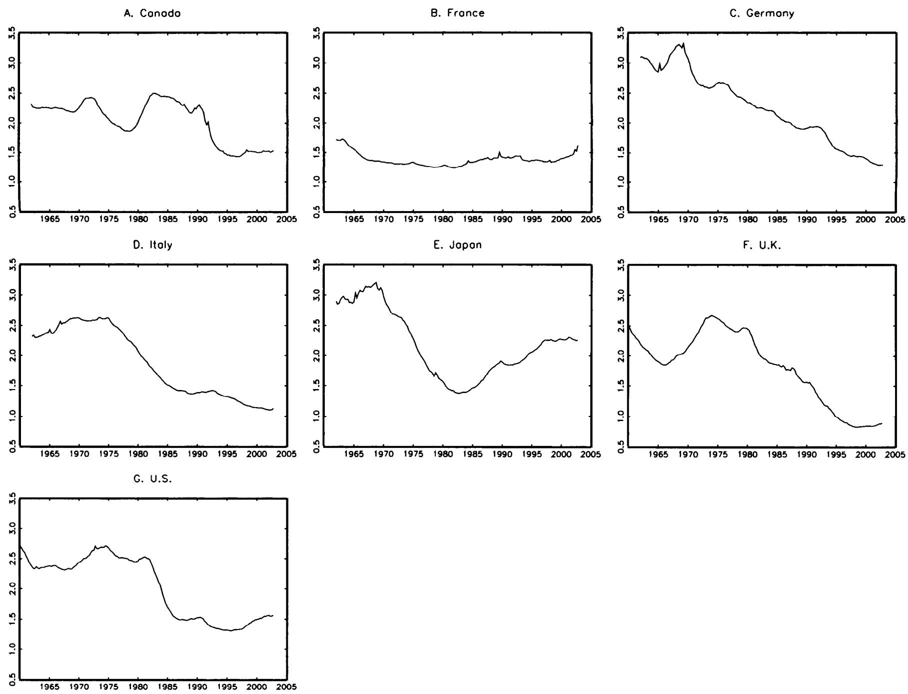

FIGURE 4. Estimated instantaneous standard deviation of four-quarter GDP growth.

rejects at the 5% significance level. The final block of Table 2 tests an alternative specification in which the innovation variance is modeled as a linear function of time with a discrete jump at an unknown break date, thereby nesting the single-break and linear time-trend specifications.

The results in Table 2 indicate widespread instability in both the conditional mean and the conditional variance of these autoregressive models for GDP: In five of the seven countries, the hypothesis of a constant conditional mean is rejected at the 5% level, and in all countries but Japan the hypothesis of a constant conditional variance is rejected. For the US, the results in the final block suggest that the break model is preferred to a linear time trend: in the nested specification, the break is significant but the time trend is not.

This finding does not generalize to the other countries, however. For example, for Germany neither the trend term nor the break term are individually significant in the nested specification. This finding does not imply that the variance for Germany was constant, for the test in panel B rejects the no-break specification at the 1% level and the estimated instantaneous variances in Figure 4 indicates a substantial reduction in volatility over this period; rather, the nonrejections for Germany and Japan—and the significance of both terms for the UK—suggests that neither the single-break nor the linear-decline model provides a good summary of

TABLE 2. Test for breaks in autoregressive parameters.

	Conditional mean			Conditional variance: break model			Conditional variance: Trend + break model		
	p	Break date	67% confidence interval	p	Break date	67% confidence interval	p: trend	p: break	Break date
Canada	0.00	1972:4	1972:2–1973:2	0.00	1991:2	1990:4–1993:1	0.00	0.25	
France	0.00	1974:1	1973:3–1974:3	0.03	1968:1	1967:3–1970:3	0.92	0.65	
Germany	0.39			0.00	1993:1	1992:3–1995:2	0.49	0.91	
Italy	0.00	1979:4	1979:2–1980:2	0.00	1980:1	1979:3–1982:4	0.00	0.27	
Japan	0.00	1973:1	1972:3–1973:3	0.32			0.34	0.11	
UK	0.00	1980:1	1979:3–1980:3	0.00	1980:1	1979:4–1982:1	0.00	0.00	1970:4
US	0.99			0.00	1983:2	1982:4–1985:3	0.67	0.01	1983:2

Notes: These results are based on AR(4) models estimated using $\Delta\ln(GDP_t/GDP_{t-1})$. The results shown in the columns labels "Conditional Mean" refer to changes in the AR coefficients, and the results shown in the columns labeled "Conditional Variance" refer to changes in the variance of the AR innovations. The "break model" allows a one-time break in the variance; the "Trend + break model" allows the variance to contain a linear trend and a one-time break. Columns labeled " p" are the p-value of the test statistic under the null hypothesis of no-change; "Break date" is the estimated date of a one-time shift in the parameters (reported only if the p-value is less than 5%); and the confidence interval is for the break date.

the changing volatility for these countries. Although the nested tests in the final block of Table 2 point towards the trend model for Canada and Italy, the estimates in Figure 4 look more like a series of plateaus than a linear trend. Taken together, we interpret all this as evidence that the pattern of the change in GDP volatility for most G7 countries is more complex than the single-break model that describes the US.

3.2. Persistence and Size of Univariate Shocks

Another way to look at the changing autocovariances of GDP growth in these countries is to examine changes in the variance of the AR innovation and in the sum of the AR coefficients, which measures the persistence of a shock to GDP growth. Changes in the variance of GDP growth imply that its spectrum has changed; an increase in the sum of the AR coefficients implies an increase in the relative mass at frequency zero, while a change in the innovation variance implies a shift in the level (but not necessarily the shape) of the spectrum.

We use two methods to capture time variation in the AR. The first method allows for a discrete break in 1984. Although a break in 1984 describes the U.S. data well, variation in the other countries is more subtle, so this single-break approach is best thought of as simply providing results for the first and second halves of the sample. The second method uses AR models estimated over rolling samples. The rolling regression estimated at date t is estimated by weighted least squares using two-sided exponential weighting, where the observation at date s received a weight of $\delta^{|t-s|}$, where we used a value of $\delta = 0.97$.[5] Both the split-sample and rolling AR models use four lags and are estimated using the detrended GDP growth rates from (1).

Table 3 shows the sum of the coefficients and the one-step-ahead forecast standard error for the split-sample AR models. The sum measures the persistence of an innovation to GDP, and by this measure GDP innovations have become substantially more persistent for Canada, France, and the UK. Persistence has increased slightly for the US and Italy, while it has declined for Germany and Japan. For all countries except Japan, the magnitude of the GDP innovations, as measured by the standard error of the regression, has decreased substantially: one-quarter-ahead forecasts based on univariate autoregressions have become more accurate for the G7 countries.

Figure 5 summarizes the results for the rolling AR models. Panel (a) presents estimated time paths for the sum of coefficients and panel (b) plots the estimated innovation standard deviation. These plots are consistent with the two-sample evidence in Table 3. In all countries, the innovation variance fell substantially,

5. Similar results are obtained using the non-Gaussian smoother estimates based on (2), and for values of δ ranging from 0.95 to 0.98. The two-sided exponential weighting scheme is used here for comparability with the two-sided VAR estimates reported in Sections 4 and 6.

TABLE 3. Autoregressive parameters for GDP growth rates: Sums of AR coefficients and standard error of the regression.

$$\Delta y_t = \alpha(L)\Delta y_{t-1} + \varepsilon_t$$

	Sum of AR coefficients ($\hat{\alpha}(1)$)		Standard error of the regression ($\hat{\sigma}_\varepsilon$)	
	1960–1983	1984–2002	1960–1983	1984–2002
Canada	0.00	0.56	3.82	2.27
France	−0.36	0.43	2.95	1.79
Germany	0.04	−0.18	5.42	3.39
Italy	0.02	0.13	4.03	2.16
JP	0.38	0.09	4.08	3.79
UK	0.03	0.65	4.81	1.84
US	0.30	0.47	3.98	1.96

Notes: These results are based on AR(4) models (excluding a constant) estimated using the detrended growth rates described in Section 2.

although it increased again during the 1990s in Japan. In Canada, France, and the UK, persistence has increased substantially, while persistence has been roughly constant for the US. The timing of these changes differs across countries, a result consistent with the different patterns of declining variances in Figure 4.

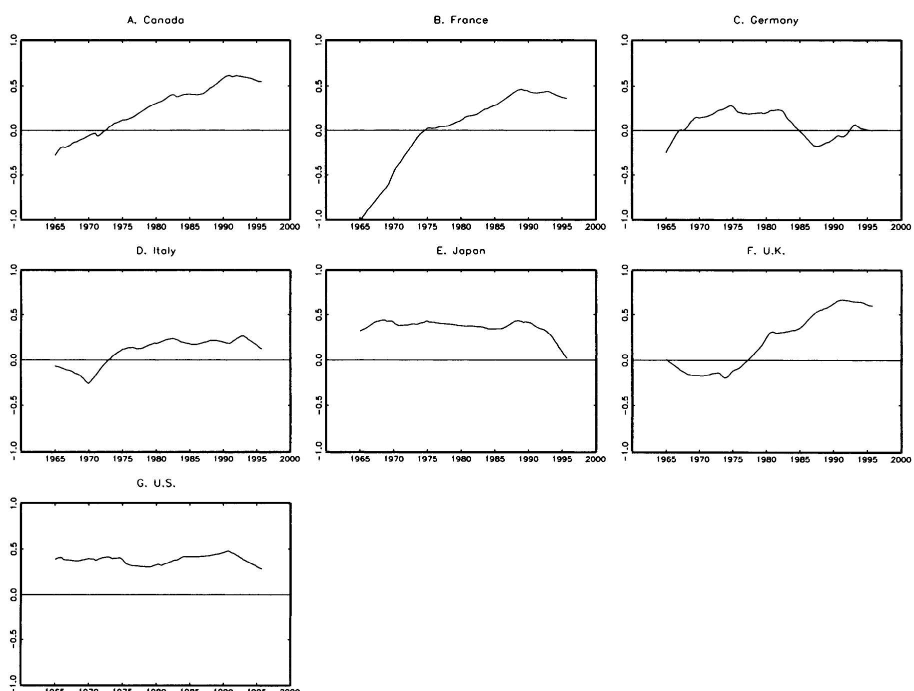

FIGURE 5a. Rolling autoregressions: Sum of AR coefficients ($\hat{\alpha}(1)$).

