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Stock returns and real activity in the G-7 countries: did the relationship change during the 1980s?

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Abstract

Several empirical studies show that a substantial fraction of the changes in growth rates of real activity can be explained by lagged aggregate stock return variations in the U.S. as well as in other G-7 countries from the 1950s to the 1990s. However, the results presented in Binswanger.[International Review of Economics and Finance 9 (2000) 387] indicate that this traditionally strong relation has disappeared in the U.S. in the early 1980s. This paper shows that a similar breakdown occurred in Canada, Japan and in an aggregate economy consisting of the four European G-7 countries. The results provide evidence in favor of the hypothesis that speculative bubbles during the 1980s and 1990s were an international phenomenon.

JEL classification: E44, G12

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1. Introduction

During the 1980s and 1990s stock markets boomed in the U.S. but also in many European countries, while in Japan there was a strong boom until 1990. A number of fundamental explanations have been offered in order to explain these recent stock market booms (see Balke & Wohar, 2001; Carlson, 1999; Carlson & Sargent, 1997; Heaton & Lucas, 2000; Kopcke, 1997; McGrattan & Prescott, 2000 for recent surveys), but several authors have also argued that the

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recent stock price movements cannot be explained by fundamental factors and that they are the result of irrational exuberance or speculative bubbles (see, e.g., Binswanger, 1999; Shiller, 2000).

Several methods have been proposed to investigate whether the recent stock price movements have indeed been governed by fundamental factors. One possible approach is to analyze whether stock returns contain significant information about subsequent growth rates of real activity over the 1980s and 1990s as done in Binswanger (2000). According to the discounted-cash-flow valuation model stock prices should reflect investors' expectations of future real economic activity. The fundamental value of a firm's stock will equal the expected present value of the firm's future payouts (dividends). And future payouts must ultimately reflect real economic activity as measured by industrial production or GDP (see, e.g., Morck, Shleifer, & Vishny, 1990; Shapiro, 1988), which can be interpreted as proxies of corporate earnings (Choi, Hauser, & Kopecky, 1999). Consequently, stock prices should lead measures of real activity as stock prices are built on expectations of these activities as long as stock price movements are related to fundamentals.

The results presented in Binswanger (2000) suggest that the traditionally strong relation between stock returns and subsequent growth rates of real activity in the U.S. (see Barro, 1990; Fama, 1990; Schwert, 1990) has disappeared in the early 1980s. Since then, the stock market does not lead real economic activity as regressions fail to establish any significant relation between stock returns and growth rates of industrial production or of GDP. This structural break in regressions of stock returns on growth rates of real economic activity coincides with the start of a prolonged stock market boom that has dominated the stock market from the early 1980s till 2000. Therefore, the results provide empirical evidence in favor of non-fundamental explanations of the stock market booms over the 1980s and 1990s, as it is difficult to reconcile the breakdown with fundamental explanations of the recent stock market boom.This paper contributes to the existing literature by investigating whether the breakdown in the relation between stock returns and growth rates of real economic activity in the United States can also be found in the other G-7 countries (Canada, France, Germany, Italy, Japan, U.K.). The results presented in Choi et al. (1999), who analyze data from 1957 to 1996, suggest that, same as in the U.S., there used to be a significant relation between stock returns and subsequent growth rates of real economic activity in all other G-7 countries with the exception of Italy.[1] Finding a breakdown of this relation in the other G-7 countries in the early 1980s would therefore support the hypothesis that speculative bubbles during the 1980s and 1990s were an international phenomenon that affected all major economies.

However, there are reasons why a breakdown may not be as easily detected as in the U.S. First, some G-7 economies are small as compared to the U.S. economy and many of the large companies that are included in the domestic stock indices partially produce goods and services abroad. Therefore, investors' expectations of future payouts are traditionally less related to the expected development of domestic real activity. Second, the magnitude of exports and imports in relation to GDP in European economies as well as in Canada is a lot higher than in Japan and the U.S. and, therefore, foreign disturbances may weaken the strength of the association between domestic stock returns and domestic real activity.

For these reasons it may prove to be more difficult to derive a conclusion concerning a possible breakdown in the relation between the stock market and real activity in these countries.

Consequently, following the suggestion of Canova and De Nicolo (1995), we also construct an aggregate stock price index as well as aggregate measures of real activity for the four European G-7 countries. Because there are strong trading patterns among these countries and the economies are closely linked to each other, we expect the relationship between stock returns and real activity to be stronger on an aggregate level that consists of the largest European economies. Moreover, there is empirical evidence about increasing convergence among European stock markets (Rangvid, 2001).

The paper is organized as follows. Section 2 describes the variables used in the following regressions and explains the choice of the regression periods. Section 3 presents the results of unit root tests and cointegration tests which motivate the use of OLS regressions as well as of a vector error correction model for an investigation of the relation between stock returns and growth rates of real activity. These results are presented in Section 4 and compared to each other. Section 5 additionally presents CUSUM tests which are based on the calculation of recursive residuals using the same OLS regressions as in Section 4. Section 6 concludes.

