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How important are fundamentals?—Evidence from a structural VAR model for the stock markets in the US, Japan and Europe

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Abstract

This paper presents a bivariate structural VAR model which includes growth rates of industrial production and stock prices. Analyzing data from 1960 to 1999 we find that real activity shocks only explain a small fraction of the variability in real stock prices in the US, Japan and an aggregate European economy since the early 1980s, while they explain a substantial proportion over the 1960s and 1970s in all areas. The results provide additional evidence for the existence of speculative bubbles over the 1980s and 1990s.

JEL classification: E44; G12

Keywords: Stock prices; Speculative bubbles; Structural vector autoregression

1. Introduction

Stock prices have seen an unprecedented rise during the 1980s and 1990s, particularly in the US. In many European countries and in Japan, prices also started to rise in the early 1980s leading to subsequent stock market booms that, with the exception of Japan, gained even further momentum over the 1990s. A number of fundamental explanations have been offered in order to explain these recent stock market booms (see Balke and Wohar,

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2001; Carlson, 1999; Carlson and Sargent, 1997; Heaton and Lucas, 2000; Kopcke, 1997; McGrattan and Prescott, 2000 for recent surveys), but several authors have also argued that the recent stock price movements cannot be explained by fundamental factors and that they are the result of speculative bubbles (see, for example, Binswanger, 1999; Shiller, 2000).

Furthermore, results from regressions presented in Binswanger (2000) suggest that there is a breakdown in the traditionally strong relation between stock returns and real economic activity in the US in the early 1980s. Such a breakdown implies the emergence of new sources of variation in stock returns during the stock market boom over the 1980s and 1990s, which are not explained by the traditional discounted cash flow valuation model according to which stock prices should lead measures of real activity (see, for example, Fama, 1990). This finding additionally supports the "bubble hypothesis",[1] as it is difficult to reconcile the breakdown with fundamental explanations of the recent stock market boom.

There are two main questions which we set out to answer in this paper and which have not been investigated so far. First, we want to find out whether the breakdown in the relation between stock returns and growth rates of real activity in the US in the early 1980s reported in Binswanger (2000) can also be found if we use a structural VAR (SVAR) model instead of single equation models. Thus, we are interested in a possible change in the relation between stock returns and growth rates of real activity if we compare the 1960s and 1970s to the 1980s and 1990s. Secondly, we set out to determine whether the breakdown in the relation between stock returns and real economic activity can also be found in Japan and Europe, or whether this is a unique feature of the US stock market. Of course, previous studies have been undertaken to examine the relation between stock returns and the growth rates of real activity or other fundamental variables in several countries by using the SVAR approach.[2] But none of the existing studies investigates whether this relation has changed during the recent stock market boom which began in the early 1980s. Looking for such a change is important because, if it can be identified, it will lend further credibility to nonfundamental explanations of the booming stock markets over the 1980s and 1990s.

The SVAR approach has become a popular tool in empirical investigations of stock prices as it allows analysis of the movements of stock prices in relation to various fundamental and nonfundamental shocks, which can be identified by imposing specific restrictions on an

[1] Some of the literature distinguishes between bubbles and fads. However, there is no general agreement concerning these terms. For example, Lee (1998) and Chung and Lee (1998), following Cochrane (1991), consider price deviations which slowly return to fundamental values as fads, whereas bubble price deviations are expected to continue until the bubble bursts. According to this definition, the empirical work of Lee (1998) and Chung and Lee (1998) identifies fads in the stock markets of the US, Hong Kong and Singapore, while it identifies bubbles in Japan and Korea. However, Shiller (1988) defines "a bubble as a fad if the contagion of the fad occurs through price". According to this definition bubbles are a subcategory of fads. In this paper, we will not make a distinction between bubbles and fads and instead interpret persistent deviations of stock prices from fundamental value as bubbles.

[2] Recent contributions in this field are Lee (1995a), Gjerde and Saettem (1999), Groenewold (2000), and Rapach (2001), who estimate SVAR models including stock prices and measures of real activity, and Lee (1995a, 1995b, 1998), Chung and Lee (1998) and Allen and Yang (2003) who estimate SVAR models including stock prices and dividends and/or earnings.

estimated VAR that includes stock prices and other variables that are supposed to indicate the change in market fundamentals (dividends, earnings or measures of real activity, interest rates, risk premia). Furthermore, aggregate stock prices can be decomposed into fundamental and nonfundamental components and the development of the fundamental component of stock prices can be simulated over various time periods. This is an advantage over single equation regressions, where it is difficult to interpret the results with respect to the question as to how much stock price movements are actually caused by changes in market fundamentals. For example, Fama (1990, p. 1090) leaves it to the reader to judge whether the identified explanatory power of the variables is good or bad news about market efficiency (implying that stock prices correctly reflect the underlying fundamentals on average).