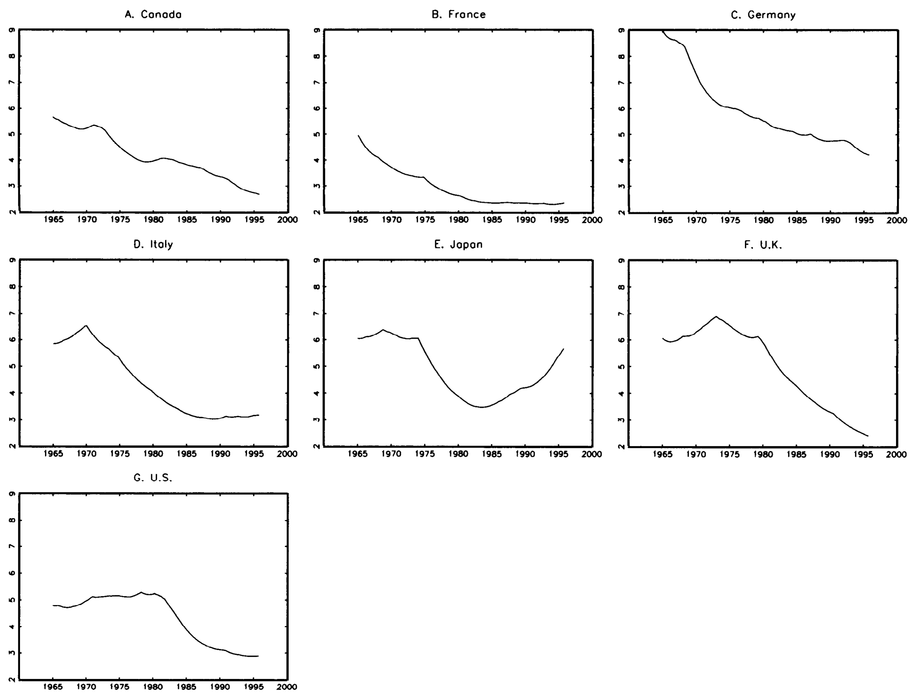

FIGURE 5b. Rolling autoregressions: Innovation standard error ($\hat{\sigma}_\varepsilon$).

4. Changes in Synchronization

This section reports various measures of time-varying international comovements of GDP. To facilitate comparisons with the analysis of Sections 5 and 6 using the FSVAR, these measures are estimated using a reduced form seven-country VAR. The section begins by describing the reduced form VAR, then turns to the measures of time-varying correlations.

4.1. Reduced Form VAR

A conventional VAR(p) with all seven countries would have $7p$ coefficients in each equation, where p is the number of lags. With the short quarterly data set at hand, this many coefficients would induce considerable sampling uncertainty even with small values of p. One solution to this dimensionality problem would be to consider VARs specified in terms of subsets of countries, but doing so would limit the international spillovers and common shocks that can be studied in a single model. Another solution is to specify a model for all seven countries but to impose additional restrictions on the VAR coefficients, as is done in many papers in this literature, for example Helg et al. (1995). We take this latter route and consider two such sets of restrictions.

For the main results, the restriction we use is for lagged foreign GDP growth to enter with a different number of lags than domestic GDP growth. Specifically, let Y_t be the vector of detrended quarterly GDP growth rates. The reduced form VAR is

$$Y_t = A(L)Y_{t-1} + v_t, \text{ where } Ev_t v_t' = \Sigma \tag{3}$$

where the diagonal elements of the matrix lag polynomial $A(L)$ have degree p_1 and the off-diagonal elements have degree p_2. Denote the resulting (restricted) VAR by VAR(p_1, p_2). The AIC and BIC, computed for the 1960–1983 and 1984–2002 subsamples, point to a VAR(4,1) specification. Because $p_1 \neq p_2$ the VAR was estimated using the method of seemingly unrelated regressions.

The second restriction we considered further restricts the coefficients on the lags of foreign GDP to be proportional to their trade shares, an approach taken by Elliott and Fatás (1997) and Norrbin and Schlagenhauf (1996). Results from that VAR are reported as part of the sensitivity analysis in Section 6.

The second moments of interest in this paper can all be computed directly from estimates of the VAR parameters in (3). The spectral density matrix of quarterly growth Y_t is $S_{YY}(\omega) = C(e^{i\omega})\Sigma C(e^{-i\omega})'/2\pi$, where $C(L) = [\mathrm{I} - A(L)L]^{-1}$. The implied spectral density matrix of four- quarter GDP growth is $|1+e^{i\omega}+e^{2i\omega}+e^{3i\omega}|^2 S_{YY}(\omega) = \{s_{ij}^{(4)}(\omega)\}$, so that $s_{ij}^{4}(\omega)$ is the cross-spectrum (spectrum when $i = j$) between four-quarter GDP growth in country i and country j at frequency ω. The implied spectral density matrix of the BP- filtered level of the logarithm of GDP is $|b(e^{i\omega})/(1 - e^{i\omega})|^2 S_{YY}(\omega)$, where b is the idealized BP filter so that $|b(e^{i\omega})|^2 = 1$ for $\omega_0 \leq \omega \leq \omega_1$, where the frequencies ω_0 and ω_1 respectively correspond to periodicities of 32 and 6 quarters, and $|b(e^{i\omega})|^2 = 0$ otherwise. Thus, for example, the contemporaneous correlation $\rho_{ij}^{(4)}$ between four-quarter growth rates in countries i and j is

$$\rho_{ij}^{(4)} = \frac{\int_{-\pi}^{\pi} s_{ij}^{(4)}(\omega)d\omega}{\left(\int_{-\pi}^{\pi} s_{ii}^{(4)}(\omega)d\omega\right)^{1/2} \left(\int_{-\pi}^{\pi} s_{jj}^{(4)}(\omega)d\omega\right)^{1/2}}. \tag{4}$$

As in Section 3, time variation in the VAR is captured by estimating the VAR parameters over the 1960–1983 and 1984–2002 subsamples and by rolling estimates of the VAR parameters. The rolling VAR parameters were estimated by weighted least squares using the two-sided exponential weighting scheme described in Section 3 for the rolling ARs.

Tests for instability in the parameters of the reduced-form VAR(4,1) are summarized in Table 4. Each cell in the table presents the p-value for the test of the hypothesis that the values of the parameters indicated in the column heading for the equation of that row are the same during 1960–1983 as they are during 1984–2002. The p-values are computed two ways: first, treating the 1984 break date as fixed (determined exogenously), and second (in brackets) treating the 1984

TABLE 4. Tests for a break in the reduced-form VAR parameters in 1984.

	All coefficients	Own lags	Other lags	Variance
Canada	0.00	0.00	0.35	0.00
	[0.05]	[0.03]	[0.97]	[0.00]
France	0.00	0.00	0.39	0.01
	[0.05]	[0.02]	[0.98]	[0.13]
Germany	0.14	0.71	0.05	0.00
	[0.75]	[1.00]	[0.41]	[0.05]
Italy	0.17	0.06	0.22	0.00
	[0.80]	[0.46]	[0.88]	[0.03]
Japan	0.62	0.46	0.55	0.60
	[1.00]	[0.99]	[1.00]	[1.00]
UK	0.00	0.00	0.22	0.00
	[0.02]	[0.04]	[0.88]	[0.00]
US	0.54	0.65	0.42	0.00
	[1.00]	[1.00]	[0.99]	[0.00]

Notes: The main entries are p-values for split-sample Chow–Wald tests of the hypothesis that the indicated set of VAR(4,1) parameters are the same in the 1960–1983 period as they are in 1984–2002 period, where the p-values were computed treating the break date as known a priori. For example, the cell, "Canada/Own lags" tests the hypothesis that the four coefficients on the lags of Canadian GDP growth in the equation for Canadian GDP growth are the same in the two periods. The degrees of freedom of the tests, by column, are 10, 4, 6, and 1. The entries in brackets are "sup-Wald" p-values computed using the conservative assumption that the 1984 break date was selected to maximize the break F-statistic in that particular cell, with 15% trimming at both ends of the full sample.

break date as having been chosen to maximize the value of the test statistic in that particular cell. To the extent that the break date was selected by examining the data, the first set of p-values overstate the statistical evidence of parameter instability, but because there is a single break date, not one selected to maximize any individual cell entry, the p-values in brackets are conservative and understate the evidence of parameter instability. In fact, we chose the 1984 break date based on the large body of evidence for the US so, for countries other than the US, the fixed-date p-values arguably are a better approximation than the conservative p-values in brackets. In any event, qualitatively similar conclusions are reached using both sets of p-values. There is evidence of changes in the VAR variances for Canada, Germany, Italy, the UK, and the US (the results for France depends on which p-value is used). There is also evidence of coefficient instability in the equations for Canada, France, and the UK. The hypothesis that all the VAR parameters are the same in the two subsamples is rejected at the 1% significance level using both the fixed break date and conservative critical values.

The changes in the VAR coefficients are difficult to interpret directly, so instead we turn to the implications of these changes for international output growth correlations.