2. Variables and sample periods

In this paper we run regressions with data from all G-7 countries. The data are from the International Financial Statistics of the IMF and consist of the aggregate stock price indices, seasonally adjusted industrial production indices, consumer price indices and seasonally adjusted nominal GDPs for the G-7 countries. As results may be sensitive with respect to the chosen real-activity variable we use industrial production as well as GDP in the following tests. The nominal stock price indices and GDP are converted into real data by dividing by the consumer price index for each country. All of the following tests use log levels of stock prices, industrial production and GDP. Growth rates are the log differences of quarterly observations and real stock returns are continuously compounded quarterly real returns. All data series run from 1960 to 1999.

The aggregate G-7 European real stock returns are constructed as an average of real stock returns in France, Germany, Italy and the U.K. weighted by market capitalization in 1993 US$ as in Canova and De Nicolo (1995).[2] The aggregate GDP measure is the sum of GDP of the four European G-7 countries. Finally, we use the European industrial production index provided by OECD as the proxy for aggregate industrial production in the G-7 European countries.[3]

In this paper we concentrate on tests using quarterly observations. Results in Fama (1990) as well as in Binswanger (2000) suggest that monthly stock returns possess little explanatory power for subsequent growth rates in real activity. The explanation offered by Fama (1981, 1990) is that information about a certain production period is spread over many previous periods. Therefore, short horizon returns only explain a fraction of future production growth rates but this fraction gets larger the longer is the time horizon of returns. Consequently, evidence concerning the relation between real stock returns and real economic activity mainly comes from regressions using quarterly and annual observations. However, we do not use annual observations because some of the investigated subsample periods are rather short and we would be forced to use overlapping annual observations in those regressions. And international comparisons of regressions using overlapping observations based on R^2-statistics and F-statistics could be misleading because the error term will be serially correlated and lead to biased estimates of these statistics.

In the following regressions we first test, whether stock prices lead real economic activities in each country over the whole sample period, which ranges from 1960 to 1999. This sample period is similar to the one investigated by Choi et al. (1999), who run their regressions with data ranging from 1957 to 1996. Then, we run separate regressions for recent boom periods in each country and test whether the results are significantly different over these periods if compared to the results for the whole sample and for a subsample covering the period before the recent stock market boom.

Real stock prices were increasing in all countries since 1983 but the actual patterns of real stock price development vary quite a lot among the G-7 countries. The boom is most prevalent in the U.S., where real stock prices steadily increased between 1983 and 1999 with only a short interruption in 1987. Of course, Japanese stock prices grew even more during the 1980s but they continuously declined again since 1990 when the boom suddenly came to an end, which makes Japan a rather special case.

As real stock prices started rising in all countries in 1983 with the exception of Italy (where they started rising in 1985), we take the first quarter of 1983 as the starting point of the current stock market boom and divide the 1960–1999 sample in two subsamples which cover the time periods from 1960 to 1982 and 1983 to 1999, respectively. However, because the results of regressions over the 1983–1999 subsample could be driven by the stock market crash in 1987, results probably look different if we exclude this event from the subsample representing the recent stock market boom. Therefore, we also run regressions over a 1989–1999 subsample.

3. Testing for cointegration

In order to test for a possible cointegrating relationship between stock prices and real activity we have to find out whether the time series under investigation are actually non-stationary and to what degree they are integrated if the null hypothesis of non-stationary cannot be rejected. According to the augmented Dickey–Fuller unit root tests all variables are $I(1)$ and, therefore, non-stationary in levels but stationary in first differences.[4] Consequently we can test for cointegration, which we do by using Johansen's VAR-based cointegration test as well as the two-step procedure proposed by Engle and Granger (1987). We present the results of both tests because they can lead to different results as, for example, outlined in Dickey, Jansen, and Thornton (1994, p. 36), and it is still controversial, which of the tests is more appropriate (see, e.g., Miyao, 1996; Timmermann, 1995). However, most of the previous studies rely on only one test procedure. The results in Choi et al. (1999), Kwon and Shin (1999) and Rapach (2001) are based on the Engle–Granger test, while Cheung and Ng (1998) and Nasseh and Strauss (2000) use the Johansen method.

Johansen's VAR-based cointegration test uses a vector error correction specification including k lags of the following form:

$$\Delta z_t = a + bd_t + \Pi z_{t-1} + \sum_{i=1}^{k-1} \Gamma_i \, \Delta z_{t-i} + \varepsilon_t \tag{1}$$

where z_t is the 2×1 vector of the $I(1)$ variables, which are either real stock prices and industrial production or real stock prices and real GDP; d_t is a time trend; ε_t is a white noise error; and the vectors and matrices of parameters (a, b, Π, Γ_i) are dimensioned conformably.

