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International evidence on the stock market and aggregate economic activity

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Abstract

Using the Johansen cointegration technique, we find empirical evidence of long run comovements between five national stock market indexes and measures of aggregate real activity including the real oil price, real consumption, real money, and real output. Real returns on these indexes are typically related to transitory deviations from the long run relationship and to changes in the macroeconomic variables. Further, the constraints implied by the cointegration results yield some incremental information on stock return variation that is not already contained in dividend yields, interest rate spreads, and future GNP growth rates. © 1998 Elsevier Science B.V. All rights reserved.

1. Introduction

It is often observed that stock prices tend to fluctuate with economic news, and this observation is supported by empirical evidence showing that macroeconomic variables have explanatory power for stock returns. Fama (1981, 1990), Chen et al. (1986), Barro (1990), Schwert (1990) and Ferson and Harvey (1991) have found that U.S. stock returns and its aggregate real activity are correlated. Asprem (1989), Beckers et al. (1992), Ferson and Harvey (1993), Cheung et al. (1997a,b) have reached a similar conclusion using other international market data.

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These studies emphasize the short-run relationship among stock returns, macroeconomic variables, and financial variables. Little work, however, is focused on the long run comovement between stock prices and the underlying economic forces that drive these asset prices through time. Nor any transitory deviations from this long run relationship are, in general, incorporated in these studies of stock return variability [1]. In this paper we investigate the interactions between national stock market prices and aggregate economic variables by examining their empirical long run relationship. Such evidence not only provides an insight about their long run behavior, but also sheds light on the nature of their short run variation. This study adopts the cointegration concept, originally developed by Engle and Granger (1987), to investigate the long run comovement of the stock market level and aggregate economic variables. If these variables are cointegrated, they tend to move together in the long run, while experiencing short run transitory deviations from this long run relationship. From a cointegrated system, we can derive an error correction model (ECM) that allows us to study both the short run dynamics and the effect of long run restrictions on stock return variation. The Johansen (1991) procedure, which is more efficient than the Engle–Granger two-step method, is used to test for the existence of cointegration.

We further integrate our work with the approach of Fama (1990). In his study, Fama determines the extent to which variations in expected future cash flows and changes in discount rates can explain the stock return variation in the U.S. market. He estimates the relationship between stock returns and the variables which are proxies for changing expected returns, shocks to expected returns, and changing investors' expectations about future cash flows. We complement his study by examining whether adding information derived from our cointegration analysis to these measures of return variation can help improve the model for stock returns.

Using quarterly data of Canada, Germany, Italy, Japan, and the U.S., we find evidence of long run comovements between the national stock market index levels and country-specific aggregate economic real variables such as the real oil price, real output, real money supply, and real consumption. We also find that the constraint implied by the cointegration result provides some incremental information that is not already captured by variables that are proxies for the three sources of equity return variation, as suggested by Fama (1990).

The rest of the paper is organized as follows. In the next section, we document the data source and provide the rationale for selecting the aggregate economic variables. Section 3 briefly describes the statistical procedures. We report the cointegration results in Section 4. In Section 5, we examine the performance of ECM in the presence of other measures of stock return variation that are previously found to have explanatory power. We summarize in Section 6.

[1] A notable exception is Campbell and Shiller (1988).

2. Data

This study obtains quarterly stock index and macroeconomic data of Canada, Germany, Italy, Japan, and the U.S. from both the International Financial Statistics data tape and CITIBASE. The macro variables which are proxies for measures of aggregate economic activity include the real oil price, real gross national product (GNP), real money supply, and real consumption [2], and the motivation for selecting these variables will be discussed below. All data series are in natural logarithms. We use the countries' respective consumer price indexes to convert their nominal variables to real terms. A description of the data is given in Table 1.

Since asset pricing theories do not specify what the underlying economic forces are that drive the asset prices, we draw on the existing literature to select macro variables which are proxies for these forces. The GNP measures the economy's overall economic activity that affects stock prices through its influence on future cash flows. The money supply is related to the stock market in several ways. For instance, the portfolio balance model suggests that an increase in money supply leads to a portfolio shift from non-interest bearing money to financial assets including equities. Money supply fluctuations can also affect the stock market through their effects on inflation uncertainty. Mandelker and Tandon (1985) show that future growth rates in real GNP and money growth rates have a positive impact on real stock returns in six major industrialized countries. Asprem (1989) finds that expectations about future real activity and measures of money are positively related to stock prices in ten European countries.

The linkage between consumption and the stock market activity is theoretically established by the consumption-based capital asset pricing model (CAPM). The model assumes that the state variables determining asset prices also covary with marginal utility and, hence, are inversely related to consumption (see Lucas, 1978; Breeden, 1978; Hansen and Singleton, 1982, 1983). Wheatley (1988) finds that consumption risk is significantly priced in the sample of 18 countries and concludes that the consumption based CAPM holds.