4.2. International Synchronization

Table 5 presents various measures of international output comovements. Panels A and B tabulate the correlation of four-quarter GDP growth rates across countries, first using the raw data then based on the estimated model. The average absolute difference between the correlations in panel (a) and their counterpart in panel (b) is 0.04 in the first subsample and 0.10 in the second subsample, indicating that the reduced form VAR(4,1) captures most of the business cycle comovements of these series; the biggest exception is that the VAR(4,1) estimated correlation considerably exceeds the sample correlation between the US and French four-quarter GDP growth in the second period. Panel (c) of Table 5 presents the correlations among BP filtered GDP estimated using the reduced form VAR(4,1); the entries in panel (c) and correlations estimated directly from estimated BP- filtered data (not reported) differ by an absolute average of 0.08 in both the first and second periods.

TABLE 5. Correlations of GDP growth across countries.

(a) Four-quarter growth rates, simple correlation coefficients							
	Canada	France	Germany	Italy	Japan	UK	US
				1960–1983			
Canada	1.00						
France	0.31	1.00					
Germany	0.50	0.56	1.00				
Italy	0.30	0.59	0.35	1.00			
Japan	0.20	0.40	0.46	0.28	1.00		
UK	0.26	0.54	0.53	0.13	0.48	1.00	
US	0.77	0.39	0.52	0.21	0.32	0.46	1.00
				1984–2002			
Canada	1.00						
France	0.33	1.00					
Germany	0.12	0.59	1.00				
Italy	0.38	0.77	0.59	1.00			
Japan	−0.05	0.28	0.38	0.34	1.00		
UK	0.72	0.33	0.11	0.47	0.09	1.00	
US	0.80	0.26	0.22	0.29	0.02	0.58	1.00
Difference, 1984–2002 vs. 1960–1983 (std. error in parentheses)							
Canada							
France	0.02						
	(0.17)						
Germany	−0.37	0.03					
	(0.24)	(0.20)					
Italy	0.08	0.18	0.25				
	(0.13)	(0.15)	(0.24)				
Japan	−0.25	−0.12	−0.08	0.07			
	(0.21)	(0.24)	(0.20)	(0.23)			
UK	0.46	−0.21	−0.42	0.34	−0.39		
	(0.18)	(0.14)	(0.20)	(0.14)	(0.22)		
US	0.03	−0.13	−0.30	0.08	−0.30	0.11	
	(0.08)	(0.19)	(0.23)	(0.16)	(0.23)	(0.19)	

TABLE 5. CONTINUED

(b) Four-quarter growth rates, implied by reduced form VAR(4,1)							
	Canada	France	Germany	Italy	Japan	UK	US
				1960–1983			
Canada	1.00						
France	0.31	1.00					
Germany	0.57	0.56	1.00				
Italy	0.35	0.52	0.33	1.00			
Japan	0.33	0.29	0.39	0.22	1.00		
UK	0.31	0.52	0.50	0.12	0.44	1.00	
US	0.72	0.38	0.53	0.20	0.33	0.42	1.00
				1984–2002			
Canada	1.00						
France	0.56	1.00					
Germany	0.09	0.54	1.00				
Italy	0.45	0.79	0.49	1.00			
Japan	−0.02	0.15	0.33	0.13	1.00		
UK	0.70	0.58	0.18	0.56	0.03	1.00	
US	0.81	0.64	0.24	0.42	0.04	0.68	1.00
(c) BP-filtered GDP, implied by reduced form VAR(4,1)							
	Canada	France	Germany	Italy	Japan	UK	US
				1960–1983			
Canada	1.00						
France	0.36	1.00					
Germany	0.62	0.60	1.00				
Italy	0.39	0.56	0.37	1.00			
Japan	0.35	0.31	0.41	0.24	1.00		
UK	0.34	0.56	0.55	0.15	0.47	1.00	
US	0.76	0.44	0.59	0.23	0.34	0.46	1.00
				1984–2002			
Canada	1.00						
France	0.50	1.00					
Germany	0.08	0.59	1.00				
Italy	0.42	0.79	0.53	1.00			
Japan	−0.03	0.20	0.42	0.18	1.00		
UK	0.68	0.51	0.15	0.54	0.04	1.00	
US	0.79	0.59	0.23	0.37	0.07	0.64	1.00

Notes: These results are based on the detrended growth rates described in Section 2. Panel (a) shows the simple correlation coefficients estimated from four-quarter averages of the quarterly growth rates. The final block in panel (a) reports the difference in the correlations between the two subsamples and the standard error of that difference computed using the Newey–West estimator with a lag length of 6. Panels (b) and (c) are based on parameters from the VAR (4,1) model estimated over the two subsamples. Panel (b) shows the implied values of the correlations for the four-quarter growth rates from the VAR. Panel (c) shows the implied values of the correlation for the ideal (infinite order) 6–32-quarter band-pass filter.

Rolling correlations between own-country BP-filtered GDP and US and German BP- filtered GDP, based on the reduced-form VAR(4,1), are plotted in Figure 6; like the other plots of rolling estimates, the plotted date corresponds to

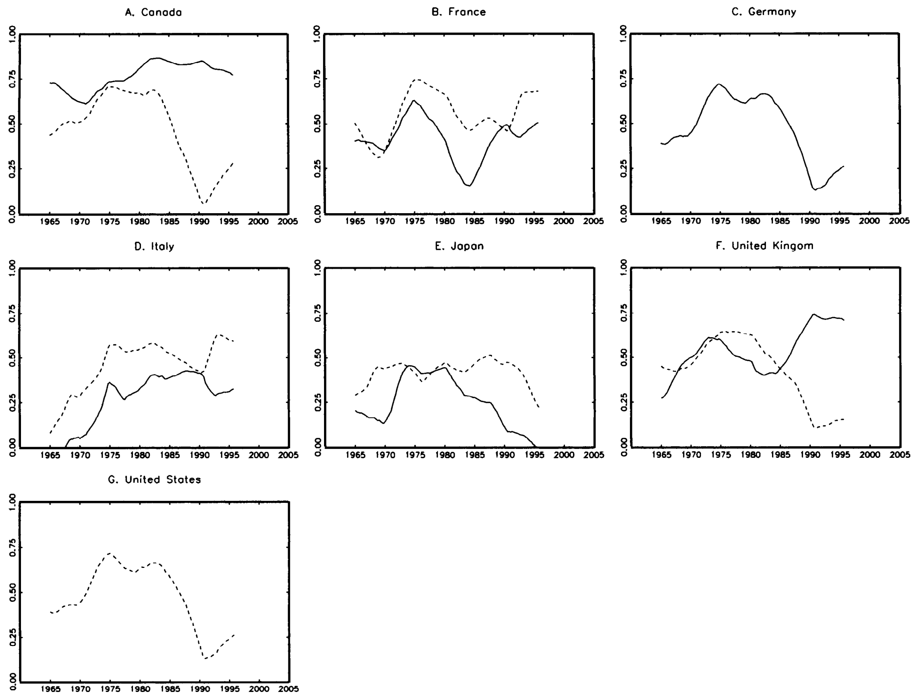

FIGURE 6. Band-pass GDP growth: Rolling correlation with US (solid line) and Germany (dashed line).

the center of the rolling window, the date with the greatest weight in the rolling exponential weighting scheme.

Three aspects of Table 5 and Figure 6 bear emphasis. First, as emphasized by Doyle and Faust (2002, 2005); Heathcote and Perri (2004); and Kose, Prasad, and Terrones (2003), there is no overall tendency towards closer international synchronization over this period: depending on the correlation measure used, the average cross-country correlation either is unchanged between the two subsamples or drops slightly. The final section of panel (a) reports the changes in the raw correlations over the two subsamples, along with their heteroskedasticity- and autocorrelation-consistent (HAC) standard errors (computed treating the 1984 break date as fixed). The average change across subsamples in these raw correlations is -0.05 (HAC standard error $= 0.10$).

Second, despite a lack of an overall increase in correlations, there appears to have been a shift in the pattern of comovements among the G7 economies. As Doyle and Faust (2005) emphasize, changes in correlations for individual pairs of countries are imprecisely measured, as can be seen by the large standard errors in the final part of panel (a). Nevertheless, there is evidence of the emergence of Euro-zone and English-speaking regional groups. Based on the correlations in

Table 5(a), during the first subsample the average correlation within the two groups was 0.50 (continental Europe) and 0.50 (English-speaking), and the average cross-group correlation was 0.38. In the second period the average correlations within the two groups rise to 0.65 (continental Europe) and 0.70 (English-speaking), while the average cross-group correlation drops to 0.28. Thus the average within-group correlation rose by 0.18 and the average cross-group correlation fell by 0.10. This contrast—the difference between the average change of the within-group correlations and the average change of the cross-group correlations—is 0.28 (HAC standard error = 0.10) and is statistically significant at the 1% level. A major source of this change is the decline in the correlation between UK GDP growth and that of France and Germany, and an increase in its correlation with the North American economies. The emergence of the two regional groups, English-speaking and Euro-zone, also is evident in Figure 6 through the increasing French–German and Italian–German correlations and the increasing correlation between the UK and the US (and their decreasing correlations with Germany).

Third, the synchronization of Japanese cycles with the rest of the G7 has been low throughout this 40-year period and recently decreased further. The average correlation between four-quarter GDP growth in Japan and that in the remaining G7 countries fell from 0.36 during 1960–1983 to 0.18 during 1984–2002. Although this decline of 0.18 is estimated imprecisely (HAC standard error = 0.17), it is large in economic terms and is consistent with the rolling correlations in Figure 6 and with other VAR-based evidence presented below that fluctuations in the Japanese economy became detached from those in the other G7 economies during the 1990s.

5. The Factor-Structural VAR model

There are several frameworks available for developing a time series model with enough structure to permit answering the questions of interest here, such as the fraction of a country's cyclical variance that is due to international shocks and how that has changed over time. Before discussing the specific framework used in this paper, a factor structural VAR, it is useful to discuss the competing modeling options and to assess their strengths and weaknesses.

