Breakdowns and revivals: the long-run relationship between the stock market and real economic activity in the G-7 countries

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Abstract

Using monthly observations of industrial production and stock market indices from January 1961 to May 2012, we analyse the long-run relationship between the stock markets and real economic activity in the G-7 countries. In particular, this analysis uses the Toda and Yamamoto (1995) approach with the leveraged bootstrap methodology that was proposed by Hacker and Hatemi-J (2006). Our results indicate that although the expected long-run relationship holds for most of the G-7 countries, a break in this relationship occurred in the 1980s, followed by a subsequent revival after 2001.

Keywords: real economic activity, stock markets, G-7, long-run relationship

JEL Classification: E44, G15, O50

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Introduction

One of the fundamental premises of stock markets is that they are expected to predict the real economic activity. Under the present value model, the stock prices of a firm should be equal to the expected present value of the firm's future payouts, which relate to the state of the real economy. Investors' expectations theoretically account for all of the available information regarding future payouts and discount rates; in response to new information, stock prices react immediately, whereas changes in the real economy occur naturally but in a somewhat delayed fashion (e.g., Humpe and Macmillan, 2009).

We note that this model suggests the existence of a long-run relationship between stock prices and real output; this model can also be used to explain short-run changes, i.e., relationships between stock returns and output growth. Following the publication of a seminal study by Fama (1990), this short-run relationship has been the focal point of the empirical research that has addressed this topic. Fama (1990), Schwert (1990), and Darrat and Dickens (1999) strengthened the belief in the existence of a "return – growth" relationship in the US. In assessments of all of the G-7 countries, Choi *et al*. (1999) found evidence of this relationship (but not Italy), Hassapis and Kalyvitis (2002) found some evidence of this relationship for five of the G-7 countries (but not Italy and France), and Panopoulou *et al*. (2010) found evidence of this relationship in six out of the G-7 countries (but not France). However, the studies of Binswanger (2000; 2004), Canova and De Nicoló (2000) and Dufour and Tessier (2006) suggested that the existence of the "return – growth" relationships is far from certain for the G-7 countries. More specifically, Binswanger (2000; 2004) argued that since the mid-1980s, this relationship has weakened (as a consequence of stock market bubbles) in most of these nations.

Although cointegration only allows for the characterisation of rather restrictive types of relationships, the long-run relationship has nonetheless typically been assessed using cointegration-based techniques; for instance, this type of approach was adopted by Darrat and Dickens (1999), Choi *et al*. (1997), Nasseh and Straus (2000), and Humpe and Macmillan (2009), among others.

In this article, we assess the long-run "price – output" relationship in the G-7 countries, utilising the lag-augmented vector auto-regression model of Toda and Yamamoto (1995). The remainder of the article is organised as follows. Section 2 presents a description of the data and the samples, and Section 3 describes the methodology of this study. Section 4 provides our empirical findings, and Section 5 concludes the paper.

I. Data and sample description

Seasonally adjusted monthly industrial production and consumer price indices were obtained from the OECD's Main Economic Indicators, and monthly stock market indices were obtained from the OECD's Monthly Financial Statistics. The real stock market indices were adjusted using the CPI values. In the subsequent analysis, we used logs of the real stock market and industrial production indices (which are henceforth referred to as SMI and IP, respectively).

Our first sample consists of all of the available data, i.e., from January 1961 to May 2012 (T = 617). Our second sample ranges from January 1983 to December 2001 (T = 228). This starting date was chosen because it roughly corresponds to both a rapid increase in prices in the G-7 stock markets and the occurrence of structural changes in certain economic variables, as suggested by previous studies, e.g., Faroque *et al.* (2012). The third sample that was assessed by Binswanger (2004) also ranges from January 1983 to December 1999. Domian and Louton (1997) argue that decreases in output may occur more quickly than increases (which could be caused by various phenomena, including enhancements to production capacities and the implementation of new technologies); therefore, stock market prices tend to predict economic slowdowns more accurately than these prices predict economic upticks. Taking this notion into account, the end of the second sample of this study was set to December 2001. This sample therefore includes not only the start of the dot-com bubble (which is typically considered to have peaked in the first quarter of 2000) but also the burst of this bubble and the subsequent terrorist attacks of September 11th. Our third sample starts in January 1989 and ends in December 2001 (T = 156); thus, this sample excludes the 1987 stock market crash that was present in our second sample. Finally, the fourth sample of this study begins in January 2002 and ends in May 2012 (T = 125).