Oil prices capture possible effects of external shocks on output and price developments in these industrialized countries. Chen et al. (1986) suggest oil prices as a measure of economic risk in the U.S. stock market, and Hamao (1988) and Brown and Otsuki (1990) find that oil price changes play an important role in

[2] Due to data availability, the Italian gross national product data are constructed from the annual series. The oil price series of Canada, Germany, Japan, and Italy were constructed based on both the U.S. oil price data and the country's respective exchange rates. Given the violation of purchasing power parity in the short run, using actual exchange rates may reflect the cost of oil in local currencies better. It is recognized that tax rates imposed on the price of oil differ across countries. However, using the before-tax price of oil is in accord with the extant empirical literature on the effects of macroeconomic factors, including oil prices, on stock market returns. Furthermore, tax rate variations on the price of oil would not affect our current study that mainly looks at the comovement of a country's stock market index and its aggregate economic variables.

Table 1
Glossary and definition of data

Country	Description of the stock indexes (sample period)
Canada	closing quotations from the Toronto Stock Exchange for a composite of 300 shares from Bank of Canada (1957:1–1992:2)
Germany	average of daily quotations covering approximately 95% of common shares of industrial companies with headquarters in Germany provided by the Federal Statistical Office (1960:1–1992:1)
Italy	average of daily quotations of common shares of 40 major companies on the Milan Exchange provided by the Bank of Italy (1970:1–1991:1)
Japan	average of daily quotations for all shares listed on the Tokyo Stock Exchange and are provided by the Bank of Japan (1957:1–1992:2)
U.S.	a Laspeyres-type index of Standard and Poor's Corp. for 400 industrials on the New York Stock Exchange, based on daily closing quotations, provided by the U.S. Department of Commerce (1957:1–1992:2)
Symbol	**Measures of aggregate economic activity**
O	crude petroleum price index
M	money supply as defined by M1
GNP	gross national product
C	total personal consumption

The sample periods of the national macro variables are the same as those of the national stock indexes. Due to data availability, the GNP for Italy is constructed from the annual series.

pricing Japanese equities. Ferson and Harvey (1993) find that changes in U.S. crude oil prices contribute a significant source of global economic risk in 18 national equity markets.

It is well documented that variables such as dividend yields, the default spread, and the term structure of interest rates have predictive power for stock returns. These variables typically have the same order of integration as the stock return and, hence, have a different order from that of the stock price level. Because of differences in the order of integration, dividend yields and interest rate spreads cannot cointegrate with stock prices, and the former variables are therefore excluded from our cointegration analysis of stock prices. However, we will include these variables when investigating the determinants of stock return variability in Section 5.

3. Cointegration analysis

The Johansen (1991) maximum likelihood (ML) procedure, which has been shown to have good large- and finite-sample properties [3], is used to conduct the

[3] Phillips (1991) examines its asymptotic properties, while Cheung and Lai (1993) and Gonzalo (1994), for example, examine its finite-sample properties.

cointegration analysis. To implement this procedure, we first estimate the least squares residuals from

$$\Delta Y_t = \mu_1 + \sum_{i=1}^{k-1} \Gamma_i \Delta Y_{t-i} + \varepsilon_{1t}, \tag{1}$$

and

$$Y_{t-k} = \mu_2 + \sum_{i=1}^{k-1} \Gamma_i \Delta Y_{t-i} + \varepsilon_{2t}, \tag{2}$$

where $Y_t \equiv (s_t\, X_t)'$, s_t is the real stock index, X_t is a vector of macro variables, μ_1 and μ_2 are constant vectors, and Δ is the difference operator. Next, we compute the eigenvalues, $\lambda_1 \geq \dots \geq \lambda_n$, of $\Omega_{21}\Omega_{11}^{-1}\Omega_{12}$ with respect to Ω_{22} and the associated eigenvectors, $v_1, \dots, v_n$, where the moment matrices $\Omega_{ij} = T^{-1}\sum_t \hat{\varepsilon}_{it}\hat{\varepsilon}'_{jt}$ for i, j = 1, 2. The trace statistic,

$$t_r = -T \sum_{j=r+1}^{n} \ln(1-\lambda_j),\ 0 \leq r \leq n, \tag{3}$$

tests the hypothesis that there are at most r cointegration vectors. In testing the hypothesis of r against the alternative hypothesis of $r+1$ cointegration vectors, we use the maximum eigenvalue statistic,

$$\lambda_{r|r+1} = -T\ln(1-\lambda_{r+1}). \tag{4}$$

The eigenvectors $v_1, \dots, v_r$ are sample estimates of the cointegration vectors.

