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# THE TRANSMISSION OF TORONTO AND PARIS STOCK EXCHANGES PRICE MOVEMENTS TO INTERNATIONAL STOCK MARKETS

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## 1. Introduction

The last decade was one of globalization of financial markets throughout the world, leading to a considerable level of joint price movements and a strong interdependence between the world stock markets. This phenomenon is bringing increased attention to the studies on the efficiency and the dynamic relationship between equity markets across the world.

Arshanapalli and Doukas (1993) use cointegration test procedures to provide evidence of a substantial increase in the level of the international stock market comovements for the post October 1987 period for the French, German, U.K. and U.S. markets. Employing unit root tests, Dwyer and Hafer (1988) investigate the connection among stock price comovements in Germany, Japan, the U.K. and the U.S. Using both daily stock prices from July 1987 to January 1988 and the monthly stock prices from 1957 through 1987, they show that there is no link among the levels, but among stock price changes across national markets. Kryzanowski and Liu (1993) use causality and cointegration tests based on monthly returns and prices from 1957 to 1990 to investigate the interrelationships among stock markets of eight major developed countries (Australia, Canada, France, Germany, Japan, U.K., U.S. and Switzerland). Using the *International Financial Statistics* database, their results suggest that some inter-

relationships exist among national stock markets in the short run, but that these markets exhibit no long-run comovement. Eun and Shim (1989) analyze the international transmission mechanism of stock market movements by estimating a nine-market vector autoregression. The nine markets considered are the eight developed markets studied by Kryzanowski and Liu (1993) plus Hong Kong. Using daily data from 1980 through 1985, their results indicate a substantial amount of multilateral interaction among the stock markets. However, their analysis of the international transmission of stock market movements is concentrated on the responses of the various markets to a shock in the U.S. market only. Their investigation reveals the leading role of the U.S. in the world stock market.

Chan et al. (1992) use unit root and cointegration tests to examine the relationships among the stock markets in major Asian markets (Hong Kong, South Korea, Singapore, Taiwan and Japan) and the U.S. Their findings, using data from the *Far Eastern Economic Review* suggests that the stock prices in major Asian markets and the United States are weak-form efficient individually and collectively in the long run. Fisher and Palasvirta (1990) provide a test for interdependence between the time series of stock markets indices of 23 countries to investigate the integration hypothesis of the world markets. Employing a cross-spectral analysis, they find a statistically significant level of interdependence between stock markets, and a leading role of the U.S. price index. While there are many other studies devoted to the investigation of comovements of developed stocks markets, there exists no in-depth analysis of the dynamic relationships between non-U.S. stock markets and the emerging stock markets. Therefore, the purpose of this paper is, first, to present evidence concerning the weak-form efficiency and the number of common stochastic trends in the seven industrialized countries stock markets (Canada, France, Germany, Italy, Japan, the U.K. and the U.S.) and nine emerging markets throughout the world (Argentina, Brazil, Chile, Greece, India, Korea, Mexico, Thailand, Zimbabwe). Secondly, we test for the interdependence relationships and dynamic interactions between the 16 markets considered to assess the relative importance of Toronto and Paris stock markets in generating fluctuations in stock returns in their own and other markets. Finally, the influence of the Toronto and Paris stock exchanges is measured by a dynamic impulse response analysis.

Using monthly data from December 1975 to December 1992 from the International Financial Statistics (for seven developed markets) and from the International Finance Corporation's *Emerging Markets Data Base* (for nine emerging markets), we employ unit root tests, cointegration tests, and vector autoregressive analysis to address the issues stated above. Unlike most of the previous studies, monthly data are used to minimize the biases associated with nonnormality, nonoverlapping trading sessions, nonsynchronous trading and to some extent the existence of the bid-ask spread. The rest of the paper is organized as follows. In the next section, the methodology is explained. The empirical results are presented and analyzed in section three, and the conclusion follows in section four.

## 2. Methodology

Four statistical tests are used: First, a cointegration test introduced by Engle and Granger (1987) and Johansen (1988, 1992) to examine the long-run relationships among the 16 stock markets. Testing for cointegration involves first to determine the order of integration, i.e., to test for the unit roots in each of the stock market price series. Therefore, two approaches, the Phillips-Perron (1988) and Dickey-Fuller (1981) procedures, are used to detect the presence of unit roots in the series. Second, by setting up a vector autoregression (VAR) system, the likelihood ratio test is used to investigate the Granger-causal direction among international stock markets. Third, once the sixteen-market VAR system is estimated, we follow Lee (1992) and Eun and Shim (1989) by using the Sim's (1980) approach to trace out the dynamic responses of each of the markets to shocks in the Toronto and Paris markets. Finally, the innovative accounting technique of the VAR allows us to assess the relative importance of each market in generating fluctuations in their own and other markets.

