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World Business Cycles under Fixed and Flexible Exchange Rates

THE PURPOSE OF THIS PAPER is to examine by cross-spectral methods the dynamic interrelationships between innovations in monthly industrial production indices in a set of economies during the time period 1963 to 1986. More specifically, the paper attempts to shed light on the issue to what extent, and in what frequency bands, output fluctuations have been correlated during the recent periods of fixed and flexible exchange rates.

An often-heard argument before the introduction of the current system of managed exchange rates was that exchange rate flexibility would greatly reduce macroeconomic interdependence across countries. Modern stochastic macroeconomic theory, however, asserts that economic disturbances and policy are in general transmitted across countries, although the channels of transmission and the exact way in which economies respond to foreign shocks may depend on the exchange rate regime.[1] In an interesting recent article, Flood and Hodrick (1986) have convincingly argued that the variability of output may well be higher during a regime of fixed exchange rate than during a regime of flexible exchange rates, once one recognizes that fixed exchange rate regimes are never truly fixed. This conclusion is striking since it contrasts so sharply with the pre-1973 standard wisdom.

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[1]See Flood and Marion (1982) for a particularly detailed study of transmission under alternative exchange rate regimes. Recent work extending the real business cycle literature to open economies is also relevant (see Cantor and Mark 1986 and Stockman and Svensson 1987).

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While measurement of the degree of correlation of output movements across countries is thus of interest for the literature on the international transmission of economic disturbances under alternative exchange rate regimes, there are few studies that provide empirical evidence on international business cycles (see Saidi and Huber 1983; Swoboda 1983; Quah 1986). Additional work on this topic seems therefore warranted.

It should be clear that this paper does not attempt to test formally any hypothesis concerning the importance of the exchange rate regime for the behavior of international business cycles, but merely to establish some statistical regularities with respect to the output behavior in open economies during the recent periods of fixed and floating exchange rates. In order to test any such hypotheses, we would need to test restrictions derived from structural economic models econometrically. In the absence of any clear idea of what an estimable structural model would look like, it is difficult to derive such restrictions. It therefore appears more appropriate to pursue this investigation in an atheoretical manner, hoping to characterize the dynamic interactions between the variables in such a way as to generate hypotheses for future work. Naturally, this "measurement without theory" approach is subject to problems. In particular, it is not possible to attribute differences in the time series behavior of output in the two time periods directly to the change in the exchange rate regime that occurred in the early 1970s. In order to do so, we would require that the characteristics of the stochastic disturbances impinging on the economies in the two time periods were identical, and that the regime change in 1973 was, in some sense, "fundamental." Both of these conditions are problematic. The occurrence of common external disturbances in the 1970s in the form of the two oil shocks suggests that the first condition is violated, and the recurrent devaluations during the Bretton Woods system and the systematic intervention in foreign exchange markets since 1973 as well as the establishment of the EMS cast severe doubts on the second.

The paper is organized as follows. In the following section, the methodology is discussed. In addition, we test for the presence of stochastic trends, and examine the variability of the business cycles in individual economies. In section 2, the estimated coherences are displayed and commented upon. The third section offers some concluding remarks.

1. PRELIMINARIES

The theory of cross-spectral analysis is well documented in the statistical literature and an extensive discussion is superfluous here.[2] Briefly, the central idea is to decompose a time series into a finite number of orthogonal components, each associated with cyclical movements at a particular frequency or periodicity. The coherence at a given frequency is the squared correlation coefficient of the components of the two time series associated with this frequency. It is thus a measure of

[2]Good introductions can be found in Fuller (1976) and Sargent (1979).

how strongly correlated two time series are at different frequencies, and is bounded between zero and unity. The reasons for the choice of spectral analysis as research methodology instead of other atheoretical methods is that the notion of a world business cycle can be given an appealing definition in the frequency domain. Sargent (1979) characterizes business cycles in closed economies as the phenomenon of high pair-wise coherences in the frequency band corresponding to periodicities of three to five years. As pointed out by Saidi and Huber (1983), Sargent's definition is appropriate also for world business cycles. In what follows, world business cycles are defined as the high coherence in the business cycle frequency band of innovations in industrial production indices across countries. An additional reason for the use of cross-spectral analysis is that it is plausible that the correlations of output movements vary across frequencies. In particular, it is plausible that the coherences are strongest in the low frequency band. If so, estimated correlation coefficients—which are measures of correlations averaged across frequencies—would tend to underestimate the importance of comovements in economic activity across countries.