Tests based on asymptotic critical values tabulated in Johansen and Juselius (1990), however, tend to reject the no-cointegration null too often in finite samples. Since our data series are not relatively long, using proper finite-sample critical values is important. To minimize this potential bias, we employ finite-sample critical values that have been adjusted for sample size effects, the system's dimension, and the lag order (Cheung and Lai, 1993).

4. Results from cointegration analyses

We apply the augmented Dickey–Fuller statistic and Kwiatkowski et al. (1992) (KPSS) trend stationarity statistic to test for unit roots in the data. While the augmented Dickey–Fuller statistic tests the unit root null hypothesis, the KPSS statistic tests the trend stationarity null hypothesis against the unit root alternative. Given different specifications of the null and alternative, the KPSS test can provide useful additional information on the unit root persistence of the data. In performing these tests, we include both a time trend and a constant in the

Table 2
Results of the Johansen cointegration test on each country's stock index and aggregate economic variables

Country	Maximum eigenvalue statistic			Trace statistic		
H_0:	$r = 2$	$r = 1$	$r = 0$	$r \leq 2$	$r \leq 1$	$r = 0$
Canada	16.0615	26.4331	32.4805	20.5433	46.9764	79.4569 *
Germany	14.9766	32.3054 *	39.5721 *	23.6615	55.9668 *	95.5389 *
Italy	16.4985	41.6161 *	53.1859 *	22.9627	64.5788 *	117.7647 *
Japan	9.3694	19.5822	47.4491 *	17.5896	37.1718	84.6209 *
U.S.	12.1826	19.1005	48.8281 *	18.4950	37.5955	86.4236 *

The maximum eigenvalues and trace statistics for cointegration are reported. For each country, the system consists of real stock index, real oil price, real consumption, real money supply, and real gross national product. All quarterly data series are in natural logarithms. The significance of the above sample statistics are evaluated based on the critical values in Cheung and Lai (1993). The lag order k used in the Johansen test is determined by the Schwarz information criterion.

* Significance at the 5% level.

regression. Both the KPSS test and augmented Dickey–Fuller test (not reported) yield similar results and indicate that the variables under examination are integrated of order one [4].

Table 2 summarizes the Johansen test results. Cointegration between stock indexes and macro variables is evident in all five countries. For Canada, Japan, and the U.S., there exists a single cointegration relationship between the country's stock index and aggregate economic variables, whereas for Germany and Italy, there exist two cointegration relationships [5]. Table 3 presents these estimated cointegration vectors, which are normalized so that the coefficient of each stock index is unity.

In four of the five cointegrated systems, the coefficients of stock and oil prices have the same sign, implying that oil prices are negatively correlated with stock prices. The finding is consistent with the common observation that increases in oil prices generally would cause a rise in production costs and a subsequent fall in aggregate economic activity. This, in turn, would lower expected future cash flows. Stock prices and consumption, however, move in the same direction, lending support to the consumption-based CAPM. In contrast, effects of both real money supply and real GNP on the stock markets are ambiguous. As discussed in Section 2, money supply can affect the stock market in several ways, and the inconclusive result probably reflects investors' perceived differences in the monetary policy implemented across countries.

[4] Finite sample critical values provided by Cheung and Lai (1995) are used to evaluate the significance of the unit root test. See Cheung and Ng (1996) for the estimated test statistics and a more detailed discussion of the test results.

[5] Note that the Canadian result is based on the trace statistic only.

Table 3
The estimated cointegration relationships of each country's stock index and aggregate economic variables

	Real stock price	Real oil price	Real consumption	Real money	Real GNP
Canada	1.0000	−1.2847	−3.9301	−2.5312	6.1094
Germany	1.0000	−0.2951	10.4420	−2.1113	−7.8096
	1.0000	4.8622	−104.5737	10.7563	89.3691
Italy	1.0000	0.3373	−31.6692	−5.7264	38.1036
	1.0000	3.4839	−3.2102	31.4812	−15.7699
Japan	1.0000	3.6689	34.1300	3.9321	−28.5861
U.S.	1.0000	7.5356	−23.0109	30.9030	13.8012

The cointegrating vectors are estimated from the cointegrated systems reported in Table 2. These vectors are normalized so that the coefficient of each stock index is unity.

Given the cointegration results, it is interesting to examine the error correction terms given by the product of the cointegration vectors and the system variables. The resulting error correction terms can be interpreted as a measure of the deviation from the long run relationship. Table 4 reports some descriptive statistics for the estimated error correction terms, and Fig. 1 plots these estimated series.

Based on the coefficient of variation (a unit free measure of variability relative to the mean), we infer that the error correction terms for Germany and Italy tend to fluctuate quite substantially, whereas that for Japan exhibit the largest variability. Except for Japan, the autocorrelation coefficients suggest no evidence of unit root persistence in these series. Note that the error correction term of a cointegrated system should follow a stationary $I(0)$ process. Due to sampling uncertainty, however, it is possible that the Japanese error correction term constructed from the estimated cointegration vector retains some long term persistence of the individual component variables and, hence, exhibits some kind of $I(1)$ behavior.