The basic issue to be resolved is the best way to identify a world (or G7) shock. One approach is simply to define a world shock to be the innovation in a univariate time series model of world (or G7) GDP growth. While this approach has the advantage of being easy to implement, because US output receives great weight in G7 GDP it confounds world shocks with US shocks and idiosyncratic shocks to other large economies. Suppose there were in fact no common shocks and no trade; this identification scheme would nevertheless attribute a large fraction of US fluctuations to a common shock as an arithmetic implication of its construction.

A second approach is to use a parametric dynamic factor model in which the number of shocks exceeds the number of series, and the comovements across series at all leads and lags are attributed to the common shock. This results in an unobserved components model that can be estimated using Kalman filtering and related methods. This approach has been widely used in the international fluctuations literature; recent contributions include Kose, Otrok, and Whiteman (2003); Carvalho and Harvey (2002); Monfort et al. (2002); Kose, Prasad, and Terrones (2003); Luginbuhl and Koopman (2003); and Justiano (2004). This framework has several advantages. In the hypothetical case of no economic spillovers and no common shocks, there would be no comovements and the common shock would correctly be identified as having zero variance. This framework also captures the differences in dynamic responses of different economies to a world shock. On the other hand, because all cross-dynamics are attributed to the world shock, this approach is not well suited to identifying the separate effects of a common world shock and spillovers arising through trade: if there were in fact no world shocks but idiosyncratic shocks were transmitted through trade, the parametric dynamic factor model would incorrectly estimate a nonzero world shock.[6]

A third approach is to use nonparametric methods to estimate a dynamic factor model. If a large number of series have a dynamic factor structure, then the common component or the common dynamic factor can be estimated using principal components (Stock and Watson 2002b) or dynamic principal components (Forni et al. 2000). This strategy is used by Helg et al. (1995) to extract European industry and country shocks as principal components of reduced-form VAR errors, and by Helbling and Bayoumi (2003) to estimate the importance of common factors in G7 fluctuations. Prasad and Lumsdaine (2003) also adopt this strategy, using a weighting scheme rather than principal components to extract the innovation in a single common G7 factor. In principle the principal components/nonparametric approach has the advantages of the second approach without the disadvantage of assuming that all comovements stem from the common disturbance rather than through trade spillovers; in practice, however, if this approach is implemented using only G7 data then individual countries are necessarily heavily weighted leading to the same problems as the first approach, in particular finding a common factor even if there is none.

A fourth approach, the one used here, is to adopt a VAR framework for the lagged effects but to identify world shocks as those that affect all countries within the same period. Thus country-specific shocks can lead to spillovers, but those spillovers are assumed to happen with at least a one-quarter lag. This results in

6. Monfort et al. (2002) partially address this drawback by considering, as an alternative to their main analysis, a specification with regional shocks that interact dynamically and thus allow cross-region spillovers. Going further down this route and fully relaxing the lag dynamics would lead to the FSVAR model discussed later.

an overidentified factor structural VAR, in which the shocks are identified by imposing a factor structure on the reduced-form errors. Examples of papers using this approach (in a regional or international context) include Altonji and Ham (1990); Norrbin and Schlagenhauf (1996); and Clark and Shin (2000).

By defining international shocks to be the common components of the innovations in the seven-country VAR, the FSVAR identification scheme has several desirable features. In a world in which all shocks are country-specific and international transmission takes at least one quarter, no common shocks would be identified and this scheme would correctly conclude that there are no international shocks. This would be true even if lagged trade effects produce dynamic international comovements. Moreover, the lagged spillover effects of a country-specific shock would be correctly captured by the VAR dynamics. For example, monetary policy shocks are often modeled as having real effects after no shorter a lag than one quarter (e.g., Christiano, Eichenbaum, and Evans 1999), so under this standard identification assumption a surprise monetary contraction in the US that subsequently affects Canadian economic activity would be identified correctly by the FSVAR as an country-specific shock followed by a spillover, not as a common shock. The definition of what constitutes a common shock, however, does depend on the frequency of the data. For example, a financial crisis that starts in one country but spills over into other G7 financial markets within days would be identified in our quarterly FSVAR as a global shock (if it had real effects). Also, an international shock that affects one country first and the others after only a lag of a quarter or more would be misclassified by the FSVAR as an idiosyncratic shock, transmitted via spillovers.

We consider the FSVAR model consisting of the VAR model (3) in which the errors have the factor structure

$$\begin{gathered} v_t = \Gamma f_t + \xi_t, \text{ where} \\ E(f_t f_t') = \operatorname{diag}(\sigma_{f_1}, \ldots, \sigma_{f_k}) \text{ and } E(\xi_t \xi_t') = \operatorname{diag}(\sigma_{\xi_1}, \ldots, \sigma_{\xi_1}), \end{gathered} \tag{5}$$

where f_t are the common international factors, Γ is the $7 \times k$ matrix of factor loadings, and ξ_t are the country-specific, or idiosyncratic, shocks. In (5), the common international shocks (the common factors) are identified as those shocks that affect output in multiple countries contemporaneously. We estimate the FSVAR using Gaussian maximum likelihood.

The FSVAR specification (5) is overidentified, so empirical evidence can be brought to bear on the number of factors k. Likelihood ratio tests of the overidentifying restrictions are summarized in Table 6. In both subsamples and in the pooled full sample, the hypothesis of $k = 1$ is rejected against the unrestricted alternative (that is, against Σ_v having full rank) at the 1% significance level, but the null hypothesis of $k = 2$ is not rejected at the 10% significance level. These

TABLE 6. Tests of k-factor FSVAR vs. unrestricted VAR.

		1960–2002		1960–1983		1964–2002	
Number of Factors (k)	d.f.	LR Statistic	p-value	LR Statistic	p-value	LR Statistic	p-value
1	14	47.32	0.00	33.36	0.00	39.29	0.00
2	8	12.78	0.12	13.05	0.11	12.68	0.12
3	3	2.25	0.52	2.69	0.45	1.59	0.66

Notes: Entries are the likelihood ratio test statistic and its p-value testing the null hypothesis that the VAR(4,1) error covariance matrix has a k-factor structure, against the unrestricted alternative. The degrees of freedom of the test are given in the second column. These results are based on the detrended growth rates described in Section 2.

results suggest that $k = 2$ is appropriate, so we adopt a specification with two common international shocks.

6. Empirical Results

This section presents empirical results based on the two-factor FSVAR, including an analysis of the sensitivity of the results to some modeling decisions.

6.1. Changing Importance of Common and Country-Specific Shocks

The factor structure permits a decomposition of the h-step ahead forecast error for GDP growth in a given country into three sources: unforeseen common shocks, unforeseen domestic shocks, and spillover effects of unforeseen domestic shocks to other G7 countries. Because the country shocks and the common shocks are all uncorrelated by assumption, this decomposition in turn permits a threefold decomposition of the variances of the h-step-ahead forecast error and other filtered versions of GDP.

Table 7 summarizes these variance decompositions for GDP growth and for BP-filtered GDP. At the one-quarter horizon, international spillovers account for none of the GDP growth forecast error variance: this is the assumption used to identify the international shock. At longer horizons, spillovers typically account for between 5% and 15% of the variance of GDP growth, depending on the country and the subsample. Most of the variance of GDP growth is attributed to the common and idiosyncratic domestic shocks, but their relative importance varies considerably across countries. In the first period, the effects of international shocks at the four-quarter horizon are estimated to be the greatest for Canada, France, and Germany, and the least for Italy and Japan. In the second period, almost all the forecast error variance in Japan is attributed to domestic shocks, a result consistent with the declining correlation between GDP in Japan and in other countries in the second period reported in Section 4.

The relative importance of international sources of fluctuations, either common shocks or spillovers, can be measured as one minus the share of the forecast

TABLE 7. Variance decompositions based on the two-factor FSVAR: Common shocks, spillovers, and own-country shocks.

		1960–1983				1984–2002			
		Forecast error standard deviation	Fraction of forecast error variance due to:			Forecast error standard deviation	Fraction of forecast error variance due to:		
Country	Horizon		Int'l shocks	Spillovers	Own shock		Int'l shocks	Spillovers	Own shock
(a) GDP Growth									
Canada	1	3.37	0.36	0.00	0.64	2.04	0.97	0.00	0.03
	2	2.70	0.45	0.09	0.46	1.77	0.92	0.05	0.03
	4	2.01	0.50	0.16	0.34	1.71	0.89	0.09	0.02
	8	1.43	0.52	0.17	0.31	1.59	0.83	0.15	0.02
France	1	2.66	0.97	0.00	0.03	1.62	0.96	0.00	0.04
	2	1.90	0.87	0.11	0.02	1.23	0.93	0.04	0.03
	4	1.29	0.82	0.16	0.02	1.11	0.91	0.06	0.03
	8	0.87	0.81	0.18	0.01	1.06	0.88	0.10	0.02
Germany	1	4.81	0.24	0.00	0.76	3.12	0.26	0.00	0.74
	2	3.35	0.33	0.10	0.57	2.02	0.31	0.06	0.63
	4	2.32	0.38	0.15	0.47	1.26	0.34	0.07	0.59
	8	1.73	0.41	0.16	0.43	0.90	0.39	0.08	0.53
Italy	1	3.86	0.10	0.00	0.90	1.96	0.33	0.00	0.67
	2	3.11	0.10	0.02	0.88	1.40	0.41	0.05	0.54
	4	2.42	0.12	0.04	0.84	1.09	0.45	0.08	0.47
	8	1.59	0.15	0.06	0.80	0.88	0.51	0.12	0.37
Japan	1	3.96	0.17	0.00	0.83	3.62	0.01	0.00	0.99
	2	3.01	0.19	0.02	0.79	2.52	0.00	0.03	0.97
	4	2.49	0.20	0.02	0.78	1.84	0.00	0.03	0.97
	8	1.98	0.20	0.03	0.77	1.37	0.01	0.04	0.95
UK	1	4.66	0.24	0.00	0.76	1.69	0.03	0.00	0.97
	2	3.22	0.23	0.03	0.74	1.56	0.10	0.00	0.90
	4	2.35	0.24	0.03	0.72	1.31	0.20	0.02	0.78
	8	1.71	0.25	0.04	0.71	1.22	0.29	0.03	0.68
US	1	3.95	0.27	0.00	0.73	1.74	0.22	0.00	0.78
	2	3.23	0.31	0.01	0.68	1.41	0.29	0.05	0.66
	4	2.55	0.33	0.02	0.65	1.29	0.38	0.12	0.50
	8	1.84	0.35	0.02	0.63	1.20	0.47	0.17	0.36
(b) BP-filtered GDP									
Canada		1.19	0.50	0.20	0.30	1.16	0.80	0.18	0.02
France		0.74	0.77	0.21	0.01	0.76	0.85	0.14	0.02
Germany		1.34	0.41	0.19	0.41	0.72	0.39	0.10	0.51
Italy		1.46	0.14	0.06	0.81	0.67	0.49	0.13	0.38
Japan		1.49	0.20	0.03	0.77	1.07	0.02	0.05	0.93
UK		1.33	0.25	0.05	0.71	0.86	0.33	0.05	0.62
US		1.51	0.34	0.03	0.63	0.87	0.44	0.20	0.37