Detrending

As the original time series are all nonstationary, and theoretical spectra for nonstationary time series are not well defined, the issue of how to detrend the time series arises. There exists by now a long literature on optimal detrending of macroeconomic time series and the serious consequences of using inappropriate techniques.[3] Briefly, the central issue is whether the source of the nonstationarity is deterministic or stochastic trend growth. Given that inappropriate detrending induces important bias in the results, we first tested for the presence of a stochastic trend. Nelson and Plosser (1982) have shown that it is possible to test for the type of nonstationarity by estimating the equation.

$$y_t = \alpha + \beta t + \rho y_{t-1} + A(L)\Delta y_{t-1} + \epsilon_t \tag{1}$$

where, in our case, y_t denotes the log of industrial production, Δ is the first difference operator ($\Delta y_t = y_t - y_{t-1}$), and $A(L)$ is a polynomial in the lag operator. They argue that a test of the joint hypothesis that $\beta = 0$ and $\rho = 1$ is a test for the presence of stochastic growth, in which the data should be detrended by differencing.

Equation (1) was estimated using monthly industrial production data spanning the time period 1963 to 1986 for Belgium, Canada, France, Germany, Italy, the Netherlands, Norway, Sweden, the United States, and also for an European OECD index and a Total OECD index.[4] Since it is possible that the trend behavior of the series changed between the relatively stable fixed exchange rate period and the relatively volatile flexible exchange rate period, the sample was split in two, with the

[3]See the discussions in Beveridge and Nelson (1981), Chan, Hayya, and Ord (1977), Nelson and Kang (1981), and Nelson and Plosser (1982).

[4]The data come from the IMF's IFS tape and the OECD's Main Economic Indicators, and are available on request.

TABLE 1

TESTS FOR STOCHASTIC TRENDS

Model: $y_t = \alpha + \beta t + \rho y_{t-1} + A(L)\Delta y_{t-1}$ + seasonals + ϵ_t

H_0: $\beta = 0$, $\rho = 1$

Country Time Periods	F-Statistics 63:9–73:2/73:3–86:3	MSL for Q-Statistic[a] 63:9–73:2/73:3–86:3
Belgium[b]	1.16/3.91	.159/.217
Canada	2.87/4.00	.754/.072
France	3.00/4.03	.995/.637
Germany	1.94/2.38	.118/.756
Italy	1.57/5.47	.996/.981
Netherlands[c]	5.76/4.08	.776/.977
Norway	4.08/5.54	.963/.114
Sweden	2.98/.667	.117/.952
United States	1.90/3.49	.343/.818
OECD Europe[d]	1.91/6.74	.250/.810
OECD Total[d]	2.43/6.37	.578/.471

NOTES: The polynomial $A(L)$ is given by $= a_1L + \ldots + a_6L^6$, where L denotes the lag operator. The number of observations is 115 for the first time period and 157 for the second. According to Dickey and Fuller (1981, p. 1063, Table VI), the critical value for the F-test for 100 observations (250 observations) at the 95 percent level is 6.49 (6.34). The degrees of freedom for the Q-statistics are 30 for the first period and 36 for the second period.
a) Marginal significance level for the Q-Statistic.
b) First period was estimated with 8 lags in $A(L)$, the sample period is 1963:10–1973:2, and the nobs is 113. The second period was estimated with 14 lags.
c) First period was estimated with 12 lags in $A(L)$, the sample period is 1964:2–1973:2 and the nobs is 109.
d) Second period was estimated with 8 lags in $A(L)$.