Table 4
Descriptive statistics for the error correction terms

	Canada	Germany		Italy		Japan	U.S.
CoVa	0.042	−0.223	0.061	0.022	0.126	0.143	−0.034
Skewness	−0.317	0.235	0.078	−0.294	−1.038	−0.300	0.136
Kurtosis	−0.104	−0.053	−0.623	0.657	3.423	−1.297	−0.076
$\rho(1)$	0.749	0.624	0.669	0.591	0.519	0.982	0.289
$\rho(2)$	0.600	0.514	0.577	0.183	0.359	0.961	0.157
$\rho(3)$	0.476	0.399	0.575	−0.116	0.238	0.941	−0.039
$\rho(4)$	0.434	0.461	0.568	−0.112	0.453	0.923	0.19S
$\rho(5)$	0.287	0.183	0.420	−0.322	0.052	0.904	−0.162
$\rho(10)$	0.112	0.167	0.321	0.084	−0.231	0.812	−0.019

Some descriptive statistics for the error correction terms generated from the cointegration relationships in Table 2Table 3 are report. 'CoVa' gives the coefficient of variation (which is a unit free measure of variability), and $\rho(k)$ is the autocorrelation at lag k.

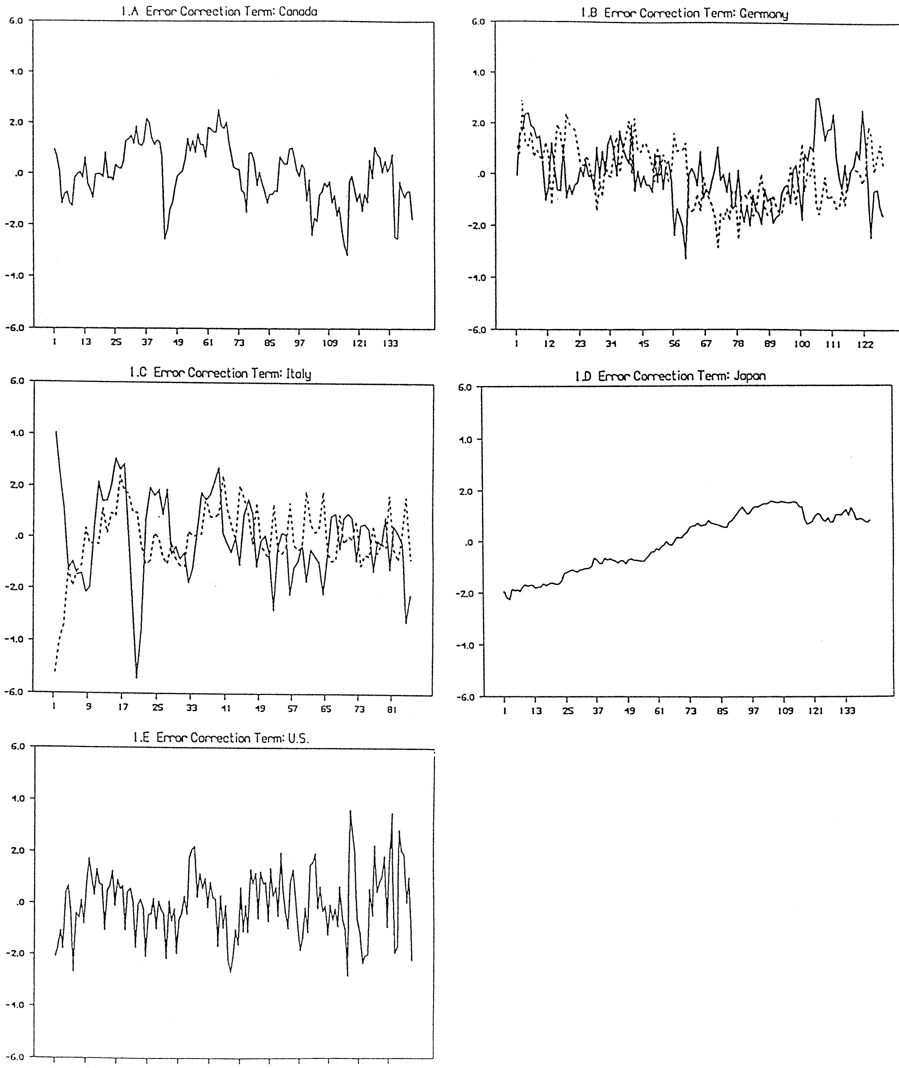

Fig. 1. The error correction terms.