In this paper we estimate a bivariate SVAR model which includes growth rates of industrial production and real stock prices. In order to identify structural shocks we will use a long-run restriction à la Blanchard and Quah (1989) that excludes long-run influences of pure innovations in stock prices on real activity. The idea behind this restriction is that permanent changes in stock prices can be caused by changes in fundamentals as well as nonfundamentals, while permanent changes in real activity can only be caused by changes in fundamentals. In our model, changes in fundamentals are caused by shocks to real activity, which may have permanent effects on real activity as well as on stock prices. Stock market shocks, which are the nonfundamental shocks in our model, may also temporarily affect real activity, but in the long-run they can only affect stock prices. However, the identified nonfundamental shocks may also include some changes in fundamental components not directly related to real activity (i.e. time-varying interest rates, time-varying risk premia) and for this reason we prefer to term the identified shocks *real activity shocks* (a major category of fundamental shocks) and *other shocks*, instead of using the terms *fundamental shocks* and *nonfundamental shocks*.

The rest of the paper is organized as follows: Section 2 outlines the bivariate SVAR model used in the subsequent tests. Section 3 presents the data as well as stationarity tests. In Section 4 we present the estimation results which include impulse responses, forecast error variance decompositions, and the real stock price historical decompositions. Section 5 concludes and interprets the findings of the paper.

2. The SVAR model and the decomposition of shocks into fundamental and non-fundamental components

In this paper we consider a two-variable model that consists of the first differenced log of real stock prices, s, and the first differenced log of industrial production x. Real economic activity may be expressed by industrial production or by real GDP. The majority of the previous studies concentrates on industrial production as the variable representing real economic activity and we will follow this tradition here because the relation between stock returns and growth rates of industrial production appears to be more significant than the relation between stock returns and growth rates of real GDP (Binswanger, 2001). The model is similar to the SVAR model estimated by Groenewold (2000) who analyses the relation between growth rates of real GDP and growth rates of the stock market in Australia.

The reduced form of the VAR model is represented by the bivariate system

$$\Delta Z_t = \begin{bmatrix} \Delta x_t \\ \Delta s_t \end{bmatrix} = \begin{bmatrix} a_{10} \\ a_{20} \end{bmatrix} + \begin{bmatrix} A_{11}(L) & A_{12}(L) \\ A_{21}(L) & A_{22}(L) \end{bmatrix} \begin{bmatrix} \Delta x_{t-1} \\ \Delta s_{t-1} \end{bmatrix} + \begin{bmatrix} e_{1t} \\ e_{2t} \end{bmatrix} \tag{1}$$

where a_{i0} are the parameters representing intercept terms, L is the lag operator with $L^i x_t \equiv x_{t-I}$, $A_{ij}(L)$ are polynomials in the lag operator L (for example, $A_{11}(L) = a_{11}(0) + a_{11}(1)L + a_{11}(2)L^2 + \cdots + a_{11}(p)L^p$) where p is the number of lags included in the VAR, and e_{1t}, e_{2t} are the observed error terms of the reduced VAR model (reduced form residuals or reduced form shocks). The error terms e_{1t}, e_{2t} are white noise disturbances which, however, will usually be correlated unless there are no contemporaneous effects between Δx_t and Δs_t.

Ignoring intercept terms we can write (1) in a more compact notation as

$$\Delta Z_t = A(L)L\,\Delta Z_t + e_t \tag{1a}$$

Given the fact that the time series Δx_t and Δs_t are both covariance-stationary and assuming that $A(L)$ is invertible, we can write

$$\Delta Z_t = [I - A(L)L]^{-1} e_t \tag{2}$$

which is the bivariate infinite order moving average representation (BMA) of (1).

As the residuals e_{1t}, e_{2t} are usually correlated, they cannot be structural innovations, which are supposed to be uncorrelated with each other. The unobservable structural innovations come from a VAR representation of the structural form (SVAR), which we suppose can be written as

$$B(L)\Delta Z_t = u_t \tag{3}$$

where $B(L)$ is a matrix of structural parameters derived by identifying restrictions and u_t is a vector of the uncorrelated white noise disturbances u_{1t}, u_{2t} which are the structural shocks or structural innovations.

If the matrix polynomial $[I - A(L)L]$ is invertible, so is the matrix polynomial and the SVAR can also be expressed as BMA

$$\Delta Z_t = [B(L)]^{-1} u_t = C(L) u_t \tag{4}$$

or, more precisely

$$\begin{bmatrix} \Delta x_t \\ \Delta s_t \end{bmatrix} = \begin{bmatrix} C_{11}(L) & C_{12}(L) \\ C_{21}(L) & C_{22}(L) \end{bmatrix} \begin{bmatrix} u_{1t} \\ u_{2t} \end{bmatrix} \tag{4a}$$

where $C_{ij}(L)$ are the infinite polynomials in the lag operator L.