breakpoint given by the date of the switch to a system of managed exchange rates. The first time period, 1963:2–1973:2, is referred to as the fixed exchange rate period, while the latter time period, 1973:3–1986:3, is referred to as the flexible exchange rate period. Since the test statistics appear unaffected by the presence of seasonal dummies, only the results for the regressions with seasonal dummies are presented in Table 1. Dickey and Fuller (1981, p. 1063, Table VI) report critical values for the F-test for 100 observations as 6.49 and for 250 observations as 6.34. Since the number of observations is 115 for the first time period and 157 in the second, the relevant critical value for the F-test is somewhere between these two critical values. As can be seen, in only one case, OECD Europe for the flexible rate period, is the hypothesis of stochastic growth rejected, while in another, OECD Total for the flexible rate period, the test statistic is in the range of the critical value. Given that we perform 22 tests at the 5 percent level, the few rejections suggest that the series are subject to stochastic growth, in which case it is appropriate to render the series stationary by differencing.

Analysis of Variance

Before estimating the cross-spectra, the variance of the log first differences of the industrial production indices were examined for the two time periods. As is revealed in Table 2, the variance is typically higher for the flexible exchange rate period than for the fixed rate period.[5] An obvious explanation for this phenomenon is the presence of common supply shocks in the latter sample period. In addition, it

[5]There are large outliers in the French data set for 1963:3–4 and 1968:5–7. The reported variance is the variance for the errors from a filtering regression in which Δy_t was regressed on a constant and five dummies for the above months. The variance of the raw data is 2.369.

TABLE 2

VARIANCE $\times\ 10^3$ OF Δy_t

Country Time Periods	Fixed Rate Period 1963:2–1973:2	Flexible Rate Period 1973:3–1986:3
Belgium	1.131	1.611
Canada	.117	.190
France	.171[a]	.341
Germany	.348	.341
Italy	.625	.980
Netherlands	.214	.672
Norway	1.245	1.436
Sweden	.738	1.363
United States	.060	.100
OECD Europe	.112	.110
OECD Total	.040	.058

NOTE: a) Using the adjusted data as explained in footnote 5.

appears that industrial production in the smaller, more open, economies is more variable. This may be so for two reasons. First, larger economies are presumably more diversified than smaller economies. A shock that is sectoral in a large economy may well be aggregate in a smaller economy. Thus, large economies could be expected to be subject to less variability than smaller economies. Second, more open economies may be subject to more foreign economic disturbances.[6]

In order to shed some light on the determination of the variability of industrial production, some simple regressions were run.[7] To test the hypothesis that the variability has increased from the fixed to the flexible exchange rate period, we regressed the variance of industrial production on a constant, ten country dummies and a time dummy. As can be seen in Table 3, regression I, the time dummy is statistically significant, as are seven of the unreported country dummies. Thus, there is evidence that industrial production was more variable in the flexible exchange rate period than in the fixed rate period. In order to test the effects of "openness" and "size," we next introduced measures of these variables into the regression. For "openness" the ratio of the sum of exports and imports to GNP in 1968 for the first time period and 1979 for the second period were used. For "size," GNP of 1968 and 1979, measured in terms of 1980 U.S. dollars using 1980 exchange rates, were used.[8] The results are displayed as regression II. As can be seen, all the t-statistics are insignificant, presumably because of the few number of observations. In order to conserve degrees of freedom by dropping insignificant country dummies, we performed an F-test of the hypothesis that the four country dummies whose t-values were less than 0.5 in absolute value were, as a group, zero. As the hypothesis was not

[6]This argument can be formalized as follows. Suppose that the level of domestic industrial production, y, is determined by both domestic, d, and foreign factors, f, and that the impact of foreign factors depend on some "openness" factor, δ, i.e., $y = (1 - \delta)d + \delta f$. The variance of domestic industrial production is thus given by $\mathrm{Var}(y) = (1 - \delta)^2\mathrm{Var}(d) + \delta^2\mathrm{Var}(f) + 2\delta(1 - \delta)\mathrm{Cov}(d,f)$. Thus, an increase in the degree of "openness" may or may not increase the variability depending on the relative sizes of the variances and the covariance. If the variances of the two factors are similar and if the covariance is positive, which empirically seems reasonable, an increase in the openness will increase the variability of domestic industrial production.