To facilitate comparison, the estimated error correction terms are plotted on the same vertical scale, with each series normalized by its sample mean. There is no apparent pattern detected in these series. The most persistent series is the Japanese error correction term — an observation that is in accordance with results reported in Table 4. The volatility of the German and U.S. error correction terms increases slightly over time, while that of the Italian one decreases. There seems little

Table 5
Estimation results of the error correction model

	c	Δs_{-1}	ΔO_{-1}	ΔM_{-1}	$z1_{-1}$	$z2_{-1}$	Q_{10}	$\bar{R}^2$
Canada	0.3754	0.2734	−0.0929		−0.0139		9.834	0.107
	(3.12)	(3.16)	(2.24)		(3.13)		(0.46)	
	−0.0002	0.2540	−0.0851				13.57	0.066
	(0.04)	(2.81)	(2.11)				(0.19)	
Germany	0.1436	0.3314			−0.0118	−0.0113	6.248	0.118
	(1.40)	(3.20)			(1.90)	(2.08)	(0.79)	
	−0.0013	0.2932					8.013	0.084
	(0.19)	(2.64)					(0.63)	
Italy	1.0984	0.4125	−0.1816		−0.0170	0.0184	11.85	0.277
	(2.22)	(4.40)	(1.98)		(2.71)	(2.11)	(0.30)	
	−0.0069	0.3703	−0.2042				10.38	0.204
	(0.56)	(3.81)	(2.32)				(0.41)	
Japan	−0.0079	0.2873	−0.1100	0.3078	0.0012		4.024	0.138
	(0.17)	(3.10)	(1.84)	(3.10)	(0.21)		(0.91)	
	0.0018	0.2873	−0.1102	0.3048			3.987	0.144
	(0.28)	(3.11)	(1.84)	(3.21)			(0.95)	
U.S.	0.4458	0.2453		0.4985	−0.0115		13.68	0.138
	(2.02)	(2.73)		(2.54)	(2.00)		(0.18)	
	0.0025	0.2903		0.3196			13.55	0.101
	(0.48)	(3.43)		(1.94)			(0.19)	

Only significant coefficients in the error correction models are reported. For each country, the dependent variable is the real stock index return (ΔS), the independent variables are the lagged real stock index return (ΔS), the lagged growth rate in real oil price index (ΔO_{-1}) and real money supply (ΔM_{-1}), and the lagged error correction terms ($Z1_{-1}$ and $Z2_{-1}$). Absolute values of heteroskedastic-consistent t-statistics are reported in parentheses below the parameter estimates. Q_{10} is the Ljung–Box statistic computed from the first ten autocorrelation coefficients of the residuals with its p-value given in parentheses. $\bar{R}^2$ is the coefficient of determination adjusted for the number of independent variables considered in each regression.

comovement in the two German error correction terms as well as the two Italian error correction terms, with sample correlation coefficients of −0.0945 and −0.2159, respectively.

Overall, the cointegration analysis indicates the existence of long run comovements between the selected real economic variables and real stock market prices. However, measures of aggregate real activity exhibit varying effects on the national stock market indexes. This diverse phenomenon may be attributable to differences in industrial structures, stock exchange trading systems, regulations on stock markets, and the fiscal and monetary policies of the five countries examined.

Given the cointegration results, we estimate the following ECM:

$$\Delta s_t = c_1 + \sum_{i=1}^{q} \gamma_i z_{i,t-1} + \sum_{i=1}^{p} \alpha_i \Delta s_{t-i} + \sum_{i=1}^{k} \phi_i \Delta X_{t-i} + \omega_{1t} \tag{5}$$

c_1, γ_i, α_i, and ϕ_i, are coefficient matrices of appropriate dimensions. $z_{i,t-1}$ is the error correction term and w_{1t} is the regression error term. k, p, and q are the lag parameters. Table 5 reports the estimates of Eq. (5), detailing only the significant first-difference terms with absolute values of heteroskedastic-consistent t-statistics in parentheses. The Ljung–Box test statistic, Q_{10} (computed from the first ten autocorrelation coefficients of the residuals), provides no evidence of serial correlation in the residuals, thereby indicating that the model is adequately specified.

Table 5 shows that the error correction terms, lagged stock returns, changes in oil prices, and changes in money stock have significant effects on stock market movements, while changes in consumption and output have weak explanatory power. Except for Japan, z_i's are significant and negative, implying that these systems have a tendency to revert back to their empirical long run relationship. The real oil prices in Canada, Italy, and Japan are negatively correlated with their stock returns, indicating that any rise in oil prices would have a negative effect on stock returns in both the long and short runs. In contrast, the results demonstrate that any rise in money stock would have a positive effect on Japanese and U.S.