The two-step Engle–Granger cointegration test is based on the estimation of the following regressions:[5]

$$x_t = a + bp_t + \mu_t, \tag{2}$$

$$\Delta\mu_t = c\mu_{t-1} + \sum_{i=1}^{k} d_i\, \Delta\mu_{t-i} + \theta_t, \tag{3}$$

where x_t and p_t represent the log levels of real activity and real stock prices, respectively, k is the number of lags included, and μ_t and θ_t are error terms. Eq. (3) tests for stationarity of the error term μ_t of Eq. (2) and shows the autoregressive representation which is used in the augmented Dickey and Fuller test. The results of the Johansen test and the Engle–Granger test are shown in Table 1.

There is mixed evidence concerning cointegration between real stock prices and industrial production and between real stock prices and real GDP at quarterly frequencies. Both, the Johansen test and the Engle–Granger test indicate that the null hypothesis of no cointegration cannot be rejected for stock prices and industrial production in the U.S., Canada, France, Germany and the U.K. at the 5% level,[6] while it can be rejected in Japan and the G-7 European

Table 1
Cointegration tests (sample: 1960–1999)

Country	Real stock prices and industrial production		Real stock prices and real GDP	
	Johansen test	Engle–Granger test	Johansen test	Engle–Granger test
U.S.A.	14.31 (4)	−2.86 (4)	14.98 (3)	−2.25 (1)
Canada	12.59 (3)	−2.37 (4)	12.65 (4)	−2.22 (4)
France	11.91 (4)	−2.07 (1)	27.71** (2)	−4.59** (1)
Germany	9.38 (4)	−2.71 (4)	9.07 (5)	−3.36* (1)
Italy	16.28* (2)	−0.96 (6)	15.89* (2)	−0.96 (6)
Japan	31.72** (5)	−2.94* (2)	27.55** (3)	−2.89* (1)
U.K.	8.02 (3)	−2.47 (4)	6.66 (8)	−2.04 (2)
G-7 Europe	18.32* (4)	−3.39* (4)	11.97 (4)	−3.65** (1)

(*,**)Rejection of the null hypothesis of no cointegration at the 5 (1) % level.

The Johansen test assumes a linear deterministic trend in the data. The test statistic shown in the table is the likelihood ratio (trace statistic). The optimal lag length has been determined according to the Akaike information criterion from an unrestricted VAR, which includes the variables of the model expressed in levels. The Akaike information criterion has been used rather than the Schwarz criterion because it usually suggests the inclusion of a larger number of lags and the Johansen test can be misleading if too few variables are included (see Maddala & Kim, 1998, pp. 214–220).

The test statistic shown in the Engle–Granger test columns is the augmented Dickey and Fuller (ADF) statistic using Eq. (3). The optimal lag length in the ADF test has been determined according to the Akaike information criterion. See Maddala and Kim (1998, pp. 77–78), or Stock (1994, p. 2781), for a discussion of different criteria of choosing the appropriate lag length in Dickey–Fuller tests.

The number of lags included in the test are given in parentheses.

aggregate.[7] For stock prices and real GDP both test procedures indicate that the null hypothesis of no cointegration cannot be rejected for the U.S., Canada and the U.K. at the 5% level while it can be rejected for France and Japan. For the other countries and the G-7 European aggregate the results are inconclusive as the two test procedures differ in their results. Moreover, neither the results of the Johansen test nor the results of the Engle–Granger test are very robust with respect to the chosen sample period and with respect to the number of lags included in the test. Overall, the results do not provide a firm basis to decide in favor or against cointegration in most countries. Therefore, we will run OLS regressions as well as estimate a vector error correction model in order to obtain robust results concerning the possible breakdown in the relation between stock prices and real activity in the early 1980s.

4. OLS regressions and a vector error correction model

In this section we present the results of the OLS regressions of quarterly growth rates of industrial production and of quarterly growth rates of real GDP on quarterly real stock returns as well as a vector error correction model including quarterly growth rates of industrial production or real GDP, quarterly real stock returns and an error correction term.