[7]I am grateful to an anonymous referee for suggesting this section.

[8]The data is from the OECD's *National Accounts: Main Aggregates*, Vol. 1, 1960–1985, Paris, 1985.

TABLE 3

VARIANCE REGRESSIONS

	Model: $\text{Var}_{i,t} = \alpha_i + \beta_t + \gamma z_{i,t} + \theta y_{i,t} + \epsilon_t$ $i = 1, \ldots, 10;\ t = 0, 1$				
No.	β	γ	θ	Adj. R^2	D.o.f.
I.	.218 (3.220)			.906	10
II.	.130 (1.035)	.946 (1.549)	$-.136\text{E}-03$ (-1.637)	.936	8
III.	.147 (2.451)	.802 (5.078)	$-.124\text{E}-03$ (-6.785)	.946	12

NOTES: $z_{i,t}$ denotes the "openness" variable, $y_{i,t}$ the "size" variable, and t-statistics are in parentheses. All equations are estimated with a constant and ten country-specific dummies in regressions I and II. Equation III was arrived at by dropping the country dummies in Equation II with t-values less than 0.5 (i.e., Belgium, Italy, Norway, and Sweden). A formal F-test that these dummies are zero yields $F(4, 8) = 0.534$, MSL $= 0.715$. The constant and the remaining six country dummies in Equation III all have t-statistics larger than 4 in absolute value.

rejected, the equation was reestimated without these dummies.[9] The results are presented as regression III. While the time dummy remains significant, the other variables—including the unreported constant and country dummies—are also significant. The results indicate that larger economies experience less variability, presumably because they are more diversified, and that countries more engaged in international trade tend to experience larger output fluctuations. Given the few observations, the results should obviously be interpreted with caution. Nevertheless, the results are in striking conformity with the hypothesis.

2. CROSS-SPECTRAL ANALYSIS

In this section we estimate cross-spectral densities in order to shed some light on two issues. First, is there a world business cycle, defined as high coherences in the business cycle frequency band of the innovations in the industrial production indices under study? Second, have output movements in individual countries become more or less correlated during the recent period of managed floating exchange rates?

One estimation problem that arises concerns the choice of "reference country." A first possibility is to estimate the coherences between individual countries. However, as pointed out by Saidi and Huber (1983), this may be inappropriate. Although output movements in two economies may be strongly correlated with each other, they may be uncorrelated with aggregate business cycles. Thus, such comparisons do not necessarily capture output fluctuations across a set of economies. A second possibility is to estimate the cross-spectra between an individual economy and an aggregate index. However, by construction, an aggregate index includes the indices for the individual economies. Estimating the coherences between the aggregate index and one of its components, thus, may lead to overestimates of the coherences. To avoid this problem, one possibility is to construct for each country a "rest-of-the-world" index. Here we take a more modest approach. The coherences are first

[9] The results for the test is $F(4,8) = 0.534$, MSL $= 0.715$.

estimated with respect to the U.S. index. This is of interest given the dominant role of the U.S. economy in the world economy, and especially in light of the prevalent European view that macroeconomic developments in the European economies are intimately linked to economic events in the United States. Second, the coherences vis-á-vis the Total OECD index are also estimated.