Notes: This table shows the standard deviation and three-way decomposition of variance of filtered versions of GPD. Panel (a) shows results for FSVAR forecast errors at the one-, two-, four-, and eight-quarter horizon. Panel (b) shows results for the ideal (infinite order) 6–32-quarter band-pass filtered values of GDP. The standard deviations in panel (a) are in percentage points at an annual rate ($(400/h)$ times the forecast error, where h is the forecast horizon), and the standard deviations in panel (b) are in percentage points. These results are based on the FSVAR model estimated using the detrended growth rates described in Section 2.

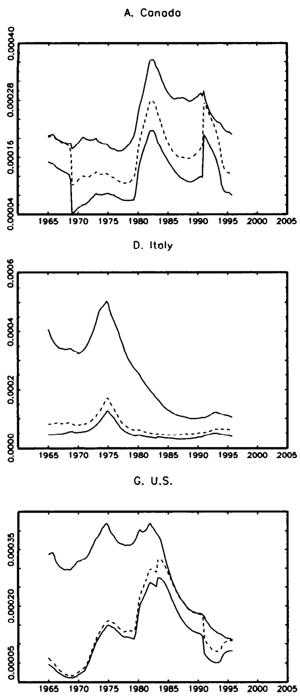

FIGURE 7. Time-varying variances of BP-filtered GDP growth due to: international shocks (lower); international shocks + spillovers (middle); and total (top). (Computed using rolling estimates of the two-factor FSVAR).

error variance attributed to domestic shocks; a small domestic share corresponds to a relatively larger role for international rather than domestic disturbances. Only Japan and Germany show a marked increase in the fraction of the variance attributed to domestic shocks, while Canada, Italy, and the US show a marked decrease. The variance decompositions for BP- filtered GDP yield similar conclusions to the variance decompositions of GDP growth at the four- and eight-quarter-ahead horizon.

Figure 7 presents time-varying estimates of the variance decomposition of BP-filtered GDP, based on rolling estimates of the two-factor FSVAR (as before, using exponential weighting). The units in Figure 7 are those of the variance; the lower line is the contribution to the variance of the international shocks, the middle line is the sum of the contributions of the international shocks and spillovers, and the top line is total variance, so the gap between the top and middle lines is the contribution to the variance of domestic shocks. For Germany, the UK, and the US, the recent decline in the overall volatility tracks a decline in the variance arising from international shocks. For Italy, the large historical decline in the variance is associated with a declining importance of domestic shocks. For Japan, international shocks have become unimportant, and domestic shocks explain nearly all of its volatility in the 1990s and are the source of its recent increase in volatility.

The correlations presented in Section 4 suggest the emergence of a Euro-zone cluster in the second period. This raises the question of whether one of the factors in the second period might be interpreted as a "Euro-zone only" factor. The hypothesis that one of the common factors loads only on France, Germany, and Italy provides three testable restrictions on the FSVAR. In the FSVAR estimated over 1960–1983, this restriction is rejected at the 5% significance level ($p = 0.02$), but when estimated over 1984– 2001, the restriction is not rejected at the 10% significance level ($p = 0.31$). Thus, the hypothesis that one of the two factors corresponds to a continental Europe factor can be rejected in the first period but not in the second, providing a precise interpretation of the apparent emergence of the Euro-zone cluster.

6.2. *Changes in Volatility: Impulse or Propagation?*

In principle, the contribution of international shocks to output volatility could decrease because the variance of the international shocks has decreased, because a shock of a fixed magnitude has less of an effect on the economy, or both. Said differently, the variance of GDP growth in a given country can change because the magnitude of the shocks impinging on that economy have changed or because the effects of those shocks have changed.

In this section, we decompose the change in the variance from the first subsample to the second into changes in the magnitudes of the shocks ("impulses") and changes in their effect on the economy ("propagation"). To make this precise, let V_p denote the variance of the 4-quarter-ahead forecast errors in a given country in period p, where $p = 1, 2$ corresponds to 1960–1983 and 1984–2002. The variance decomposition attributes a portion of V_p to each of the nine shocks in the model, so we can write, $V_p = V_{p,1} + \cdots + V_{p,9}$, where $V_{p,j}$ is the variance in period p attributed to the jth shock. Thus the change in the variance between the two periods is $V_2 - V_1 = (V_{2,1} - V_{1,1}) + \cdots + (V_{2,9} - V_{1,9})$. In identified structural VARs, the variance component $V_{p,j}$ always can be written as $a_{pj}\sigma^2_{pj}$, where a_{pj} is a term depending on the squared cumulative impulse response of GDP to shock j in period p and σ^2_{pj} is the variance of shock j in period p Thus the change in the contribution of the jth shock can be decomposed exactly as

$$V_{2j} - V_{1j} = \left(\frac{a_{1j} + a_{2j}}{2}\right)\left(\sigma^2_{2j} - \sigma^2_{1j}\right) + \left(\frac{\sigma^2_{1j} + \sigma^2_{2j}}{2}\right)(a_{2j} - a_{1j}). \quad (6)$$

That is, the change in the variance can be decomposed into the contribution from the change in the shock variance plus the contribution from the change in the impulse response. The decomposition (6) is additive so these contributions can be aggregated into variance changes arising from the common shocks, spillovers, and own shocks, with each type of shock in turn decomposed into changes in variances arising from changing shock variances and from changing impulse responses; this yields a six-way decomposition of the change in the variance of GDP forecast errors from the first period to the second.

This decomposition (and the counterfactual calculations in the next subsection) requires that the covariance matrix of the factors, Σ_{ff}, and the factor loadings, Γ, are separately identified. We identify the factors by assuming that they are uncorrelated (so that Σ_{ff} is diagonal) and that the second factor has no impact effect on the US (so that $\Gamma_{US,2} = 0$). These restrictions yield factors with a plausibly stable interpretation across the two subsamples. The dramatic changes in Europe suggest that other identifying assumptions, such as $\Gamma_{\text{France},2} = 0$, are unlikely to yield factors with the same interpretations across subsamples. The scale of the factors is identified by the restriction that each column of Γ has unit length, that is $\Gamma_i'\Gamma_i = 1$ for $i = 1, 2$. We investigate alternative assumptions in Section 6.5.

Table 8 presents this six-way decomposition of the change in variances of four quarter- ahead forecast errors in GDP. Standard errors, computed using parametric bootstrap simulations, are shown in parentheses. Evidently, the decline in the variance between the two periods is to a great extent attributed to a decline in the magnitudes of the shocks. For all countries except Japan changes in the variance of shocks led to a large and statistically significant decline in volatility. Indeed, for Canada, France, the UK, and the US, the decline in the shock variances more than accounts for the drop in the variance of GDP forecast errors, in the sense that changes in the propagation mechanism worked to increase rather than to decrease the total variance across these two periods (although this increase is statistically significant only for Canada). For Germany and Italy, the net contribution of changes in propagation is small, so that most of the variance reductions in Germany and Italy are attributed to changes in the magnitudes of the shocks. The exception here, as we have seen in other aspects of this analysis, is Japan, in which the decline in the variance is largely attributed to changes in the propagation mechanism, not to changes in the size of shocks. Among the different types of shocks, reductions in the size of country-specific shocks is important in all countries except France and Japan. A reduction in the size of international shocks played a substantial role in the volatility moderation in Canada, France, Germany, and the US. In addition, in all countries a small, typically statistically insignificant portion of the moderation is attributed to smaller foreign idiosyncratic shocks.

TABLE 8. Decomposition of changes in the variance of four-quarter-ahead FSVAR forecast errors into changing impulses and changing propagation.