The estimated OLS equations are

$$g_t = a + \sum_{k=0}^{4} b_k r_{t-k} + \varepsilon_t \tag{4}$$

where g_t stands for the quarterly growth rate of real activity from $t-1$ to t, and r_{t-k} stands for the real stock return from $t-k-1$ to $t-k$. We include the same amount of lags of real stock returns as in the regressions presented in Fama (1990) and Binswanger (2000) in order to make our results directly comparable. All regressions are estimated for four different sample periods, which are 1960–1999, 1960–1982, 1983–1999, and 1989–1999. Table 2 shows the results of the regressions using the growth rate of industrial production as the dependent variable.[8]

The results for the U.S. confirm the results reported in Binswanger (2000) where the full sample ranged from 1953 to 1997. There is a breakdown in the relation between real stock returns and growth rates of real activity in the early 1980s as indicated by the results of the Chow breakpoint test. Stock returns clearly correlate with growth rates over the full sample and this relation is especially strong over the 1960–1982 subsample. However, the correlation is weak or absent if we test for the 1983–1999 as well as the 1989–1999 subsample as can be seen from the estimated values of the adjusted R^2- and the F-statistic. A result that is even more obvious if we test for a subsample that starts in 1984 instead of 1983 (results not reported) because regression diagnostics suggest that the structural break appeared in 1984 rather than in 1983 (Binswanger, 2000).

The results for Japan and the G-7 European aggregate economy are in accordance with the results for the U.S. The correlation between past stock returns and current growth rates of real activity is visible in the 1960–1982 subsample but absent in the 1983–1999 as well as in the 1989–1999 subsample and Chow breakpoint tests suggest a structural break in 1983. There also appears to be a significant relation in Canada, Germany and the U.K. from 1960 to 1982 based on the adjusted R^2-values and the results of F-tests, which, however, in some regressions can also be found for the 1989–1999 subsample or the 1983–1999 subsample. Chow breakpoint

Table 2
Regressions of quarterly growth rates of industrial production on quarterly real stock returns

Country		1960–1999	1960–1982	1983–1999	1989–1999
U.S.A	Adj. R^2	0.26	0.39	0.10	0.04
	S.E.	0.01	0.02	0.01	0.01
	F-statistic (*p*-value)	0.00	0.00	0.04	0.27
	Chow-test (*p*-value)	0.00			
Canada	Adj. R^2	0.23	0.36	0.06	0.18
	S.E.	0.02	0.01	0.02	0.01
	F-statistic (*p*-value)	0.00	0.00	0.11	0.03
	Chow-test (*p*-value)	0.49			
France	Adj. R^2	0.00	0.01	0.04	0.09
	S.E.	0.03	0.03	0.01	0.01
	F-statistic (*p*-value)	0.41	0.36	0.20	0.13
	Chow-test (*p*-value)	0.40			
Germany	Adj. R^2	0.08	0.20	0.00	0.32
	S.E.	0.02	0.02	0.02	0.01
	F-statistic (*p*-value)	0.01	0.00	0.43	0.00
	Chow-test (*p*-value)	0.05			
Italy	Adj. R^2	0.03	0.01	−0.00	0.13
	S.E.	0.03	0.03	0.02	0.02
	F-statistic (*p*-value)	0.21	0.21	0.44	0.08
	Chow-test (*p*-value)	0.34			
Japan	Adj. R^2	0.20	0.32	0.08	0.06
	S.E.	0.02	0.02	0.01	0.01
	F-statistic (*p*-value)	0.00	0.00	0.06	0.20
	Chow-test (*p*-value)	0.00			
U.K.	Adj. R^2	0.10	0.10	0.02	0.17
	S.E.	0.02	0.02	0.01	0.01
	F-statistic (*p*-value)	0.00	0.02	0.29	0.05
	Chow-test (*p*-value)	0.71			
G-7 Europe	Adj. R^2	0.06	0.20	−0.06	−0.08
	S.E.	0.01	0.01	0.01	0.01
	F-statistic (*p*-value)	0.02	0.00	0.94	0.81
	Chow-test (*p*-value)	0.00			

Adj. R^2 stands for the adjusted *R*-squared, S.E. stands for the standard error of the regression, *F*-statistic gives the significance level at which the null hypothesis that all coefficients are zero can be rejected. The Chow-test refers to the Chow breakpoint test and gives the significance level of the *F*-test at which the null hypothesis of no subsample instability can be rejected for the first quarter in 1983.

tests do not provide evidence for a structural break in these countries. Moreover, the relation between growth rates of real activity and stock returns never seemed to be very strong in France and Italy and, consequently, the lack of evidence for a structural break is not surprising as growth rates of real activity and past stock returns did not correlate neither before nor after 1983.

As cointegration between real stock prices and real activity is a possibility in most countries, we also estimate the following vector error correction model

$$\begin{bmatrix} g_t \\ r_t \end{bmatrix} = \begin{bmatrix} \sum_{k=1}^{4} a_{1k} g_{t-k} + \sum_{k=1}^{4} b_{1k} r_{t-k} + c_1 \rho_{t-1} + \varepsilon_{1t} \\ \sum_{k=1}^{4} a_{2k} g_{t-k} + \sum_{k=1}^{4} b_{2k} r_{t-k} + c_2 \rho_{t-1} + \varepsilon_{2t} \end{bmatrix} \tag{5}$$

where g_{t-k} stands for the quarterly growth rate of real activity from $t-k-1$ to $t-k$, r_{t-k} stands for the real stock return from $t-k-1$ to $t-k$, and ρ_t stands for the error correction term, which is the residual from the estimated cointegrating equation $x_t = \alpha + \beta p_t$ where x_t represents the log level of industrial production or real GDP and p_t stands for the log level of real stock prices.