The structural shocks u_{1t}, u_{2t} can be approximated by first estimating the finite-order reduced form VAR (1) (which is an approximation of the infinite-order VAR) and then transforming the reduced form residuals e_{1t}, e_{2t}. In order to make this transformation, it is necessary to impose restrictions on the structural model. As the structural shocks are supposed to be uncorrelated, the variance–covariance matrix of the structural shocks must be diagonal. Furthermore, without loss of generality, the standard deviations of the structural shocks are normalized to 1 leading to an orthonormalized BMA. Generally, making these assumptions yields $n(n + 1)/2$ restrictions, which in the case of a BMA amounts to three restrictions. However, at least n^2 independent restrictions on parameters of the structural form are needed to exactly identify the system. Thus, in an orthonormalized BMA we need just one additional restriction. This final restriction, which is the "structural" part of our SVAR, is a long-run restriction as proposed by Blanchard and Quah (1989).[3] Referring to (4a) we impose the restriction

$$C_{12}(L) = 0 \tag{5}$$

from which it follows that u_{2t} has no long-run impact on x_t.

Once the reduced form VAR (1) is estimated, the structural shocks can be obtained by using the restriction imposed on the SVAR as expressed by (5). As shown in Lee (1995b) the shocks u_{1t}, u_{2t}, can also be interpreted as fundamental and nonfundamental shocks if restriction (5) is imposed on the estimated VAR.[4]

Of course, our bivariate SVAR model cannot reveal all sources of variations in fundamentals.[5] Typically, macroeconomic VAR models also include interest rates in order to indicate changes in fundamentals related to changes in the discount rate. However, this would require further identifying restrictions and the results presented in Lee (1995a, 1998) suggest that time-varying interest rates do not help to explain stock price movements that cannot be related to changes in earnings or dividends (in the USA) or real activity (in Japan). However, there may also be variations in the discount rate caused by variations in excess stock returns (i.e. caused by changes in domestic or global risk premia), which Lee (1998) does find to be important in the USA. Finally, stock prices may also be affected by foreign influences such as foreign real activity. Therefore, the nonfundamental shocks in our model are likely to also include some fundamental changes and for this reason we term the identified shocks *real activity shocks* (fundamental shocks) and *other shocks*, which mainly consist of nonfundamental shocks but also include fundamental shocks not intrinsically related to domestic real activity.

The choice of a simple bivariate model instead of a VAR including further fundamental variables is motivated by the fact that we only need to impose one additional identifying restriction (5), which does not require a strong a priori assumption based on a specific

[3] Potential problems of this approach are outlined in Lippi and Reichlin (1993), or Faust and Leeper (1997), who concentrate on identification problems, and in Cochrane (1998), Rudebusch (1998), or Cooley and Dwyer (1998), who criticize the lack of robustness.

[4] According to traditional stock valuation models the fundamental stock price at t is determined by discounted expected future cash flows (for which expected future real activity is a proxy) conditional on information available at t. However, as shown in Lee (1995b), provided that we treat conditional expectations as equivalent to linear projections on information, we can use a lemma for covariance stationary stochastic processes proved in Hansen and Sargent (1980) that allows the expression of the expected value of future fundamental stock prices based on past fundamental shocks. And changes in the log of industrial production (our fundamental variable) Δx_t are a covariance stationary stochastic process in our model, where using restriction (5), $\Delta x_t = C_{11}(L)u_{1t}$.

[5] According to the discounted cash flow valuation model, these fundamentals are the expected present value of the firm's future cash flows and, as outlined by Fama (1990), there are three possible sources of variations in stock price fundamentals: (a) shocks to expected cash flows, (b) shocks to discount rates and (c) predictable return variation due to predictable variation through time in the discount rates (for example caused by the risk premium) that price expected cash flows.

economic theory. As soon as we estimate VARs including more variables, we have to impose $n(n-1)/2$ additional restrictions in order to identify the structural shocks. In the case of VARs which include measures of real activity, this requires assumptions concerning the validity of specific economic theories which are not always straightforward. Rapach (2001), for example, estimates a VAR for the USA that includes real stock returns, the growth rates of real GDP, the real interest rates and the inflation rate and imposes six additional long-run restrictions which are motivated by the natural rate hypothesis. However, the imposed structure is not innocuous. What the paper actually shows is how stock prices could be decomposed into various components if the natural rate hypothesis were true, which arguably represents a strong a priori assumption.

3. Data and unit root tests

The relation between growth rates of real activity and growth rates of real stock prices is analyzed for the USA, Japan and an aggregate European economy that consists of the four major European economies (France, Germany, Italy, United Kingdom). Analyzing the relation for an aggregate European economy was suggested by Canova and De Nicolo (1995) because there are strong trading patterns among these countries and the economies are closely linked to each other. This would suggest a weak relation between stock returns and real activity for the single European countries but a possibly strong relationship for the aggregate European economy, which is confirmed by the results shown in Canova and De Nicolo (1995) and Binswanger (2001). In each case we simulate the effects on real activity and stock prices of each of these two shocks. We then go on to use this identification scheme to decompose stock prices into two components, which also allows comparison of our results with those decompositions reported in the literature.