	Variances			Contribution of change in shock variance				Contribution of change in impulse response function			
	1960–1983	1984–2002	Change	Int'l	Spillover	Own	Total	Int'l	Spillover	Own	Total
Canada	4.06	2.93	−1.13	−3.37	−0.71	−2.40	−6.48	3.94	0.31	1.09	5.35
	(0.88)	(0.67)	(1.09)	(1.32)	(0.38)	(0.99)	(1.72)	(1.62)	(0.49)	(0.57)	(1.94)
France	1.66	1.22	−0.44	−1.28	−0.32	−0.02	−1.63	1.04	0.12	0.03	1.19
	(0.35)	(0.28)	(0.45)	(0.50)	(0.15)	(0.30)	(0.62)	(0.65)	(0.19)	(0.14)	(0.75)
Germany	5.39	1.59	−3.80	−1.04	−0.43	−1.41	−2.88	−0.47	−0.27	−0.18	−0.92
	(1.16)	(0.36)	(1.21)	(0.49)	(0.21)	(0.52)	(0.71)	(0.65)	(0.30)	(0.38)	(0.92)
Italy	5.86	1.19	−4.67	−0.65	−0.32	−3.18	−4.15	0.45	0.20	−1.18	−0.52
	(1.31)	(0.27)	(1.34)	(0.43)	(0.19)	(0.81)	(0.89)	(0.67)	(0.28)	(0.61)	(1.00)
Japan	6.18	3.39	−2.79	−0.42	−0.24	−0.01	−0.67	−0.78	0.20	−1.54	−2.13
	(1.39)	(0.83)	(1.63)	(0.42)	(0.29)	(0.97)	(1.00)	(0.70)	(0.40)	(1.06)	(1.43)
UK	5.52	1.73	−3.80	−0.80	−0.23	−5.03	−6.07	−0.19	0.08	2.39	2.27
	(1.24)	(0.41)	(1.31)	(0.52)	(0.15)	(1.33)	(1.35)	(0.74)	(0.22)	(1.04)	(1.44)
US	6.51	1.66	−4.84	−1.40	−0.54	−3.26	−5.20	−0.13	0.62	−0.13	0.36
	(1.49)	(0.38)	(1.53)	(0.93)	(0.29)	(1.26)	(1.37)	(1.17)	(0.39)	(0.67)	(1.50)

Notes: The first three columns give the variance of BP-filtered GDP (in percentage points) in the first and second subsample, using the estimated FSVAR (identified as described in Section 6.2), and their difference. The remaining columns decompose this difference into changes in the impulse response functions and changes in the variances of the shocks themselves. The sum of the "international," "spillover," and "own" columns equals the "total" column, and the sum of the two "total" columns equals the "change" column. Estimated standard errors are shown in parentheses.

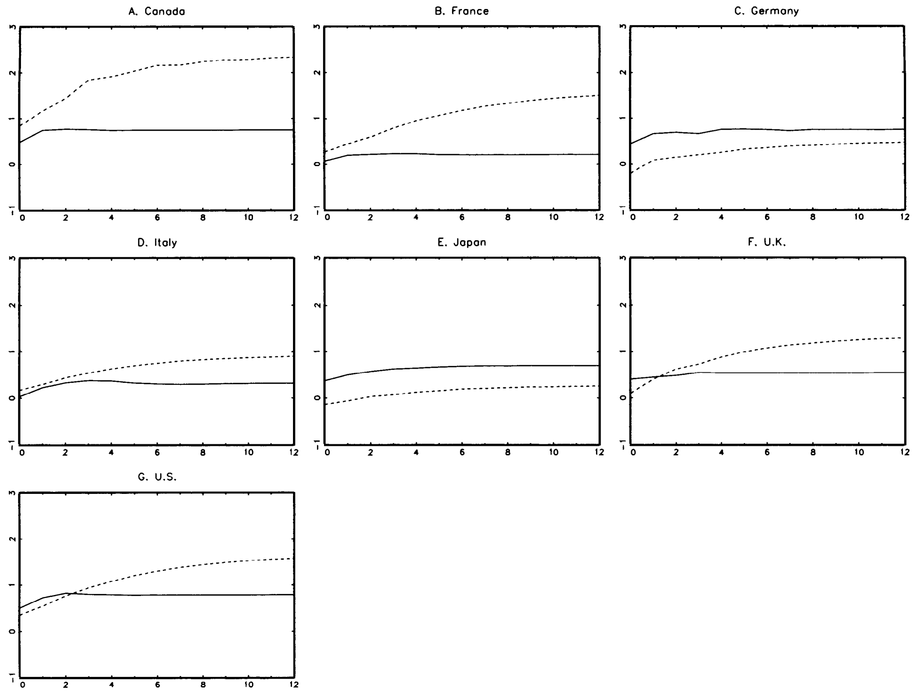

FIGURE 8a. Cumulative impulse response of country GDP growth with respect to the first common factor in 1960–1983 (solid line) and 1984–2002 (dashed line).

One lesson from Table 8 is that there have been important changes in the effect of an international shock of a fixed magnitude on some of these economies. This changing effect is examined further in Figure 8, which presents the impulse response functions for the different countries in the two subsamples with respect to the first common factor (Figure 8a) and the second common factor (Figure 8b). For the first factor there is a large estimated increase in the magnitude of the effect of the common shocks and in its persistence for Canada, France, Italy, the UK, and the US. The second factor has become more important for France, Germany and Italy and generally less important for the other countries. Again, Japan is different than the rest of the G7, with the estimated responses to both shocks being nonzero in the first period but nearly zero in the second.

6.3. Counterfactuals: Second-Period Propagation, First-Period Shocks

The foregoing analysis indicate that much of the moderation is attributable to declines in the variance of the common international shocks. This raises the counterfactual question: what would the volatility and cross-correlations have been in

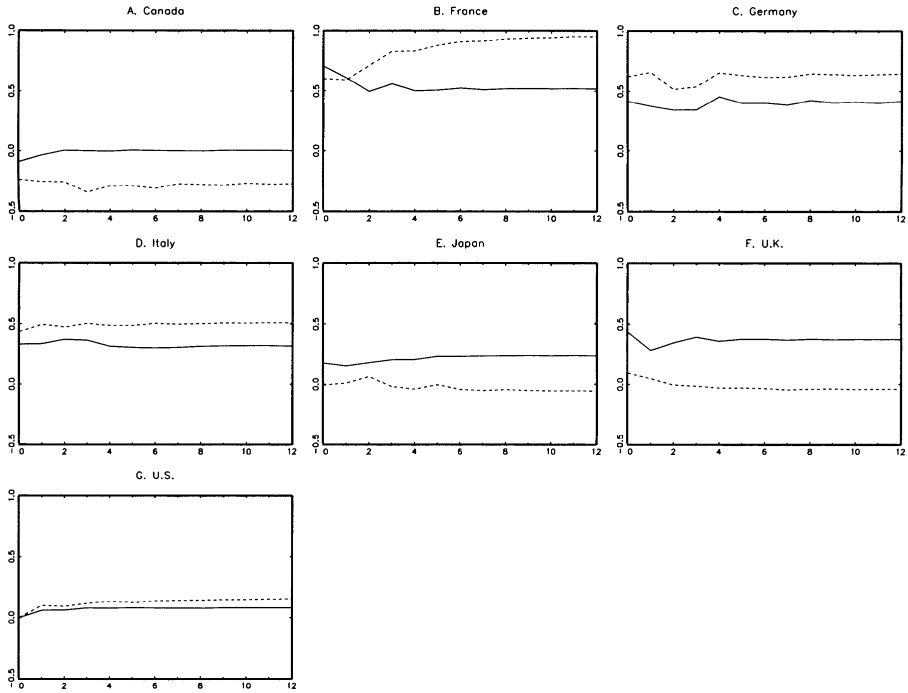

FIGURE 8b. Cumulative impulse response of country GDP growth with respect to the second common factor in 1960–1983 (solid line) and 1984–2002 (dashed line).

1984–2002, had the G7 economies been confronted with common international shocks as large as those experienced in 1960–1983?

This counterfactual question can be addressed by suitably combining the impulse responses from the second period FSVAR with the shock variances from the first-period FSVAR, then computing the implied moments. The resulting estimated variances are summarized in Table 9. Comparing the first line of each panel (the estimated standard deviations based on second-period impulse response functions and second-period shock variances) with the second line (in which the first-period variance of the common shocks is used) reveals that all countries, except again Japan, would have had considerably greater volatility over the past two decades had the world experienced the first-period shocks. For example, the standard deviation of four-quarter GDP growth in the US would have been approximately 2.2 percentage points, compared with the actual value of 1.6 percentage points; the standard deviation of French four-quarter GDP growth, which in reality was essentially constant over the two periods, would have increased from 1.4 to 2.2 percentage points had the second period experienced international shocks of the same magnitude as the first period.

TABLE 9. FSVAR-based counterfactual volatility measures during 1984–2002 using common and country shock variances from 1960–1983.

(a) Standard deviations of four-quarter GDP growth

Period for shock variances		Standard deviation of four-quarter GDP growth						
Common shocks	Country shocks	Canada	France	Germany	Italy	Japan	UK	US
84–02	84–02	2.06	1.43	1.34	1.21	1.87	1.59	1.60
60–83	84–02	3.35	2.24	1.68	1.64	1.91	2.06	2.23
60–83	60–83	4.26	2.64	2.17	2.50	2.11	3.61	3.34

(b) Standard deviations of BP-filtered GDP

Period for shock variances		Standard deviation of BP-filtered GDP						
Common shocks	Country shocks	Canada	France	Germany	Italy	Japan	UK	US
84–02	84–02	1.16	0.76	0.72	0.67	1.07	0.86	0.87
60–83	84–02	1.89	1.19	0.90	0.91	1.09	1.13	1.20
60–83	60–83	2.40	1.38	1.16	1.37	1.19	1.97	1.82

Notes: Entries in panel (a) are the standard deviations of four-quarter GDP growth (in percentage points at an annual rate) based on the estimated FSVAR impulse response functions (identified as described in Section 6.2) estimated using data from 1984–2002, using the shock variances estimated over the sample indicated in the first two columns. The first row is the model-based estimate of the actual standard deviation during 1984–2002; the remaining rows are counterfactuals. The entries in panel (b) are analogous to those in panel (a) but pertain to BP-filtered GDP (in percentage points).

The cross-country correlations implied by this counterfactual scenario are summarized in Table 10. Under the counterfactual scenario the correlations typically increase by 0.10 (Japan again is the exception). According to these estimates, had the common shocks in the second period been as large as they were during the first period, international business cycles would have been more highly synchronized than they actually were, and indeed would have been more highly synchronized than there were in the 1960–1983 period.[7]

6.4. *An Examination of the International Shocks*

Because moderation of the international shocks appears to be an important source of the moderation in G7 volatility, it is of interest to see if these international shocks can be linked to observable and interpretable time series.