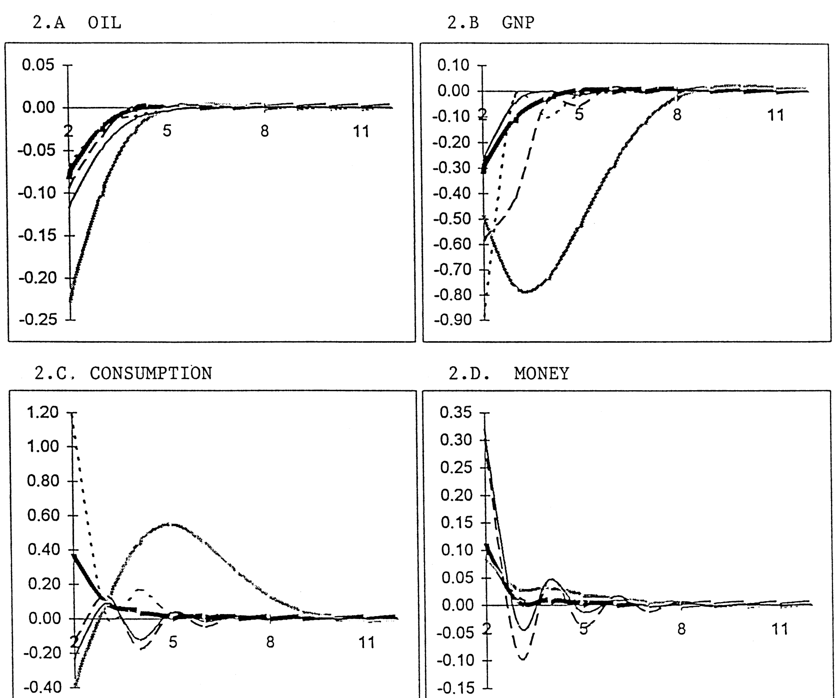

Fig. 2. Real stock price response to real macro economic shocks.

stock returns. This evidence ties in with previous results suggesting a positive relation between stock market returns and growth rates in money supply.

To gauge the importance of short run adjustments to deviations from the long run equilibrium, we re-estimate the ECM without z_t. Such a regression mimics the standard VAR approach, and Table 5 contains these estimates, as indicated in the second row of each country's entry. Although omitting z_t has only a marginal impact on the estimated coefficients, their standard errors, and Q_{10} statistics, it has a discernible effect on $\bar{R}^2$'s. In almost all countries, $\bar{R}^2$'s improve substantially when error correction terms are included in the regressions. Thus, when modeling stock return movements, it is important to incorporate the empirical long run restriction on the regression model as well.

Fig. 2 graphs the responses of changes in real stock prices to a unit shock in changes in real oil prices, real output, real consumption, and real money. These responses are generated from a vector ECM, which includes the five variables and z_t's. Notice that effects of unit shocks are usually small and decay fairly quickly. Shocks to real oil prices and real output tend to have an initial negative impact on real stock prices, whereas shocks to real money have a positive impact. However, no distinct pattern is found in real consumption shocks on stock prices [6].

5. Other sources of stock return variation

Thus far, our analysis only focused on effects of aggregate economic activity on the five national stock markets. A recent study by Fama (1990) postulates that time-varying expected returns, expected returns shocks, and shocks to expected future cash flows are the three sources of stock return variation [7]. In this section, we examine whether the reported ECMs provide any useful information that is not contained in these three measures of stock variation.

Consistent with the existing literature, we use U.S. variables as proxies for time-varying expected returns and shocks to expected returns [8]. The proxies for expected returns are: (1) the quarterly dividend yield (DY) for the U.S. value-weighted market index, (2) the default spread (DS) defined as the difference in annualized yields between the Moody's BAA corporate bond portfolio and AAA

[6] The impulse response functions are based on the following recursive structure: real stock price, real oil price, real GNP, real consumption, and real money. We also computed the impulse response functions for different orderings of these variables and found their plots are similar to those depicted in Fig. 2.

[7] The study of Fama (1990) was subsequently extended by Schwert (1990) using a century of historical U.S. data. Cheung et al. (1997a,b) have demonstrated that Fama's results are robust to some international stock markets.

[8] For example, Ferson and Harvey (1993), Cheung et al. (1994, 1997a,b) have shown that U.S. variables have better ability to explain stock return variations across different national stock markets than do the countries' own variables.