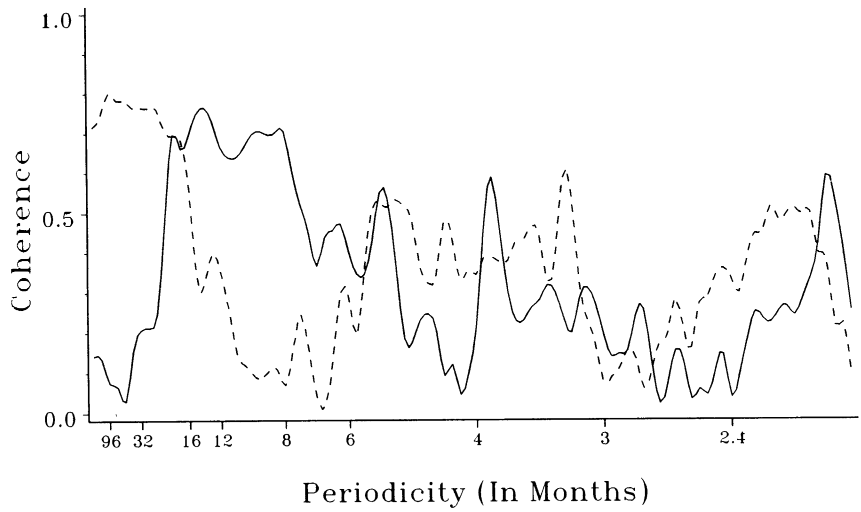

FIG. 1. Germany—United States.
Solid line indicates fixed rate period, dashes indicate flexible rate period.

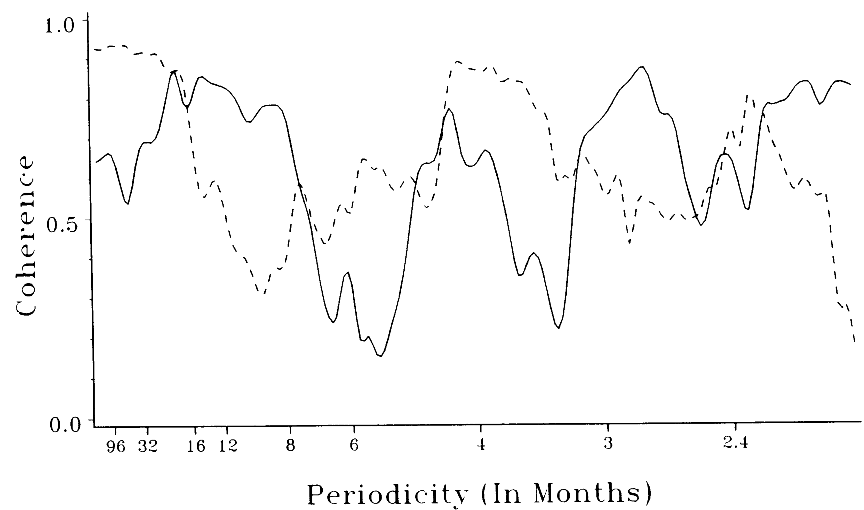

FIG. 2. Germany—Total OECD.
Solid line indicates fixed rate period, dashes indicate flexible rate period.

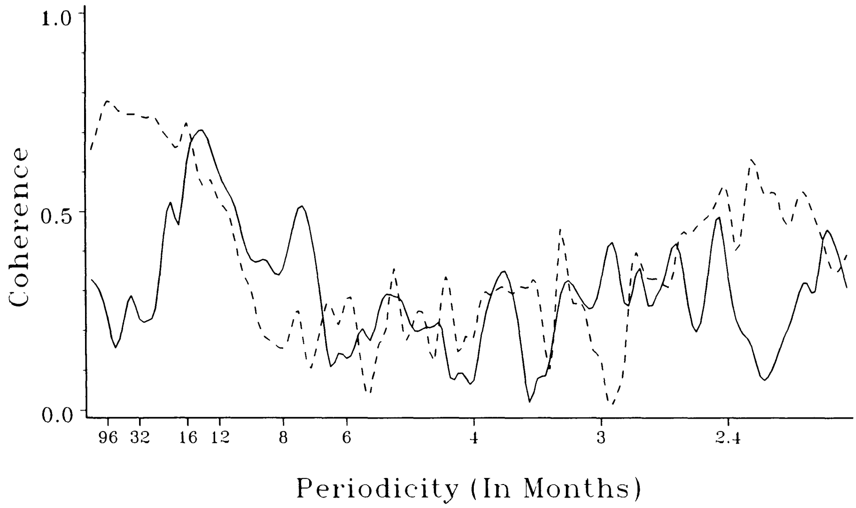

FIG. 3. OECD Europe—United States.
Solid line indicates fixed rate period, dashes indicate flexible rate period.