This section examines several candidates for such observable shocks, taken from Stock and Watson (2002a). The first candidate is US monetary policy shocks;

7. The counterfactual exercises reported here assume that the VAR coefficients and idiosyncratic shock variances do not change when the factor variances change. In some models, such as the model of Heathcote and Perri (2004), these parameters may change, raising Lucas critique caveats concerning these counterfactual calculations.

TABLE 10. FSVAR-based counterfactual correlations between four-quarter growth rates during 1984–2002 using common shock variances from 1960–1983.

(a) FSVAR estimates of actual 1984–2002 correlations

	Canada	France	Germany	Italy	Japan	UK	US
Canada	1.00						
France	0.57	1.00					
Germany	0.10	0.55	1.00				
Italy	0.48	0.80	0.48	1.00			
Japan	0.01	0.16	0.27	0.15	1.00		
UK	0.70	0.58	0.19	0.56	0.05	1.00	
US	0.81	0.66	0.23	0.53	0.13	0.70	1.00

(b) FSVAR estimates of 1984–2002 correlations using common shock variances from 1960–1983

	Canada	France	Germany	Italy	Japan	UK	US
Canada	1.00						
France	0.63	1.00					
Germany	0.17	0.66	1.00				
Italy	0.57	0.88	0.63	1.00			
Japan	0.07	0.19	0.28	0.18	1.00		
UK	0.79	0.68	0.33	0.67	0.13	1.00	
US	0.87	0.76	0.35	0.67	0.18	0.82	1.00

Notes: Entries in panel (a) are the correlations among four-quarter GDP growth based on the FSVAR estimated using data from 1984–2002. Entries in panel (b) are based on the 1984–2002 FSVAR (identified as described in Section 6.3), except calculated using the common shock variances from the 1960–1983 FSVAR.

although these are domestic shocks, were they to affect other countries within the quarter that they occur, then they would be classified as common international shocks in the FSVAR identification scheme. Many methods have been proposed for identifying monetary policy shocks; here, we adopt Christiano, Eichenbaum, and Evans' (1997) identification method. The second candidate series is US productivity shocks, identified using Galí's (1999) method; we treat this as a proxy for world productivity shocks. The third set of shocks are innovations to commodity prices, measured here by an aggregate index of commodity prices, an index for food, an index of industrial materials, and an index of sensitive material prices, all for the US. The final set of shocks are oil prices, measured in three ways: the nominal growth rate in oil prices (in the US), and Hamilton's (1996) oil price series, which is the larger of zero and the percentage difference between the current price and the maximum price during the past four quarters. For details of construction of these series, see Stock and Watson (2002a).

Table 11 reports the largest squared canonical correlations between the factors and the leads and lags of the candidate observable shock series.[8] In the first period,

8. The largest canonical correlation is the correlation between a linear combination of the factors and a linear combination of the leads and lags of the observable shock series, where the linear combinations are chosen to maximize that (squared) correlation. This measure has the advantage of not requiring additional normalizations for identifying the two factors separately.

TABLE 11. Squared canonical correlations between international factors and various observable shocks.

	1960–2001	1960–1983	1984–2001
US money (CEE)	0.103	0.100	0.024
US productivity (Galí)	0.061	0.016	0.058
Commodity prices: all	0.046	0.069	0.056
Commodity prices: food	−0.004	0.001	−0.055
Industrial materials prices	0.089	0.107	0.124
Sensitive materials prices	0.107	0.128	0.081
Oil price (nominal)	−0.028	0.156	−0.034
Oil price (Hamilton)	0.037	0.154	0.025

Notes: Entries are the largest squared canonical correlation (adjusted for degrees of freedom) between the two factors from the FSVAR model and four leads and lags of the series listed in the first column. These series are described in the text.

the common international shocks are somewhat correlated with the US monetary policy shock and with the oil price measures, but not with the other shocks. Otherwise, however, the squared canonical correlations are nearly zero or are negative (possible because of the degrees of freedom adjustment), indicating that the common international shocks in the FSVAR are in these cases unrelated to these candidate observable shocks. Admittedly Table 11 represents a rather coarse attempt to identify the source of the international factors as several of the candidate shocks examined in Table 11 are US-centric, and an obvious next step is to examine alternative measures of global shocks.

6.5. Sensitivity Analysis

This section reports the results of two checks of the foregoing results to changes in the modeling assumptions or in the statistics reported.

Trade-Weighted VAR Lag Restrictions. As a check, we considered a further restriction of the VAR in which the coefficients on foreign GDP are proportional to trade shares. Elliott and Fatás (1996) used a similar restriction to identify shocks in a structural VAR, and Norrbin and Schlagenhauf (1996) used it (as we do here) to simplify the lag dynamics. Accordingly, the restricted reduced form VAR is

$$Y_t = b(L)Y_{t-1} + d(L)WY_{t-1} + v_t, \tag{7}$$

where (a) $Ev_t v_t' = \Sigma$, where $b(L)$ and $d(L)$ are diagonal lag polynomial matrices, and W is a fixed weighting matrix. The diagonal elements of W are zero and the (i, j) element is the share of gross trade (imports plus exports) of trading partner j in all of country i's trade with G7 countries.[9]

9. Bilateral import and export data are from the IMF's IFS database.

TABLE 12. Sensitivity check: Counterfactual standard deviation of four-quarter GDP growth based on trade-weighted FSVAR.

Period for shock variances		Standard deviation of four-quarter GDP growth						
Common shocks	Country shocks	Canada	France	Germany	Italy	Japan	UK	US
84–02	84–02	1.88	1.20	1.36	1.10	1.93	1.49	1.43
60–83	84–02	2.94	1.80	1.62	1.39	1.98	1.58	1.83
60–83	60–83	3.69	1.84	2.11	2.12	2.13	3.36	2.73

Note: Entries are computed in the same way as in panel (a) of Table 9, except they are based on the FSVAR (7) with trade-weight lag restrictions.

In the restricted reduced form VAR (7), the number of coefficients per equation equals the number of own lags (the degree of $b(L)$) plus the number of lags on trade-weighted foreign GDP (the degree of $d(L)$). AIC and BIC comparisons point to four own lags and one lag of trade-weighted foreign GDP growth. The FSVAR corresponding to (7) imposes the factor structure (5) on the reduced form errors in (7), and the model is estimated by Gaussian maximum likelihood.

As a gauge of the sensitivity of the results in the previous sections, we recomputed the counterfactual variances and correlations of Tables 9 and 10 for the trade-weighted FSVAR; the results are reported in Tables 12 and 13. Although

TABLE 13. Sensitivity check: Counterfactual correlations of four-quarter GDP growth based on trade-weighted FSVAR.

(a) Trade-weighted FSVAR estimates of actual 1984–2002 correlations

	Canada	France	Germany	Italy	Japan	UK	US
Canada	1.00						
France	0.22	1.00					
Germany	−0.09	0.52	1.00				
Italy	0.15	0.68	0.43	1.00			
Japan	0.20	0.11	0.10	0.08	1.00		
UK	0.15	0.35	0.20	0.24	0.07	1.00	
US	0.75	0.34	0.09	0.24	0.30	0.17	1.00

(b) Trade-weighted FSVAR estimates of 1984–2002 correlations using common shock variances from 1960–1983

	Canada	France	Germany	Italy	Japan	UK	US
Canada	1.00						
France	0.28	1.00					
Germany	−0.12	0.60	1.00				
Italy	0.22	0.80	0.55	1.00			
Japan	0.22	0.14	0.12	0.11	1.00		
UK	0.23	0.46	0.28	0.37	0.11	1.00	
US	0.84	0.43	0.08	0.35	0.32	0.28	1.00

Note: Entries are computed in the same way as in Table 10, except they are based on the FSVAR (7) with trade-weight lag restrictions.

the numerical values for the estimated changes in variances in Tables 9(a) and 12 differ, the qualitative conclusions are the same. In most countries, the variances of four-quarter GDP growth would have been considerably larger had second-period shocks been as large as first- period shocks. The main differences between the standard deviations in Tables 9(a) and 12 is the estimated increase for the UK, which is less using the trade-weighted FSVAR than the base case FSVAR in Table 9(a). The main differences between the implied correlations in Tables 10 and 13 is that the model-based estimates in Table 13(a) (estimates of actuals, not counterfactuals) in some cases differ considerably from the actual sample correlations in Table 5(a). The trade-weighted model especially fails to capture the correlations involving Canada. In this sense, the trade-weighted FSVAR does not fit the data as well as the FSVAR(4,1). Still, the main conclusion from Table 10—that international synchronization would have been substantially greater had the common shocks in the second period been as large as they were in the first—also obtains using the trade-weighted FSVAR.

Measuring Synchronization by Average Coherences. The analysis of international synchronization so far has relied on contemporaneous cross-correlations of four-quarter GDP growth and of BP-filtered GDP as the measures of comovements, but this can mask lagged associations. An alternative measure of comovements, which is invariant to these lagged effects, is the average coherence at business cycle frequencies. Specifically, let ω_0 and ω_1 be the lower and upper frequencies that define the business cycle portion of the spectrum, and let $s_{ij}^{(4)}(\omega)$ be the cross-spectrum between the four quarter growth rates in countries i and j as in (4). One measure of the average coherence between four- quarter growth rates in countries i and j at business cycle frequencies is,

$$R_{ij}^2(\omega_0, \omega_1) = \frac{\int_{\omega_0}^{\omega_1} \left\| s_{ij}^{(4)}(\omega) \right\|^2 d\omega}{\left(\int_{\omega_0}^{\omega_1} \left\| s_{ii}^{(4)}(\omega) \right\|^2 d\omega \right)^{1/2} \left(\int_{\omega_0}^{\omega_1} \left\| s_{jj}^{(4)}(\omega) \right\|^2 d\omega \right)^{1/2}}. \tag{8}$$

This measure reduces to the usual definition of the coherence when it is evaluated at a single frequency rather than over the range ω_0 to ω_1.