The data used in this paper are from the International Financial Statistics of the IMF and consist of the aggregate stock price indices, industrial production indices (seasonally adjusted) and consumer price indices for Japan, the US and the four major European economies from 1960 till 1999. The nominal stock price indices 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 and industrial production. Growth rates are the log differences of quarterly observations and real stock returns are continuously compounded quarterly returns. The aggregate European stock index is constructed as an average of real stock prices in France, Germany, Italy and the UK weighted by market capitalization in 1993 in US$ as in Canova and De Nicolo (1995). The results are not altered significantly if the date chosen to index market capitalization is changed because the relative shares of these markets in total are approximately constant over time.

Fig. 1 shows the development of the log levels of real stock prices and industrial production from 1960 to 1999. Real stock prices have increased in Europe, Japan and the USA since 1983 but the actual patterns of real stock price development vary considerably among the G-7 countries. The boom is most prevalent in the US, 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 have subsequently declined

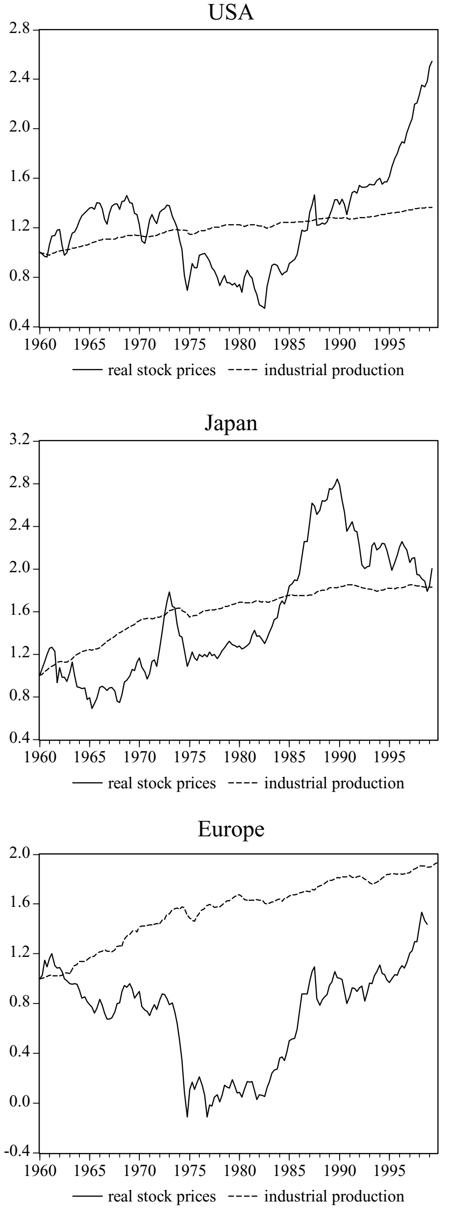

Fig. 1. Log levels of real stock prices and industrial production (normalized for 1960).

continuously since 1990 when the boom suddenly came to an end, which makes Japan a rather special case. A stock market boom over the period 1983–1999 can also be detected in Europe although it is less pronounced than in the US or in Japan. Therefore we identify the period 1983–1999 with the recent stock market boom.

Table 1
Annualized mean and standard deviation of growth rates of real stock prices and industrial production

Country	Mean		Standard deviation	
	Stock price (%)	Industrial production (%)	Stock price (%)	Industrial production (%)
1960–1982				
USA	−1.2	3.0	25.8	8.1
Japan	1.5	8.6	30.8	9.8
Europe	−3.8	2.8	28.2	6.8
1983–1999				
USA	11.0	3.6	22.7	3.9
Japan	3.9	2.1	34.4	6.0
Europe	8.9	2.0	26.5	3.7

In the following section we will test whether variance decompositions of stock prices in the SVAR model are significantly different over the recent stock market boom and present historical decompositions of stock prices over this period. Table 1 shows the annualized mean and standard deviation of growth rates of real stock prices and growth rates of industrial production for the period from 1960 to 1982 as well as for the period from 1983 to 1999. Looking at the first sample, the growth rate of real stock prices is negative (USA, Europe) or slightly positive (Japan). The growth rate of industrial production is around 3 percent in the US and Europe, while it is over 8 percent in Japan. However, things look completely different for the period 1983–1999. In the US and Europe, the growth rate of the stock market (11 percent and 8.9 percent, respectively) is more than three times the growth rate of real activity. Also in Japan, the growth rate of stock prices exceeds the growth rate of real activity, but stock price growth rates are not as high as in the US or Europe because they came down again over the 1990s. Furthermore, Table 1 also shows that the growth rate of stock prices is far more volatile than the growth rate of real activity. However, the volatility of the stock market was not higher during the period 1983–1999 than during the period 1960–1982 period.

In order to proceed with our empirical analysis, we also have to determine whether the time series under investigation are actually non-stationary and the degree to which they are integrated if the null hypothesis of non-stationary cannot be rejected. According to the results of the augmented Dickey–Fuller unit root test, as well as of the Phillips–Perron test, all variables (log levels) are I(1) and, therefore, non-stationary at their levels but stationary at their first differences.