Table 6
Results from regressing quarterly real stock returns on proxies for time-varying expected returns, expected return shocks, and expected cash flow shocks

	Canada	Germany	Italy	Japan	U.S.
c	0.1770	−0.1367	0.0712	−0.1413	−0.1114
	(6.42)	(4.20)	(0.88)	(3.92)	(4.38)
DY_{-1}	14.557	10.450	−5.9406	10.656	9.6146
	(5.50)	(3.43)	(0.91)	(3.60)	(3.71)
TS_{-1}	0.0012	0.0053	−0.0104	0.0076	−0.0004
	(0.22)	(0.98)	(1.40)	(1.97)	(0.08)
SHD	−0.0835	−0.0751	−0.0114	−0.0836	−0.0644
	(3.05)	(3.49)	(0.16)	(3.90)	(2.55)
SHT	−0.0074	0.0006	0.0023	0.0048	−0.0067
	(1.19)	(0.11)	(0.16)	(0.92)	(1.12)
ΔY	0.9706	1.0146	0.7621	0.4480	0.9808
	(1.79)	(1.73)	(0.60)	(1.16)	(2.52)
ΔY_{+1}	1.2713	0.1087	−0.5137	0.9064	1.3500
	(2.74)	(0.32)	(0.38)	(2.63)	(2.89)
ΔY_{+2}	0.7498	1.4885	1.5309	0.9258	0.6878
	(1.38)	(4.36)	(1.26)	(2.40)	(1.44)
ΔY_{+3}	0.9484	0.6617	−2.0830	0.6374	0.3869
	(1.97)	(1.74)	(1.65)	(1.90)	(0.87)
Q_{10}	8.187	12.20	31.85	11.42	12.58
	(0.51)	(0.27)	(0.00)	(0.33)	(0.25)
$\bar{R}^2$	0.280	0.182	−0.032	0.210	0.258

The dependent variable is the real stock index return (ΔS), the independent variables are the dividend yield (DY), the term spread (TS), shocks to the default and term spreads (SHD and SHT), and the current and future growth rates of real GNP (ΔY_{+i}). Absolute values of heteroskedastic-consistent t-statistics are reported in parentheses below the parameter estimates Q_{10} is the Ljung–Box statistic computed from the first ten autocorrelation coefficients of the residuals with its p-value given in parentheses. $\bar{R}^2$ gives the coefficient of determination adjusted for the number of independent variables in each regression.

corporate bond portfolio, and (3) the term spread (TS) defined as the difference between the U.S. 10-year Treasury bond yield and the 3-month Treasury bill yield. All interest rates are obtained from CITIBASE. Following Fama (1990), residuals from first-order autoregressions fitted to DS and TS are used as proxies for shocks to expected returns, which are denoted by SHD and SHT, respectively. Since real GNP is viewed as real economic activity that influences future equity cash flows, the growth rates of country-specific real GNP are used as proxies for cash flow shocks to individual national stock markets.

For comparison purposes, we first estimate the Fama (1990) regression as given by

$$\Delta s_t = c + \alpha_1 \mathrm{DY}_{t-1} + \alpha_2 \mathrm{TS}_{t-1} + \alpha_3 \mathrm{SHD}_t + \alpha_4 \mathrm{SHT}_t + \alpha_5 \Delta Y_t + \alpha_6 \Delta Y_{t+1} + \alpha_7 \Delta Y_{t+2} + \alpha_8 \Delta Y_{t+3} + \upsilon_t, \tag{6}$$

where ΔY_{t+i} is the growth rate of real GNP from $t+i-1$ to $t+i$, υ_t is the disturbance term, and the remaining variables are defined above. DS and DY are not included in the same regression as these two variables are found to be highly

Table 7
Estimation results of the augmented error correction model (ECM), incorporating the proxies for time-varying expected returns, expected return shocks, and expected cash flow shocks

	Canada	Germany	Italy	Japan	U.S.
c	0.2127	0.1211	2.4147	−0.1980	0.1888
	(1.49)	(0.78)	(3.33)	(2.76)	(0.93)
Δs_{-1}	0.1202	0.2860	0.4102	0.1770	0.1495
	(1.53)	(3.18)	(4.30)	(1.54)	(1.76)
ΔO_{-1}	−0.0663		−0.1562	−0.1008	
	(1.45)		(1.21)	(1.79)	
ΔM_{-1}				0.2274	0.1244
				(2.49)	(0.67)
$z1_{-1}$	−0.0134	−0.0265	−0.0324	0.0117	−0.0077
	(2.75)	(3.16)	(3.65)	(1.33)	(1.45)
$z2_{-1}$		−0.0165	0.0102		
		(2.21)	(0.79)		
DY_{-1}	12.345	−0.3513	−2.0471	7.1362	9.6176
	(4.32)	(0.09)	(0.33)	(2.20)	(4.08)
TS_{-1}	−0.0045	0.0060	−0.0172	0.0020	−0.0022
	(0.82)	(1.36)	(2.41)	(0.43)	(0.38)
SHD	−0.0646	−0.0504	−0.0288	−0.0569	−0.0523
	(2.22)	(2.42)	(0.52)	(3.10)	(2.20)
SHT	−0.0106	0.0021	−0.0050	0.0001	−0.0072
	(1.66)	(0.38)	(0.43)	(0.02)	(1.28)
ΔY	0.9021	1.3360	2.0066	0.3205	0.8064
	(1.88)	(2.57)	(1.69)	(0.86)	(2.07)
ΔY_{+1}	1.2567	0.2481	−2.2178	0.8027	1.2718
	(2.65)	(0.58)	(1.78)	(2.29)	(2.92)
ΔY_{+2}	0.7452	1.5920	0.4004	0.7978	0.5927
	(1.25)	(5.68)	(0.37)	(1.73)	(1.29)
ΔY_{+3}	0.9733	0.7701	−1.7607	0.8102	0.3652
	(2.13)	(2.28)	(1.45)	(2.05)	(0.85)
Q_{10}	6.511	7.141	13.06	4.939	8.737
	(0.77)	(0.71)	(0.22)	(0.90)	(0.56)
$\bar{R}^2$	0.309	0.294	0.286	0.275	0.286