In order to interpret the results of the vector error correction model we rely on similar test statistics as Choi et al. (1999).[9] In the context of this paper we are especially interested in the question whether the breakdown in the relation between past stock returns and current growth rates of real activity, that is suggested by the results of OLS regressions, can also be found if an error correction term is included in the model. Therefore we use a Wald test in order to find out about the joint significance of past stock returns in explaining current growth rates of real activity in a vector error correction model. The imposed restriction on Eq. (5) is

$$b_{11} = b_{12} = b_{13} = b_{14} = 0. \tag{6}$$

Table 3 reports the test statistic of the Wald test, which is distributed as χ^2, as well as the associated p-value, which is the level of significance at which the null hypothesis of the parameter restriction (Eq. (6)) can be rejected.

The results of the vector error correction model are similar to the results of the OLS regressions.[10] With the exception of France and Italy, past stock returns contain significant information about current growth rates of real activity in all countries and the G-7 European aggregate for the full sample and the 1960–1982 subsample where the null hypothesis of the Wald test can be rejected. However, the results are totally different for the 1983–1999 and 1989–1999 subsamples. Past stock returns only appear to be somewhat significant in the U.S. for both subsamples and in Germany for the 1989–1999 subsample. But in the case of the U.S. the significance of past stock returns disappears once we test for a subsample that starts in 1984 instead of 1983 (results not reported here).

Overall, the results of the vector error correction model support the conclusion based on the results from OLS regressions that there is a breakdown in the relation between real activity and stock returns in the U.S., Japan, the aggregate European economy because past stock returns are not significant any more in explaining growth rates of real activity in the recent subsamples. The breakdown seems to be an international phenomenon that is not restricted to the U.S. And again, there is no evidence for a breakdown in France and Italy. For Germany, the test results confirm that there is a significant relation between stock returns and real activity also in the 1989–1999 subsample, which, however, is absent in the 1983–1999 subsample.[11] The only major difference between the results reported in Tables 2 and 3 concerns the U.K., where the vector error correction model, contrary to the OLS regressions, indicates a breakdown.

The results presented in this section also suggest that the relationship between stock markets and real activity in the European G-7 countries is better understood if it is investigated on an

Table 3
Vector error correction model of industrial production and real stock prices

Country		1960–1999	1960–1982	1983–1999	1989–1999
U.S.A.	χ^2	46.84	29.60	10.22	10.04
	p-value	0.00	0.00	0.04	0.04
	Adj. R^2	0.44	0.44	0.45	0.36
Canada	χ^2	16.20	13.06	0.62	2.12
	p-value	0.00	0.01	0.95	0.71
	Adj. R^2	0.31	0.41	0.20	0.21
France	χ^2	7.17	8.51	5.99	3.93
	p-value	0.13	0.07	0.20	0.42
	Adj. R^2	0.08	0.07	0.12	0.22
Germany	χ^2	18.53	13.40	1.19	14.71
	p-value	0.00	0.01	0.88	0.01
	Adj. R^2	0.13	0.24	−0.01	0.31
Italy	χ^2	6.94	7.88	1.93	4.60
	p-value	0.14	0.10	0.75	0.33
	Adj. R^2	0.05	0.07	−0.01	0.03
Japan	χ^2	22.02	38.28	3.50	3.12
	p-value	0.00	0.00	0.48	0.54
	Adj. R^2	0.57	0.66	0.37	0.22
U.K.	χ^2	13.64	13.75	1.48	4.47
	p-value	0.00	0.01	0.83	0.35
	Adj. R^2	0.12	0.09	0.14	0.43
G-7 Europe	χ^2	24.92	26.78	1.64	2.51
	p-value	0.00	0.00	0.80	0.64
	Adj. R^2	0.28	0.36	0.11	0.32

χ^2 reports the Chi-square statistic of the Wald test and p-value reports the associated probabilities at which the null hypothesis can be rejected. Adj. R^2 stands for the adjusted R-squared and refers to the estimated vector error correction model. Changing the number of lags included in the vector error correction model does not significantly alter the test results. The estimated vector error correction model refers to Eq. (5) and includes four lags. In the Wald test restrictions are imposed on the vector error correction model as indicated in Eq. (6).

aggregate level (G-7 Europe) than if it is investigated for the single countries, which confirms the finding of Canova and De Nicolo (1995). The adjusted R^2 is higher in the regressions using data from G-7 Europe than in the regressions using data from the individual European countries in the OLS regressions (with the exception of Germany, Table 2) as well as in the vector error correction model (Table 3). Past stock returns seem to be less significant for the explanation of current growth rates of real activity if analyzed for the single countries than if analyzed for the aggregate European economy. This is most evident in the case of Italy where previous studies also have failed to find any significant relation, while there is mixed evidence about France in the existing literature (Cheung & Ng, 1998; Choi et al., 1999; Hassapis & Kalyvitis, 2002).