II. Methodology

The Granger causality tests were performed within the Toda and Yamamoto (1995) framework. First, we used the ADF-GLS test with either trends and constants (for levels) or constants alone (for differenced series) to determine the order of integration of each examined series. In the auxiliary regression, we added augmented terms until the null of no autocorrelation of residuals could not be rejected (at α = 0.05, using the Ljung-Box test with 1,…,12 lags). The finite sample critical values were calculated from the response surfaces of

Cheung and Lai (1995). All of the series have been demonstrated to be integrated of order one.

Subsequently, for each country and sample, we determined the optimal lag *p* of the unrestricted bivariate VAR using the Hatemi-J (2003) selection criteria. However, if the resulting residuals were autocorrelated, we increased the order of the system by one until the null of no autocorrelation of residuals could not be rejected (at α = 0.05, using the multivariate test by Mahdi and McLeod (2011) with 5, 10, 15, and 20 lags). In the next step, we quantified the VAR(*p*+*d*) as follows:

$$\begin{aligned} IP_t &= \alpha_{1,0} + \sum_{i=1}^{p}\alpha_{1,i}IP_{t-i} + \sum_{j=p+1}^{p+d}\alpha_{1,j}IP_{t-j} + \sum_{i=1}^{p}\beta_{1,i}SMI_{t-i} + \sum_{j=p+1}^{p+d}\beta_{1,j}SMI_{t-j} + e_{1,t} \\ SMI_t &= \alpha_{2,0} + \sum_{i=1}^{p}\alpha_{2,i}SMI_{t-i} + \sum_{j=p+1}^{p+d}\alpha_{2,j}SMI_{t-j} + \sum_{i=1}^{p}\beta_{2,i}IP_{t-i} + \sum_{j=p+1}^{p+d}\beta_{2,j}IP_{t-j} + e_{2,t} \end{aligned} \tag{1}$$

where *d* is the maximal order of integration of the series used in the system. Toda and Yamamoto (1995) demonstrated that the use of the Wald-χ^2 test to determine the parameter restrictions in (1) can be used to establish that SMI does not Granger cause (≠>) IP, i.e., $\forall\, i$: $\beta_{1i} = 0$.

However, instead of the asymptotic χ^2-distribution, we utilised the leveraged bootstrap methodology (with 10^4 bootstrap samples) that was proposed by Hacker and Hatemi-J (2006). In general, this methodology should produce better finite sample properties, even for non-normal and conditionally heteroskedastic residuals (ARCH), which were present in most of our models (ARCH effects were tested by the Mahdi and McLeod, 2011 test and normality was assessed with the kurtosis and skewness tests of Kankainen *et al.*, 2007).[1]

III. Empirical evidence

Table 1 contains the estimation results. We were able to confirm the existence of the “price – output” relationship for most of the G-7 countries from 1961 to 2012; however, we were unable to confirm this relationship in France and the United Kingdom (UK). The evidence for the “return – growth” relationship in France was already weak, but the lack of support for this relationship in the UK was somewhat surprising (see, e.g., Choi *et al.*, 1999; Binswanger, 2004; and Panopoulou *et al.*, 2010). In this particular case, it appears that from a long-term perspective, investors are good at predicting changes but not trends. To assess the

[1] The results for the analyses of the unit-roots, autocorrelations, non-normality of the residuals, and conditional heteroskedasticity of the VAR(*p*) are available upon request.

overall effects of SMI on IP, Table 1 also presents the sum of the *p*-lagged SMI values from the VAR(*p*), i.e., without augmentation. This sum was positive for all but one of the examined cases.