The square root of average coherence (8), $R_{ij}(\omega_0, \omega_1)$, was computed for the counterfactual correlations examined in Table 10, and the results are summarized in Table 14 (because the coherence has the interpretation of an R^2, using the square root of the average coherence makes this measure more directly comparable to the correlations of Table 10). Comparing panel (a) of Tables 10 and 14 shows that the coherences are higher than the correlations of four-quarter growth rates, which is not surprising because the coherences are not sensitive to phase shifts and also focus on business cycle frequencies, whereas the four-quarter growth rates contain some higher frequency noise. The qualitative conclusions from the

TABLE 14. Sensitivity check: Counterfactual coherences between four-quarter growth rates during 1984–2002 using common shock variances from 1960–1983.

(a) FSVAR estimated actuals for 1984–2002							
	Canada	France	Germany	Italy	Japan	UK	US
Canada	1.00						
France	0.73	1.00					
Germany	0.32	0.61	1.00				
Italy	0.64	0.88	0.58	1.00			
Japan	0.20	0.18	0.34	0.21	1.00		
UK	0.82	0.71	0.33	0.68	0.19	1.00	
US	0.88	0.77	0.36	0.67	0.21	0.82	1.00

(b) FSVAR estimated counterfactuals for 1984–2002 using common shock variances from 1960–1983							
	Canada	France	Germany	Italy	Japan	UK	US
Canada	1.00						
France	0.77	1.00					
Germany	0.43	0.73	1.00				
Italy	0.73	0.94	0.72	1.00			
Japan	0.26	0.22	0.33	0.25	1.00		
UK	0.89	0.77	0.44	0.76	0.25	1.00	
US	0.93	0.84	0.50	0.79	0.27	0.90	1.00

Note: Entries are square root of the average coherence at business cycle frequencies, as defined in (8), computed using the FSVAR described in Section 4. The factual (panel (a)) and counterfactual (panel (b)) scenarios are the same as in Table 10.

counterfactual exercise, however, are the same as those drawn from Table 10: under the counterfactual scenario, average business cycle coherences increase by an average of 0.07. In general, findings based on the contemporaneous correlations and the average coherence will be different. As it happens, however, the cross-country lead-lag relations evidently are modest, so these different measures give similar results.

Alternative Assumptions for Identifying the Factors. The model used for the counterfactual exercises reported above identifies the scale of the factors by assuming that the columns of Γ had unit length. Table 15 reports results for two alternative assumptions. In the first alternative, factor 1 has a unit impact on the US and factor 2 has a unit impact on France. In the second alternative, factor 1 has an average unit impact on English-speaking countries, and factor 2 has an average unit impact on Euro-zone countries. Table 15 summarizes the changes in the standard deviation of four-quarter growth rates (as in Table 9(a)) averaged across all of the series, and in average changes in correlations associated the common shock variance (that is, the difference in the elements Table 10(b) and Table 10(a)).

The first row of Table 15, shows the results for the baseline specification. Changing the 1984–2002 common shock variance to its 1960–1983 value leads to an average increase in the standard deviation of four-quarter growth rates of

TABLE 15. Summary of sensitivity to factor identification assumptions.

Factor identifying restrictions	$\sigma_{60-83,84-02} - \sigma_{84-02,84-02}$	$\sigma_{60-83,60-83} - \sigma_{84-02,84-02}$	$\mathrm{Cor}_{60-83,84-02} - \mathrm{Cor}_{84-02,84-02}$
$\lvert\Gamma_1\rvert = \lvert\Gamma_2\rvert = 1$	0.57	1.36	0.09
$\Gamma_{\mathrm{US},1} = \Gamma_{\mathrm{France},2} = 1$	1.13	1.80	0.15
$\Gamma_{\mathrm{EngSpeaking},1} = \Gamma_{\mathrm{EuroZone},2} = 1$	0.63	1.41	0.11

Notes: This table summarizes results for different assumptions used to identify the common factors. The assumptions are shown in the first column of the table. The first row shows results for the benchmark model used in the paper: the scale of the factor is determined by assuming that each column of factor loading has unit length. Alternative assumptions for the scale are shown in the last two rows. The second column of the table shows the average increase in the standard deviation of four quarter growth rates for 1984–2002 using common shock variances from 1960–1983. The third column shows the corresponding average increase using common and country-specific shocks from 1960–1983. The final column shows the average increase in the pair-wise correlation for 1984–2002 using common shock variances from 1960–1983. $\Gamma_{\mathrm{EngSpeaking}}$ is the average factor loading for Canada, the UK, and the US; $\Gamma_{\mathrm{EurZone}}$ is the average factor loading for France, Germany and Italy.

0.57 percentage points over the 1984–2002 sample period. Changing both the common and country-specific shock variances leads to an average increase of 1.36 percentage points. The average 1984–2002 pairwise correlation increases by 0.09 when 1960–1983 common shock variances are used in place of the actual common shock variances. Results for the alternative assumptions are shown in the next two rows, and they are similar to the baseline specification. The scale normalization $\Gamma_{\mathrm{US},1} = \Gamma_{\mathrm{France},2} = 1$ yields somewhat larger point estimates for the change in the variance of the common component, leading to a larger counterfactual increase in the standard deviation of GDP growth rates and a larger increase in the correlations. However, the standard errors (not reported) of the point estimates with this normalization are substantially larger than the baseline specification.

7. Discussion and Conclusion

These empirical results suggest four broad conclusions. First, although there has not been a general increase in international synchronization among G7 business cycles, there appear to have been important changes, in particular the emergence of two groups, one consisting of Euro-zone countries and the other of English-speaking countries, within which correlations have increased and across which correlations have decreased. Over this period, cyclical movements in the UK became less correlated with Euro-zone countries and more correlated with North American countries. Although the estimated magnitudes of the changes in these correlations are large from a macroeconomic perspective, the individual country-pair correlations and their changes are imprecisely estimated.

Second, common international shocks have been smaller in the 1980s and 1990s than they were in the 1960s and 1970s. According to the FSVAR, this declining volatility of common G7 shocks is the source of much of the observed moderation in individual country business cycles. Moreover, this moderation of common G7 shocks is responsible, in a mechanical sense, for the failure of business cycles to become more synchronous as one might expect given the large increase in trade over this period: had world shocks been as large in the 1980s and 1990s as they were in the 1960s and 1970s, international cyclical correlations would have increased.

Third, the Japanese experience is in many ways exceptional. For the other G7 countries, volatility generally decreased or at least stayed constant in the 1990s, but it increased in the 1990s in Japan. During the 1980s and 1990s, cyclical fluctuations in Japanese GDP became almost detached from the other G7 economies, with domestic shocks explaining almost all of the cyclical movements in Japanese GDP. This finding is consistent with Asian trade being increasingly important for the Japanese economy and with the domestic nature of the economic difficulties Japan experienced in the 1990s.

Fourth, a robust finding is that, however measured, persistence of disturbances—both reduced-form innovations and structural shocks—has increased in Canada, France, the UK. In those countries, a shock of a given magnitude would result in more cyclical volatility today than 30 years ago.

This analysis has focused on documenting the changes in the magnitudes of shocks and their effects. An important next step is sorting out the reasons for these changes and their implications for economic policy.

Appendix

Quarterly real GDP series were used for each of the G7 countries for the sample period 1960:1–2002:4. Unfortunately the data are not of uniform quality. A consistent series over the entire sample period did not exist for Canada, France, and Italy, and in these cases two series were spliced. The table below gives the data sources and sample periods for each series used. Abbreviations used in the source column are (DS) DataStream, (DRI) Data Resources and (E) for an internal OECD series from Dalsgaard, Elmeskov, and Park (2002). Some components of GDP are available only on an annual basis for some countries in the early part of the sample, and the OECD uses interpolation method to distribute these series over the quarter. See Dalsgaard, Elmeskov, and Park (2002), Doyle and Faust (2005), and OECD (2001, 2003) for a more detailed discussion of these problems.

Annual population values were interpolated to quarterly values using log-linear interpolation. The source for the annual series is given in the table. Consistent with the GDP data, population data from Germany are from West Germany prior to 1991 and for unified Germany from 1991.

There were three large outliers in the quarterly growth rates of real per capita GDP (France 1968:2–1968:3 and Germany 1991:1). These values were replaced with the series-specific full sample median growth rate.

TABLE A.1. Data sources.

Country	Series name	Source	Sample period
		Quarterly Real GDP	
Canada	cnona017g	OECD (DS)	1960:1 1960:4
	cngdp. . .d	Statistics Canada (DS)	1961:1 2002:4
France	frona017g	OECD (DS)	1960:1 1977:4
	frgdp. . .d	I.N.S.E.E. (DS)	1978:1 2002:4
Germany	bdgdp. . .d	Deutsche Bundesbank (DS)	1960:1 2002:4
Italy		OECD (E)	1960:1 1969:4
	itgdp. . .d	Istituto Nazionale di Statistica (DS)	1970:1 2002:4
Japan	jpona017g	OECD (DS)	1960:1 2002:4
UK	ukgdp. . .d	Office for National Statistics (DS)	1960:1 2002:4
US	gdpq	Department of Commerce (DRI)	1960:1 2002:4
		Annual Population	
Canada	tpopcan	OECD (DS)	1959–2002
France	tpopfra	OECD (DS)	1959–2002
Germany	tpopwgm	OECD (DS)	1959–1990
	topogma	OECD (DS)	1991–2002
Italy	tpopita	OECD (DS)	1959–2002
Japan	tpopjpa	OECD (DS)	1959–2002
UK	tpopukd	OECD (DS)	1959–2002
US	tpopusa	OECD (DS)	1959–2002

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