For space reasons, only some of the estimated coherences (which are representative) are displayed in Figures 1–3.[10] Additional coherences are tabulated in Table 4 for some particularly interesting frequencies.[11] Exact analytical confidence intervals cannot be given. However, coherences in excess of approximately 0.47 (0.56) are significantly different from zero at the 5 (1) percent level.[12] Although the coherences are somewhat unstable, several stylized facts are readily apparent.

First, comparing the fixed and flexible exchange rate periods, we note that the coherences are typically much higher during the latter time period. This finding, which corroborates the findings in Quah (1986) and Swoboda (1983), confirms the regression results discussed earlier. The additional insight provided by the use of the spectral techniques is that the coherences vary across frequency and time period. In the fixed rate period, to the extent that the coherences are significant, they are typically significant in the frequency band corresponding to periodicities of 8–16 months and, to some extent, the business cycle frequency band centered around periodicities of 48 months. In the flexible exchange period, however, the coherences are typically only significant in the low and in the business cycle frequency bands.

[10]The cross-spectra are estimated for the two time periods, 1963:2–1973:2 and 1973:3–1986:3, on the residuals from two filtering regressions in which Δy_t is regressed on seasonal dummies to remove any remaining seasonal effects. The cross-spectra are computed on 121 observations for the first period, and 157 observations for the second period, for 190 ordinates, and the cross-spectra are smoothed using a Daniell window with width 13. This window takes a moving average, with weights summing to unity, of the entries of the periodogram around each ordinate. To reduce window leakage, a trapezoidal taper is applied to the first and last 30 observations of the unpadded part of the series.

[11]Plots of all the coherences are available from the author on request.

[12]Fuller (1976, p. 315) demonstrates that a test for the hypothesis that the coherence, c, at a given frequency equals zero is given by the statistic $K = (4dc^2)/(2(1 - c^2))$, where $d = (q - 1)/2$ and where q is the width of the spectral window. K is approximately distributed as $F_{2,4d}$. Solving for c gives the critical values for the coherence quoted in the text.

TABLE 4

ESTIMATED COHERENCES

Belgium

	Against United States		Against Total OECD	
Months[a]	Fixed[b]	Flexible[c]	Fixed[b]	Flexible[c]
96	.10	.61**	.68**	.80**
64	.07	.63**	.61**	.83**
48	.09	.61**	.54*	.81**
32	.26	.64**	.69**	.82**
24	.28	.64**	.60**	.81**
16	.53*	.80**	.62**	.90**
12	.64**	.81**	.71**	.82**
8	.37	.16	.53*	.47*
6	.50*	.17	.22	.41
4	.29	.16	.72**	.64**

Canada

	Against United States		Against Total OECD	
Months[a]	Fixed[b]	Flexible[c]	Fixed[b]	Flexible[c]
96	.93**	.91**	.76**	.88**
64	.91**	.92**	.76**	.88**
48	.87**	.93**	.76**	.88**
32	.58**	.89**	.72**	.85**
24	.51*	.89**	.69**	.84**
16	.31	.82**	.58**	.80**
12	.48*	.54*	.72**	.52**
8	.09	.46	.47*	.44
6	.59**	.46	.57**	.47*
4	.74**	.39	.64**	.28

France

	Against United States		Against Total OECD	
Months[a]	Fixed[b]	Flexible[c]	Fixed[b]	Flexible[c]
96	.18	.73**	.47*	.89**
64	.19	.72**	.45	.90**
48	.21	.71**	.44	.89**
32	.39	.69**	.51*	.89**
24	.53*	.66**	.62**	.85**
16	.41	.64**	.38	.83**
12	.51*	.26	.39	.40
8	.72**	.26	.06	.47*
6	.46	.46	.36	.30
4	.20	.40	.51*	.35

Germany

	Against United States		Against Total OECD	
Months[a]	Fixed[b]	Flexible[c]	Fixed[b]	Flexible[c]
96	.07	.80**	.66**	.93**
64	.06	.78**	.59**	.93**
48	.03	.78**	.54*	.93**
32	.21	.76**	.69**	.92**
24	.25	.76**	.73**	.91**
16	.69**	.57**	.79**	.70**
12	.67**	.35	.83**	.53*
8	.70**	.07	.74**	.42
6	.40	.28	.33	.51*
4	.21	.34	.64**	.87**