For each country, the dependent variable is the quarterly real stock index return (ΔS) and the independent variables are the lagged real stock index return (ΔS_{-1}), lagged growth rates in real oil price index (ΔO_{-1}) and real money supply (ΔM_{-1}), lagged error correction terms ($Z1_{-1}$ and $Z2_{-1}$), the dividend yield (DY), the term spread (TS), shocks to the default and term spreads (SHD and SHT), and current and future growth rates of real GNP (ΔY_{+i}). Absolute values of heteroskedastic-consistent t-statistics are reported in parentheses below the parameter estimates. Q_{10} is the Ljung–Box statistic computed from the first ten autocorrelation coefficients of the residuals with its p-value given in parentheses. $\bar{R}^2$ is the coefficient of determination adjusted for the number of independent variables in the regression.

correlated (Fama and French (1989)). We re-estimate Eq. (6) using DS in place of DY. Results of these two regressions are virtually similar and thus only estimates of Eq. (6) are reported in Table 6. As indicated in the table, the estimation results are qualitatively consistent with those of the existing studies. The selected proxies can explain about 20–30% of the stock return variation in Canada, Japan, and the U.S., but explain little of that in Italy.

Next, we add these measures of return variation to Eq. (5) and obtain

$$\begin{aligned}\Delta s_t = c &+ \varphi_1 \Delta s_{t-1} + \phi_1 \Delta O_{t-1} + \phi_2 \Delta M_{t-1} + \sum_{i=1}^{q} \gamma_i z_{i,t-1} + \alpha_1 \mathrm{DY}_{t-1} \\ &+ \alpha_2 \mathrm{TS}_{t-1} + \alpha_3 \mathrm{SHD}_t + \alpha_4 \mathrm{SHT}_t + \alpha_5 \Delta Y_t + \alpha_6 \Delta Y_{t+1} + \alpha_7 \Delta Y_{t+2} \\ &+ \alpha_8 \Delta Y_{t+3} + \varepsilon_t, \end{aligned} \tag{7}$$

The estimation results are reported in Table 7. We find that coefficient estimates on the explanatory variables from the original ECMs are typically smaller, and some of which even become statistically insignificant. However, for Canada, Germany, and Italy, the error correction terms remain highly significant. This finding may indicate that effects of error correction terms on stock returns cannot be fully absorbed by proxies for time-varying expected returns and shocks to both expected returns and future cash flows. The proxies TS and SHT, in general, are insignificant. On the other hand, future quarterly growth rates of real GNP appear to have significant influences on stock returns.

These estimation results reinforce our earlier findings of empirical long run comovements between real stock prices and aggregate economic variables. The ECM contains some useful incremental information on stock movements that is not already found in the commonly-used proxies for time-varying expected returns and shocks to both expected returns and expected future cash flows.

6. Summary

This study adopts the cointegration approach to examine the empirical relationship between five national stock market indexes and their aggregate economic variables. After adjusting for possible finite-sample biases, we find that real stock market indexes are typically cointegrated with measures of the countries' aggregate real activity such as the real oil price, real consumption, real money stock, and real output. Based on the ECM, we find that real returns on stock indexes are generally related to deviations from the empirical long run relationship and to changes in macro variables. Adding error correction terms to the model substantially improves the explanatory power for stock returns. We further find that the ECM provides incremental information about the stock return dynamics not found in other measures of return variation such as dividend yields, default and term spreads, and future GNP growth rates.

Our results indicate that a proper modeling of long run stock market dynamics helps in interpreting its return variation. Another possible venue to gain a better understanding of stock price movements is to incorporate information from other countries' equity markets. For instance, Kasa (1992) reports the presence of a common long term trend driving five national stock markets. Cheung and Lai (1994), on the other hand, find that common trends in the macro variables in three European Monetary System countries are cointegrated with the common trends in these countries' stock indexes. Nonetheless, these studies do not formally analyze if the long run relationship between macro and stock price variables can be exploited to better explain stock returns in a national market. Thus, an interesting extension of the current study is to examine whether foreign variables contain incremental information about the domestic stock return variation.

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