The unit root tests provide the following results:

Country	Real stock prices		Industrial production	
	Log levels	Log differences	Log levels	Log differences
Augmented Dickey–Fuller test				
USA	0.07	−5.85[a]	−1.89	−5.43[a]
Japan	−2.49	−5.27[a]	−1.69	−6.34[a]
Europe	−1.31	−5.60[a]	−3.07	−6.45[a]

(*Continued*)

Country	Real stock prices		Industrial production	
	Log levels	Log differences	Log levels	Log differences
Phillips–Perron test				
USA	0.20	−9.12[a]	−2.34	−7.04[a]
Japan	−1.89	−9.28[a]	−1.90	−6.20[a]
Europe	−1.01	−9.52[a]	−2.50	−9.15[a]

[a] Rejection of the null hypothesis of non-stationarity at the 5 percent level. Both tests include an intercept and allow for a deterministic trend. In the Phillps–Perron test the lag truncation for the Bartlett kernel is set to 4.

4. Empirical results of the structural VAR model

From the preceding section we know that the growth rates of stock prices as well as of industrial production are stationary and, therefore, we are able to estimate a VAR including the growth rates of these two variables as well as a constant. In order to set the lag length we use four different criteria (final prediction error, Akaike information criterion, Schwarz information criterion, Hannan–Quinn information criterion) and estimate the model for lag lengths of 0–5. The results are displayed in Table 2. The final prediction error and the Akaike information criterion suggest the inclusion of more lags than the other two criteria in the case of the US and Europe, while all criteria suggest the inclusion of just one lag in the case of Japan. Therefore, we set the lag length at one for Japan. Furthermore, the lag length is set at three for the US and Europe as suggested by the final prediction error and the Akaike information criterion. The other two criteria would suggest the inclusion of only one lag, but t-statistics indicate that the second and third lag are also significant in explaining the current level of the growth rate of industrial production in the VAR.

The impulse responses of each variable to typical (one-standard-deviation) structural shocks for the full sample are presented in Fig. 2.

The growth rates of both industrial production and stock prices react positively to a real activity shock in Europe, Japan and the US. This is in accordance with the discounted cash flow valuation model, where an increase in real activity should also lead to a subsequent increase in stock prices through the effect on dividends. Furthermore, comparing the effect of real activity shocks on the growth rate of real activity and on the growth rate of stock prices in the first quarter after the shock, the effect on the growth rates of stock prices is always considerably larger than the effect on the growth rate on real activity. This probably reflects the fact that stock prices have fluctuated a lot more than real activity during the period 1960–1999. But the effect on the growth rate of stock prices rapidly declines and growth rates turn slightly negative after the second quarter. However, the accumulated (positive) effect on stock prices is still larger than the effect on real activity.

Other shocks have a positive effect on the growth rates of stock prices that is larger than the effect of real activity shocks in Europe, Japan and the US, indicating that other kinds of shocks are generally a more important determinant of stock prices than real activity