The comparatively weak relation between domestic real activity and national stock returns in the European G-7 countries (especially France and Italy) may be explained by the fact that the value of national stock markets is only partially linked to domestic real activity (Canova & De Nicolo, 1995). Many stocks included in the national stock price indices are more closely linked to foreign real activity in the other major European countries because of the strong trading patterns among these countries.

5. Recursive residuals and CUSUM tests

The test statistics of the OLS regressions presented in the last section concentrated on an in-sample analysis of the correlation between real stock returns and growth rates of real activity, which is an *ex post* property of the data. In this section we adopt an *ex ante* perspective and test for the stability of the out-of-sample predictive power by estimating recursive least squares based on the OLS Eq. (4). In recursive least squares the equation is estimated repeatedly, using ever larger subsets of the sample data. If there are k coefficients to be estimated in the b vector used in Eq. (8), then the first k observations are used to form the first estimate of b. The next observation is then added to the data set and $k+1$ observations are used to compute the second estimate of b. This process is repeated until all the sample points T have been used, yielding $T-k+1$ estimates of the b vector. At each step the last estimate of b can be used to predict the next value of the dependent variable. The one-step ahead forecast error resulting from this prediction, suitably scaled, is defined to be a recursive residual.

The CUSUM test as developed by Brown, Durbin, and Evans (1975) is based on the cumulative sum of the recursive residuals. Fig. 1 plots the cumulative sum together with the 5% critical lines for each country. The test finds parameter instability if the cumulative sum goes outside the area between the two critical lines. The CUSUM test is based on the statistic.

$$W_t = \sum_{t=k+1}^{T} \frac{w_t}{s} \qquad T = k+1, \ldots, N \tag{7}$$

where w_t is the recursive residual, s is the standard error of the regression fitted to all N sample points. The recursive residual w_t is defined as

$$w_t = \frac{y_t - x_t' b_{t-1}}{(1 + x_t'(X_{t-1}' X_{t-1})^{-1} x_t)^{-1/2}} \tag{8}$$

where X_t denotes the $t-1$ by k matrix of the regressors from period 1 to period $t-1$, and y_t the corresponding vector of observations on the dependent variable. These data up to period $t-1$ give an estimated coefficient vector, denoted by b_t. This coefficient vector gives a forecast of the dependent variable in period t. The forecast is $x_t' b_{t-1}$, where x_t' is the row vector of observations on the regressors in period t. The forecast error is $y_t - x_t' b_{t-1}$ and the forecast variance is $\sigma^2(1 + x_t'(X_{t-1}' X_{t-1}')^{-1} x_t)$.

We present results from CUSUM tests for OLS regressions using industrial production (Fig. 1).

The CUSUM test provides evidence for coefficient instability in the U.S., Japan and G-7 Europe, where W_t clearly moves outside the critical 5% line during the 1980s. These results support the finding of the OLS regressions shown in Table 2, where the Chow breakpoint test indicates a structural break between growth rates of both measures of real activity and stock