Table 1. Granger non-causality tests: monthly frequencies

Null hypothesis	01/1961-05/2012 p	$Wald\text{-}\chi^2$	Σ	01/1983-12/2001 p	$Wald\text{-}\chi^2$	Σ	01/1989-12/2001 p	$Wald\text{-}\chi^2$	Σ	01/2002-05/2012 p	$Wald\text{-}\chi^2$	Σ
$SMI_{CAN} \not\Rightarrow IP_{CAN}$	4	23.68***	0.001	2	0.90	0.017	2	2.48	0.009	4	25.24***	-0.002
$SMI_{FRA} \not\Rightarrow IP_{FRA}$	4	4.42	0.001	2	0.41	0.009	2	2.69	0.014	4	18.90***	0.008
$SMI_{DEU} \not\Rightarrow IP_{DEU}$	4	22.80***	0.005	2	3.61	0.013	2	1.29	0.020	3	21.23***	0.031
$SMI_{ITA} \not\Rightarrow IP_{ITA}$	4	15.02***	0.000	2	0.08	0.004	2	5.00*	0.004	3	24.71***	0.033
$SMI_{JAP} \not\Rightarrow IP_{JAP}$	4	20.55***	0.002	4	3.16	0.009	2	1.87	0.012	2	1.69	0.037
$SMI_{GBR} \not\Rightarrow IP_{GBR}$	4	6.88	0.002	2	0.04	0.012	2	3.73	0.015	2	5.70*	0.008
$SMI_{USA} \not\Rightarrow IP_{USA}$	4	50.44***	0.002	4	18.22***	0.012	3	14.21***	0.015	3	28.28***	0.017

*Notes: ***, ** and * denote significance at α < 0.01, α < 0.05 and α < 0.1, respectively. Note that d = 1 for all of the examined systems. Σ denotes the sum of the p-lagged SMI coefficients from the VAR(p). The country codes correspond to the ISO 3166-1 alpha-3 codes.*

Table 1 indicates that distinct per-country estimation results are obtained for different periods. In particular, for Canada, Germany, Italy and Japan, the break in the "price – output" relationship occurred during the 1980s and 1990s. Excluding the 1987 event from this subsample (as in our third sample) does little to alter the results. These results confirm the weakening of the "return – growth" relationship that was found by Binswanger (2004). Interestingly, since 2002 (our fourth sample), we have witnessed the revival of the "price – output" relationship for all of the G-7 countries except Japan (including France and the UK). Between 1990 and 2012, Japan's real stock prices declined, but the output of its economy (its industrial production index) remained nearly unchanged. In this situation, the predictive power of the stock market with respect to the output of the economy appears to be non-existent.

IV. Conclusions

In this article, we were able to confirm a breakdown of the long-run "price – output" relationship for four of the G-7 countries during the 1983 – 2001 period: Canada, Germany, Italy and Japan. However, during the most recent decade (2002 – 2012), the predictive power of the stock market regarding the output of the economy has been restored in all of the G-7 countries except for Japan. These results remain robust even if we increase the maximum order of integration to $d = 2$ or increase the optimal p by one.[2]

Fama (1990) argues that using quarterly or annual data is preferable for these types of analyses because information about the state of the future production is spread across many previous periods. We have therefore expanded our analysis through the use of non-overlapping, end-of-quarter observations. Unfortunately, this approach greatly decreases the number of observations for the third and fourth samples; therefore, estimations were performed only for the first two samples. Table 2 displays these results. We note that for the duration of the sample period, the existence of the "price – output" relationship was verified for all of the G-7 countries; however, this relationship eventually deteriorated for all of the examined countries except for the US. This confirms that the long-run "price – output" relationship was absent during the 1980s and 1990s.

Table 2. Granger non-causality tests: quarterly frequency

Countries	01/1961-05/2012			01/1983-12/2001		
	p	*Wald*-χ^2	Σ	*p*	*Wald*-χ^2	Σ
$SMI_{CAN} \nRightarrow IP_{CAN}$	3	30.28^{***}	-0.003	2	4.35	0.005
$SMI_{FRA} \nRightarrow IP_{FRA}$	2	8.36^{**}	0.002	1	1.11	0.028
$SMI_{DEU} \nRightarrow IP_{DEU}$	2	20.14^{***}	0.011	1	1.43	0.037
$SMI_{ITA} \nRightarrow IP_{ITA}$	2	13.21^{***}	-0.001	2	5.72^{*}	0.007
$SMI_{JAP} \nRightarrow IP_{JAP}$	2	12.86^{***}	0.002	1	0.72	0.022
$SMI_{GBR} \nRightarrow IP_{GBR}$	3	20.77^{***}	0.006	1	2.00	0.028
$SMI_{USA} \nRightarrow IP_{USA}$	2	55.15^{***}	0.004	2	11.50^{***}	0.031

Notes: ***, ** *and* * *denote significance at* $\alpha < 0.01$, $\alpha < 0.05$ *and* $\alpha < 0.1$, *respectively. Note that d = 1 for all systems.* Σ *denotes the sum of the p-lagged SMI values from the VAR(p).*

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[2] The results of both of these analyses are available upon request.

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