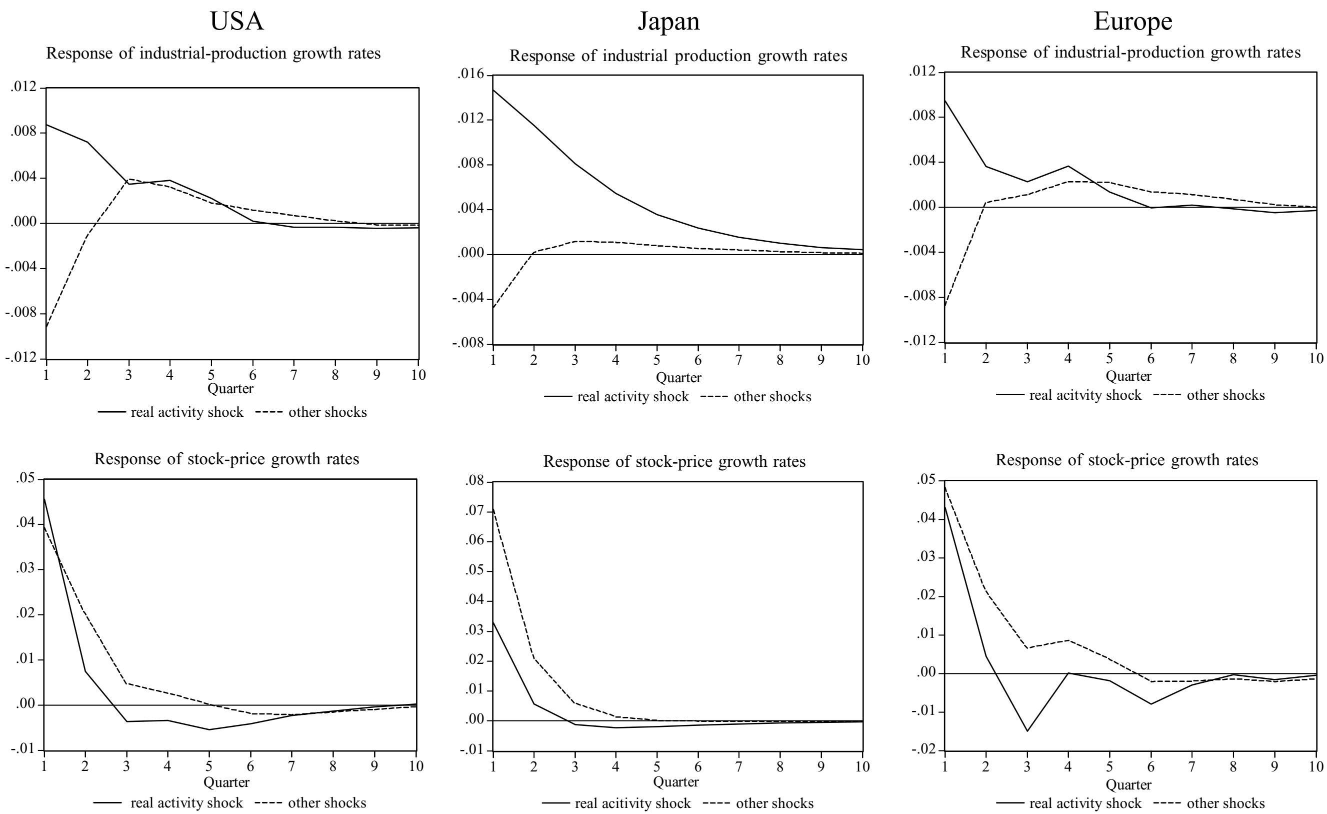

Fig. 2. Impulse responses to one-standard-deviation structural shocks (1960–1999).

Table 2
Selection of lag length

Lag	FPE	AIC	SC	HQ
USA				
0	1.06E-06	−8.081599	−8.041278	−8.065217
1	6.24E-07	−8.611959	−8.490994[a]	−8.562813[a]
2	6.43E-07	−8.581312	−8.379704	−8.499402
3	6.16E-07[a]	−8.623781[a]	−8.341531	−8.509108
4	6.19E-07	−8.620409	−8.257516	−8.472971
5	6.48E-07	−8.574258	−8.130722	−8.394057
Japan				
0	2.83E-06	−7.100528	−7.060206	−7.084146
1	1.36E-06[a]	−7.835394[a]	−7.714430[a]	−7.786248[a]
2	1.39E-06	−7.813038	−7.611431	−7.731128
3	1.43E-06	−7.785056	−7.502806	−7.670383
4	1.46E-06	−7.762218	−7.399325	−7.614780
5	1.52E-06	−7.721737	−7.278201	−7.541536
Europe				
0	1.03E-06	−8.111973	−8.069750	−8.094815
1	8.30E-07	−8.325698	−8.199030[a]	−8.274224[a]
2	8.33E-07	−8.323008	−8.111895	−8.237218
3	7.80E-07[a]	−8.388587[a]	−8.093028	−8.268480
4	8.07E-07	−8.354130	−7.974126	−8.199707
5	8.32E-07	−8.323982	−7.859532	−8.135242

[a] Indicates lag order selected by the criterion (each test at 5% level). FPE: final prediction error; AIC: Akaike information criterion; SC: Schwarz information criterion; HQ: Hannan–Quinn information criterion.

shocks. Due to our imposed long-run restriction, the accumulated effect of these shocks on real activity must be equal to zero. However, in the short run, we can observe a negative effect on real activity growth rates in all cases. A similar negative effect is also found in the four-variable SVAR model for the US estimated by Rapach (2001) and, therefore, it does not seem to be a statistical artifact of our bivariate SVAR model.

The explanation for this negative effect offered in Rapach (2001) is that an unexpected rise in stock prices that is not caused by real activity shocks will encourage investors to shift funds into the stock market from the bond market. Bond yields will then increase as bond prices fall leading to higher interest rates, which in turn should have a negative effect on investment and, therefore, real activity. However, it seems doubtful whether this effect can be observed as early as one quarter later, as the impulse response functions suggest. The negative effect may also be the result of investors‘ changing sentiments (i.e. irrational exuberance). If they suddenly become more optimistic about the future development of the stock market they will buy stocks instead of financing new investment projects, which in the short run has a negative effect on real activity.

Calculating the value for the impulse response functions over the period 1983–1999 (results are not shown here) does not lead to substantially different results but the effects of real activity shocks on stock prices become generally weaker as compared to the effect of other shocks on real activity. However, the stock price forecast error variance decompositions

Table 3
Stock price forecast error variance decompositions for different periods

Quarters-ahead	1960–1999, percent of variance attributable to		1960–1982, percent of variance attributable to		1983–1999, percent of variance attributable to	
	Real activity shocks	Other shocks	Real activity shocks	Other shocks	Real activity shocks	Other shocks
USA						
1	57.32	42.68	75.53	24.47	18.99	81.01
2	52.17	47.83	67.70	32.30	18.97	81.03
3	52.05	47.95	67.72	32.28	20.25	79.75
4	52.10	47.90	68.06	31.94	20.96	79.04
5	52.44	47.56	68.64	31.36	21.16	78.84
10	52.57	47.43	68.61	31.39	21.16	78.84
Japan						
1	17.67	82.33	36.85	63.15	1.38	98.62
2	16.89	83.11	35.78	64.22	2.26	97.74
3	16.81	83.19	35.96	64.04	2.72	97.28
4	16.87	83.13	36.15	63.85	2.91	97.09
5	16.92	83.08	36.25	63.75	2.98	97.02
10	16.97	83.03	36.31	63.69	3.01	96.99
Europe						
1	39.65	60.35	71.87	28.13	3.20	96.80
2	35.82	64.18	59.85	40.15	10.33	89.67
3	38.70	61.30	60.31	39.69	15.87	84.13
4	38.22	61.78	60.12	39.88	16.27	83.73
5	38.11	61.89	60.30	39.70	16.57	83.43
10	38.72	61.28	60.49	39.51	16.65	83.35