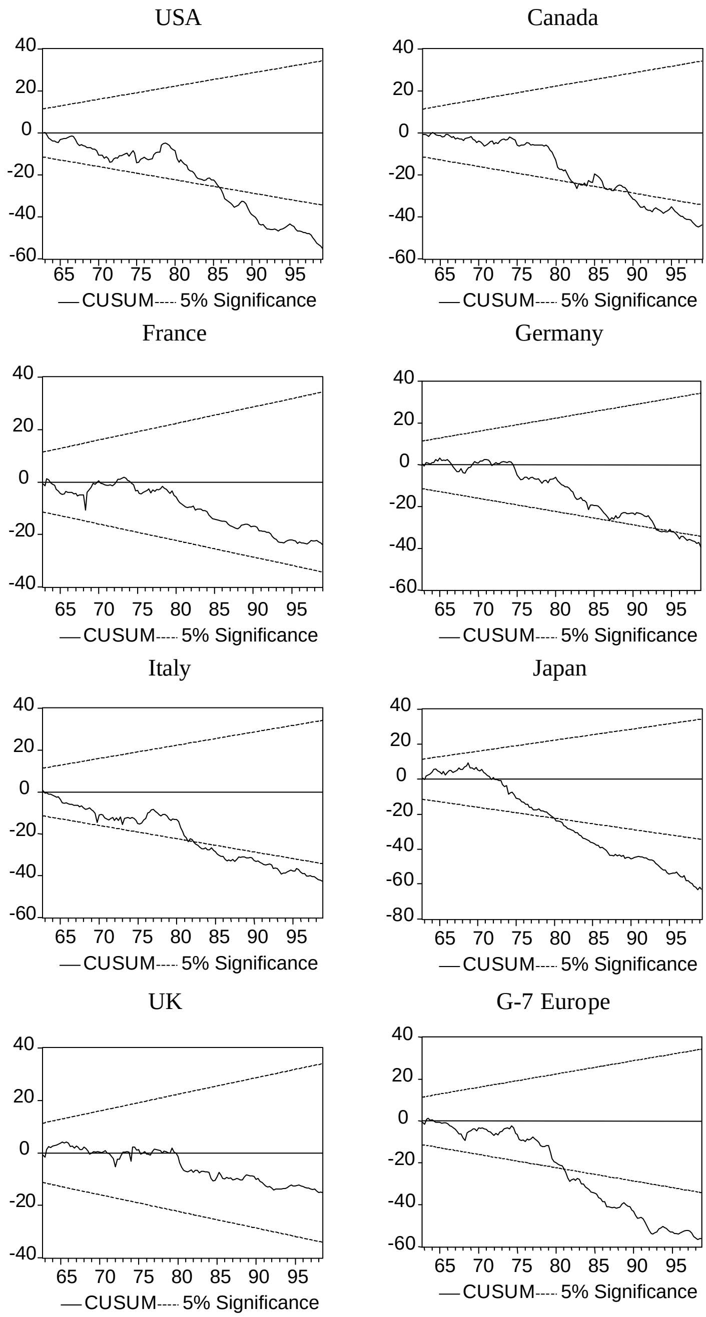

Fig. 1. CUSUM Tests for OLS regressions using industrial production as the independent variable. The test statistics were calculated by using expression (7).

returns for the U.S., Japan and G-7 Europe in 1983. However, the CUSUM test also provides some evidence for structural breaks in Canada and Italy. Furthermore, the CUSUM test indicates that the breakdown did not occur simultaneously in different areas but either before 1983 (Italy, Japan, G-7 Europe for industrial production) or after 1983 (U.S., Canada, G-7 Europe for real GDP). And indeed, the Chow breakpoint test can also detect a subsample instability in the OLS regressions using industrial production as the measure of real activity in Canada and Italy if one tests for a structural break during the years suggested by the CUSUM tests (1980 or 1981 in the case of Italy and 1985 or 1986 in the case of Canada) instead of in 1983. Therefore the results from the Chow breakpoint tests in the OLS regressions and from the CUSUM test are consistent as both tests suggest a structural break for Canada and Italy. However, in the case of Italy the structural break is not associated with an actual breakdown in the relation between stock returns and growth rates of real activity as OLS regressions as well as the vector error correction model fail to find any significant relationship before and after 1983.

Furthermore, the CUSUM test finds no coefficient instability in France and the U.K. This is in accordance with the results from Chow breakpoint tests presented in Table 2. And, same as the previous tests, the CUSUM test cannot provide a conclusive result as to whether there is coefficient instability in Germany.

6. Conclusion

In this paper we investigated whether the breakdown in the traditional relation between real stock returns and growth rates of real economic activity in the United States, that according to Binswanger (2000) occurred in the early 1980s, can also be found in the other G-7 countries (Canada, France, Germany, Italy, Japan, U.K.) as well as in the aggregate European economy consisting of the four European G-7 countries. The results from OLS regressions and the vector error correction model presented in Section 4 as well as the results from the CUSUM test presented in Section 5 all suggest that a similar breakdown occurred in Japan and in the aggregate European economy no matter whether industrial production or real GDP is used as the variable representing real activity. Since the 1980s stock markets do not lead real economic activity as regressions fail to establish any significant relation between stock returns and growth rates of real activity even if the 1987-episode is excluded from the tested samples. However, the results from the CUSUM test indicate that the breakdown did not occur exactly at the same date in different areas as there is some variation between the early and the late 1980s. A breakdown also seems to have occurred in Canada, although the results are not as clear as for the U.S., Japan and the aggregate European economy. The test results for France, Italy and the U.K. do not suggest a structural break while they are inconclusive for Germany.

The relationship between stock markets and real activity is stronger on an aggregate level (G-7 Europe) than on national levels (Germany is a possible exception) as can be seen from the results of the OLS regressions as well as from the results of the vector error correction model for the 1960–1982 period. This confirms the finding of Canova and De Nicolo (1995), who also report a closer association between stock returns and real activity on an aggregate level than for any of the individual countries. Therefore, evidence for a breakdown in the relation between stock returns and real activity cannot be found in the data for France, Italy and the U.K. but rather in the aggregated European data, where the hypothesis of a breakdown is confirmed by all tests.