Italy

	Against United States		Against Total OECD	
Months[a]	Fixed[b]	Flexible[c]	Fixed[b]	Flexible[c]
96	.32	.56*	.13	.76**
64	.34	.50*	.21	.74**
48	.22	.51*	.15	.75**
32	.20	.56*	.21	.78**
24	.15	.55*	.28	.76**
16	.43	.59**	.23	.75**
12	.67**	.47*	.52*	.57**
8	.65**	.15	.37	.07
6	.42	.07	.27	.27
4	.57**	.44	.31	.53*

TABLE 4 (*Continued*)

ESTIMATED COHERENCES

Netherlands

	Against United States		Against Total OECD	
Months[a]	Fixed[b]	Flexible[c]	Fixed[b]	Flexible[c]
96	.17	.83**	.53*	.94**
64	.26	.82**	.47*	.93**
48	.26	.82**	.39	.92**
32	.22	.77**	.47*	.88**
24	.19	.75**	.51*	.85**
16	.47*	.71**	.67**	.77**
12	.42	.32	.70**	.46
8	.34	.19	.54*	.35
6	.55*	.21	.55*	.34
4	.22	.26	.39	.76**

Norway

	Against United States		Against Total OECD	
Months[a]	Fixed[b]	Flexible[c]	Fixed[b]	Flexible[c]
96	.02	.37	.08	.44
64	.08	.43	.21	.55*
48	.22	.29	.38	.42
32	.34	.12	.57*	.30
24	.28	.13	.59**	.31
16	.70**	.20	.82**	.34
12	.65**	.14	.54*	.33
8	.29	.18	.14	.23
6	.09	.22	.14	.20
4	.19	.22	.13	.39

Sweden

	Against United States		Against Total OECD	
Months[a]	Fixed[b]	Flexible[c]	Fixed[b]	Flexible[c]
96	.14	.30	.69**	.38
64	.16	.31	.64**	.39
48	.31	.36	.69**	.40
32	.43	.29	.35	.36
24	.42	.30	.13	.37
16	.18	.24	.36	.35
12	.31	.35	.49*	.41
8	.52*	.36	.66**	.40
6	.21	.43	.55**	.18
4	.36	.22	.63**	.68**

OECD Europe

	Against United States		Against Total OECD	
Months[a]	Fixed[b]	Flexible[c]	Fixed[b]	Flexible[c]
96	.23	.78**	.81**	.94**
64	.16	.76**	.76**	.94**
48	.22	.75**	.73**	.93**
32	.23	.74**	.81**	.93**
24	.26	.74**	.85**	.93**
16	.62**	.72**	.92**	.92**
12	.60**	.53*	.90**	.81**
8	.35	.16	.79**	.62**
6	.13	.28	.67**	.63**
4	.07	.18	.91**	.91**

NOTES: a) Periodicity in months.
b) Fixed exchange rate period 1963:2–1973:2.
c) Flexible exchange rate period 1973:3–1986:3.
* (**) Denotes significance on 5 (1) percent level.

These results suggest that output movements in individual countries are correlated in the business cycle frequency band. Thus, there is evidence for a world business cycle. Furthermore, the importance of the international component is higher during the recent period of managed exchange flexibility than during the earlier period of fixed exchange rates. This is, as mentioned earlier, probably partially due to the occurrence of common external shocks in the latter sample period. However, another possible conclusion which is supported by the earlier regression tests is that increases in trade flows and the removal of capital controls have made national economies more integrated and more susceptible to external shocks. The results also suggest that studying the importance of world business cycles by estimating correlations of output movements across countries leads to underestimates of the importance of international factors since correlation techniques implicitly weigh all frequency bands equally.