shown in Table 3 provide clear evidence that the period 1983–1999 is fundamentally different from the period before. While real activity shocks explain a large proportion of the variability of real stock prices during the period 1960–1982, this proportion becomes very small in the SVAR estimated for the period 1983–1999. In the US, real activity shocks explain about two thirds of the variability in real stock prices at longer horizons over the period 1960–1982 and this proportion drops to one fifth over the period 1983–1999. In Europe, real activity shocks explain some 60 percent of the variability over the period 1960–1982 and some 17 percent over the period 1983–1999. In Japan the explained proportion drops from 36 percent to 3 percent for the different periods. The considerably lower proportion of real stock price variability that can be explained by real activity shocks in Japan can be partially attributed to the fact that only one lag (as compared to three in the US and Europe) is included in the estimated VAR. However, the proportion still remains lower than in the US and Europe even if three lags are included in the SVAR (results not reported here).

The results of the stock price forecast error variance decompositions shown in Table 3 support the finding of a breakdown in the relation between stock prices and real activity in the US in the early 1980s reported in Binswanger (2000). Furthermore, the variance decompositions of our SVAR model shown in Table 3 also indicate a similar breakdown in Europe and Japan.

Finally, we use the estimated SVAR for the full sample (1960–1999) for a historical decomposition of stock prices over the period 1983–1999. As the restriction imposed on the estimated VAR allows the recovery of the entire $\{u_{1t}\}$ and $\{u_{2t}\}$ sequences of structural shocks we can also simulate what stock prices would have looked like if they had only been influenced by real activity shocks. For this purpose we set the other shocks at zero for the period 1983–1999 and allow the real activity shock to take its historical values. These historical values of the real activity shock are then used to calculate the component of stock prices that can be explained by real activity shocks for each quarter over the period 1983–1999 by employing the SVAR model estimated for the full sample. The estimated growth rates of this component are then converted to log levels by integrating the series forward as pictured in Fig. 3.

However, we have to be aware of the fact that these simulations require some additional (implicit) assumptions usually not mentioned in the literature. First, to implement the forward integration in our simulation we face a starting-value problem. Whenever we start the simulation, we implicitly assume that stock prices adequately reflect real activity at the beginning of the simulation period. Therefore, the absolute value of the simulated stock prices should not be overemphasized. The important issue is the degree to which the stock market movements during the simulation period can be explained by real activity shocks and not the absolute value of the simulated components. Furthermore, the chosen simulation method also implies that a simulation of the "real activity component" of stock prices over the whole sample period (1960–1999) will cause this component to be equal to the value of the actual series of the stock price in the final period.

The historical decompositions shown in Fig. 3, clearly suggest that most of the stock price movements over the period 1983–1999 cannot be attributed to real activity shocks. Not surprisingly, in Japan we observe the largest difference between the simulated stock prices and the actual series which is due to the extreme movement of Japanese stock prices over the 1980s and 1990s, which are commonly referred to as a bubble that burst in 1990. Furthermore, most of the stock price movements in Europe and the US during the period 1983–1999 cannot be attributed to real activity shocks. The sharp rise in European stock prices up to 1987 in particular does not seem to be related to any changes in real activity. Real activity can, however, partially explain the increasing stock prices in Europe as well as in the US in the early 1990s. But during the second half of the 1990s stock price movements again seem to be unrelated to changes in real activity. The historical decomposition of stock prices therefore suggests a strong influence of non-fundamental shocks on stock prices in Europe, Japan and the US over the period 1983–1999. If this were not the case there should always be a relation between stock prices and real activity even if other fundamental shocks (i.e. interest rate shocks, risk premium variations, variations in foreign real activity) temporarily obscure this relation.