A possible explanation of the observed breakdown of the relation of the stock market and real activity is the deviation of stock prices from fundamental values due to irrational exuberance or the emergence of speculative bubbles (Binswanger, 1999; Shiller, 2000). However, the existence or nonexistence of bubbles or fads in stock markets is a highly controversial issue because fundamental values are unobservable and, consequently, bubbles usually cannot be distinguished from unobserved fundamentals (Hamilton & Whiteman, 1985). If we accept the existence of bubbles as an explanation of the empirical findings in this paper, our results suggest the emergence of a speculative bubble in the U.S., Japan and in Europe in the early 1980s, which subsequently dominated the development of these stock markets over the 1980s and 1990s and weakened the relation between the stock market and real activity. The difference is, of course, that the Japanese bubble burst in 1990, while the American and European bubbles survived the 1990s and even gained further momentum over this decade. But also a bursting bubble may weaken the relation between the stock market and real activity as is evident from the case of Japan during the 1990s.

Of course, there are alternative potential explanations of the observed breakdown related to shocks to discount rates or variation in risk premia (see, e.g., Lee, 1998), which are not considered in this paper. However, the results in Binswanger (2000) also suggest that, at least in the U.S., the breakdown cannot easily be explained by time-varying risk premia. But no matter how the breakdown is explained, the results presented in this paper suggest that it is an international phenomenon that affected all of the major economic areas (U.S., Japan, Europe). The breakdown may also be associated with the increasing interdependence of major stock markets since the 1980s (especially since the 1987-episode, as shown in Longin & Solnik, 1995; Meric & Meric, 1997; Wu & Su, 1998; Rangvid, 2001) and further research is needed to uncover the possible interrelation between these findings and the results of this paper.

Notes

1. Further results from OLS regressions for the G-7 countries are presented in Canova and De Nicolo (1995), Aylward and Glen (2000) and Mauro (2000). But the results do not allow for general conclusions about the relation between stock returns and growth rates of real activity.

   A number of recent empirical studies has also investigated the relation between real activity and stock returns for several countries by using multiple equation analysis such as vector autoregressions (Lee, 1992, 1995; Gjerde & Saettem, 1999; Kwon & Shin, 1999; Groenewold, 2000; Rapach, 2001; Hassapis & Kalyvitis, 2002) or vector error correction models (Cheung & Ng, 1998; Nasseh & Strauss, 2000). As far as the G-7 countries are concerned, Hassapis and Kalyvitis (2002) find a strong relation between real stock price changes and growth rates of real activity from the 1950s to the mid 1990s in all G-7 countries with the exceptions of Italy and France. Nasseh and Strauss (2000) report significant long-run relationships between stock prices and industrial production in five European countries (France, Germany, Italy, The Netherlands, Switzerland and UK) testing for data sets from 1962 to 1995. And Cheung and Ng (1998) find a significant relation between stock returns and future GNP growth rates in Canada, Germany, Japan and the U.S. but not in Italy for data from 1957 to 1992.

2. As explained by Canova and De Nicolo (1995, p. 986), the results are not altered significantly if the date chosen to index market capitalization is changed because relative shares of these markets in total are approximately constant over time. This can be seen from the data on market capitalization of shares of domestic companies provided by the International Federation of Stock Exchanges (Tables 1.3B and V.3).
3. The weights for calculating the European industrial production index from the OECD are derived from gross domestic product originating in industry and the GDP purchasing power parity for 1995. Together, the four European G-7 countries' proportion of the European industrial production index is about two thirds.
4. Results are available from the author upon request.
5. The equation does not include a time trend because according to Hansen (1992) including a time trend results in under-rejecting the null hypothesis of no cointegration when it is false even if $\grave{\imath}_t$ in fact contains a deterministic trend.
6. Actually, the lack of cointegration between stock indices and industrial production or between stock prices and real GDP in some countries already indicates that a stable relationship between these variables probably does not exist and that a breakdown may have occurred between 1960 and 1999.
7. Our results partially differ from the results reported in Choi et al. (1999), who find a significant cointegrating relationship between real stock prices and industrial production in all of the G-7 countries based on the two-step Engle–Granger cointegration test. However, one has to bear in mind that the test results appear to be sensitive with respect to the chosen sample period and our sample period is different from the sample period tested in Choi et al. (1999).
8. Using GDP-growth rates instead of industrial production as the dependent variable leads to very similar results although the correlation between real stock returns and growth rates of real GDP appears to be less pronounced than the correlation between real stock returns and growth rates of industrial production (Results are available from the author upon request).
9. Choi et al. (1999) evaluate the significance of real stock returns in explaining current growth rates of industrial production by testing whether the b_{1k} and c_1 in (5) are different from zero if the first equation of (5) is estimated as a single equation error correction model.
10. Again, the results of regressions using growth rates of GDP instead of industrial production are very similar to the results presented in Table 3.
11. Further investigations are necessary to uncover the logic behind this rather peculiar result for Germany, which may be linked to the German unification.

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