Second, comparing the coherences estimated with respect to the United States and the Total OECD index, we note that with the exception of Canada the latter coherences are stronger. This is partially due to the earlier discussed overestimate that results when comovements between an index and a component of that index are estimated. However, the drastic differences for Belgium, Germany, the Netherlands, and Sweden, which presumably are too small to affect the Total OECD index, suggest that this difference is due to the presence of a European component. This conclusion is supported by the estimated coherences in Table 4 between the European OECD index and the U.S. index (which also is displayed in Figure 3) and between the European OECD index and the Total OECD index. As can be seen, the coherences between Europe and the United States are lower than the coherences between the European and the Total OECD indices. Thus, the Total OECD index contains more of a European component than the U.S. index, which explains why the European economies are more strongly correlated with the Total OECD index.

A third interesting conclusion is that Norway and, in particular, Sweden seem less affected than the other economies by international developments. Of several possible hypotheses for this result, one is that this is due to these countries' trading patterns: neither is a member of the Common Market, both are geographically somewhat peripheral, and for both raw materials are relatively important exports. Another hypothesis is that the large government sector in these countries makes them less sensitive to external disturbances.

3. CONCLUSIONS

The main results of this empirical study of multicountry output movements under fixed and flexible exchange rates are clear. First, industrial production indices are subject to stochastic trend growth. The data should thus be rendered stationary by differencing. Second, the variances of the monthly growth rates are typically higher in the flexible exchange rate period than in the fixed exchange rate period. The variances also are statistically significantly related to the degree of openness and to national income: more open economies tend to experience more variability of industrial production while richer, more diversified economies experience less. Third, output movements have been correlated across countries under both exchange rate regimes, and there is evidence of a world business cycle. Furthermore, the comovements in output have been more important, particularly in the business cycle frequency band, during the recent years of managed exchange rate flexibility.

Data for this paper are available from the JMCB *editorial office.*

LITERATURE CITED

Beveridge, Stephen, and Charles R. Nelson. "A New Approach to Decomposition of Economic Time Series into Permanent and Transitory Components with Particular Attention to Measurement of the 'Business Cycle'." *Journal of Monetary Economics* 7 (March 1981), 151–74.

Cantor, Richard, and Nelson C. Mark. "The International Transmission of Real Business Cycles." Unpublished paper, 1986.

Chan, K. Hung, Jack C. Hayya, and J. Keith Ord. "A Note on Trend Removal Methods: The Case of Polynomial Regression versus Variate Differencing." *Econometrica* 45 (April 1977), 737–44.

Dickey, David A., and Wayne A. Fuller. "Likelihood Ratio Statistics for Autoregressive Time Series with a Unit Root." *Econometrica* 49 (July 1981), 1057–72.

Flood, Robert P., and Robert J. Hodrick. "Real Aspects of Exchange Rate Regime Choice with Collapsing Fixed Rates." *Journal of International Economics* 21 (November 1986), 215–32.

Flood, Robert P., and Nancy Peregrim Marion. "The Transmission of Disturbances under Alternative Exchange-Rate Regimes with Optimal Indexing." *Quarterly Journal of Economics* 97 (February 1982), 43–66.

Fuller, Wayne A. *Introduction to Statistical Time Series*. New York: Wiley, 1976.

Huber, Gerard, and Nasser H. Saidi. "Business Cycles under Alternative Exchange Rate Regimes." Unpublished paper, 1982.

Nelson, Charles R., and Heejoon Kang. "Spurious Periodicity in Inappropriately Detrended Time Series." *Econometrica* 49 (May 1981), 741–51.

Nelson, Charles R., and Charles I. Plosser. "Trends and Random Walks in Macroeconomic Time Series: Some Evidence and Implications." *Journal of Monetary Economics* 10 (September 1982), 139–62.

Quah, Danny. "How Are International Business Cycles Also Alike." Unpublished paper, 1986.

Saidi, Nasser, and Gerard Huber. "Postwar Business Cycles and Exchange Rate Regimes." Unpublished paper, 1983.

Sargent, Thomas J. *Macroeconomic Theory*. New York: Academic Press, 1979.

Stockman, Alan C., and Lars E. O. Svensson. "Capital Flows, Investment and Exchange Rates." *Journal of Monetary Economics* 19 (March 1987), 171–201.

Swoboda, Alexander K. "Exchange Rate Regimes and European-U.S. Policy Interdependence." *IMF Staff Papers* 30 (March 1983), 75–102.