Let us briefly compare our historical decomposition to the simulations presented in Lee (1998) and Chung and Lee (1998), which are based on SVAR models that include dividends and earnings rather than real activity. The simulation of Japanese stock prices over the full sample period (1975–1995) in Chung and Lee (1998), who also use quarterly observations, reveals that the increase in stock prices during the second half of the 1980s cannot be explained by earnings and dividends. This finding is largely in accordance with our results as changes in earnings and dividends should be linked to changes in real activity. However,

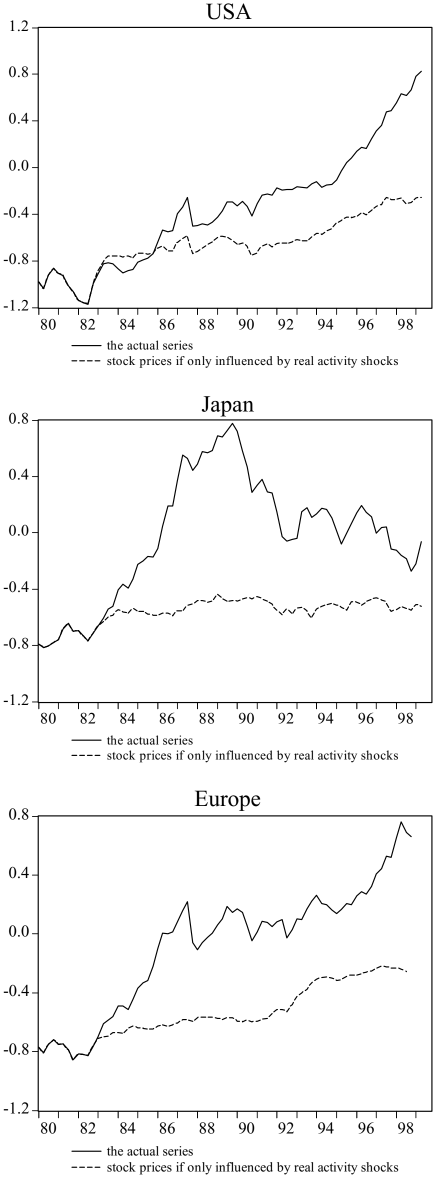

Fig. 3. Historical decomposition of stock prices (log levels normalized for 1960).

the deviation of stock prices from the simulated fundamental component is smaller than in the simulation presented in this paper and stock prices are again in line with fundamentals during the first half of the 1990s according to the simulation presented in Chung and Lee (1998).

Lee (1998) includes three fundamental variables (earnings, dividends, and time-varying discount factors) in his SVAR model for the US that uses annual observations from 1874 to 1995. The simulation over the full sample period shows that a large proportion of the rise in stock prices since the early 1980s cannot be attributed to changes in earnings or dividends. This finding is again in accordance with our results. However, the simulation also shows that part of the increase in stock prices which remains unexplained by earnings and dividends can be attributed to time-varying discount factors.[6]

5. Conclusion

In this paper we presented a bivariate SVAR model which includes growth rates of industrial production and stock prices. Imposing a long-run restriction à la Blanchard and Quah (1989) on a bivariate VAR that excludes long-run influences of the stock market on real activity allows the identification of two categories of structural shocks, which we called real activity shocks and other shocks. The first category of shocks include an important part of fundamental shocks while the second category mainly consists of nonfundamental shocks, though it also includes fundamental shocks not intrinsically related to domestic real activity. This identification scheme is motivated by our interest in the question as to how the relation between the stock market and real activity, which thus far has mainly been investigated by using single equation models, translates into a SVAR model, and whether this relation appears different over the recent stock market boom since the early 1980s if compared to the period from the 1960s to the early 1980s.

The results of the stock price forecast error variance decompositions from the SVAR model support the finding of a breakdown in the relation between stock prices and real activity in the US in the early 1980s reported in Binswanger (2000). Furthermore, the variance decompositions of our SVAR model also indicate a similar breakdown in Europe and Japan. While real activity shocks explain a large proportion of the variability of real stock prices during the period 1960–1982, this proportion becomes very small for the period 1983–1999. Furthermore, historical decompositions of stock prices over the period 1983–1999 show that most of the stock price movements during the period 1983–1999 cannot be attributed to real activity shocks. The pattern observed in the simulation can be compared to the results presented in Chung and Lee (1998) and Lee (1995a, 1998), where it was also the case that a large proportion of stock price movements over the 1980s and 1990s cannot be explained by earnings or dividends and must be attributed to nonfundamental factors or a time-varying discount rate.

A further hypothesis that may explain our findings would be the increasing globalization over the time period investigated in this paper. Globalization causes expectations of future cash flows to be less related to the development of domestic real activity and to be increasingly related to the development of the world market, where many of the transnational

[6] However, this finding may also be the result of the specific restrictions that Lee (1998, p. 10) imposes on the estimated VAR. According to these restrictions, nonfundamental shocks have no influence on time-varying discount factors and nonfundamental shocks only possess marginal explanatory power in the presence of the other structural shocks. Therefore, the model potentially underestimates the effect of nonfundamental shocks.

companies that exert a large influence on the domestic stock indices sell most of their product. We leave it to future research to further investigate whether the increasing globalization may also help to explain the observed breakdown in the relation between stock prices and real activity.

Overall, the results presented in this paper lend further credibility to the bubble hypothesis concerning the US, as well as Japan and Europe, because it is difficult to reconcile the marginal influence of real activity shocks on stock prices over the 1980s and 1990s with explanations that are based solely on changes in fundamentals. No convincing "fundamental story" has been told so far that would explain why stock price movements have simultaneously become much more independent of changes in domestic real activity in the three most important economic areas since the early 1980s.

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