To test the random walk behavior of a time series  $X_t$ , we use the Augmented Dickey-Fuller (ADF) test statistic  $t_\alpha$  for  $\alpha$  in:

$$(1 - L)X_t = \alpha_0 + \eta t + \alpha X_{t-1} + \alpha_1(1 - L)X_{t-1} + \dots + \alpha_p(1 - L)X_{t-p} + e_t \quad (1)$$

where  $L$  is the lag operator,  $e_t$  is a Gaussian disturbance term,  $t$  is a time trend and  $p$  is the optimal lag in the autoregressive representation of  $X_t$ . Instead of choosing arbitrarily the lag length  $p$ , the value of  $p$  is set sufficiently large so that serial correlation in disturbances is avoided. The Phillips-Perron procedure is also used to avoid the problems arising from serial dependence through a non-parametric correction of the estimated error variance.

A market index series is said to be integrated of order one, denoted  $I(1)$ , if it has one unit root and is stationary only in the first difference. Two  $I(1)$  series are said to be cointegrated if there exists a linear combination of them which is  $I(0)$ . The two series are said to share a common stochastic trend, that is the series cannot diverge too far from each other over time. In this sense, there is a long run relation between the time series considered. Granger (1986) has demonstrated that market efficiency, in which the price of an asset incorporates all available and relevant information, has the important implication that prices from two efficient markets for different assets cannot be cointegrated. We use this result to test for the joint efficiency of the sixteen stock markets.

If the  $n$  time series show the same order of non-stationarity, following Kryzanowski and Liu (1993), and Turtle and Abeysekera (1993), test for their cointegration can then be conducted. More specifically, the null hypothesis that the series  $X_{it}$  is not cointegrated with the other series  $X_{jt}$ , ( $j \neq i$ ,  $j = 1, 2, \dots, n-1$  countries) is tested by examining the stationarity of the cointegrating residuals,

$u_{it}$ , from the following regression using the ADF test and the  $Z_\alpha$ -Phillips-Ouliaris (1990) test statistics:

$$X_{it} = c_i + \eta_i t + \sum_{j=1}^{n-1} X_{jt} + u_{it} \quad (2)$$

The stock price index of a single country is thus tested whether it can be predicted using a combination of stock prices from one or some of the other countries, and then to assess the collective efficiency of these markets.

However, this two-step cointegration test procedure has been criticized by Phillips and Ouliaris (1990) and Johansen (1988) among others, on the grounds that it is a relatively weak test of whether the series tend to "hang together" over time. The Johansen (1988, 1992) test should be preferred since it is robust to various departures from normality. In fact, the Johansen test provides maximum likelihood estimates of the cointegrating vectors, or equivalently the number of long-run relationships among a system of  $I(1)$  series. In order to assess whether the long-run movements of the stock markets are determined by some common driving fundamentals, or whether each market reacts to its own particular set of forcing variables, the hypothesis of at least one common trend among a set of  $n$  stock markets is tested by implementing the multivariate cointegration test based on the following general VAR system written in error correction form:

$$(1 - L)Y_t = \sum_{i=1}^{p-1} \Psi_i(1 - L)Y_{t-i} + \Psi_p Y_{t-p} + \mu + \varepsilon_t \quad (3)$$

where  $Y_t$  is a  $n \times 1$  column vector of the  $n$  stock market price indices at time  $t$ ,  $\Psi_i$  is a  $n \times n$  matrix of regression coefficients,  $p$  is the lag-length selected such that the residuals are serially uncorrelated,  $\mu$  is a  $n \times 1$  vector of intercepts, and  $\varepsilon$  is the vector of residuals assumed to be white noises. The hypothesis of at most  $r$  cointegrating vectors is formulated by restricting the rank of  $\Psi_p = \alpha \beta'$ , where  $\beta$  is a  $(p \times r)$  matrix of cointegrating vectors and  $\alpha$  the  $(p \times r)$  matrix of adjustment coefficients. Johansen (1988), and Johansen and Juselius (1991) show how estimation and testing of (3) can be performed by solving an eigenvalue problem and considering the resultant ordered  $n$  eigenvalues  $\lambda_i, i = 1, \dots, n$ . The likelihood ratio test of there being at least  $r$  cointegrating vector (i.e.,  $n-r$  common stochastic trends) is obtained by computing the following test statistic:

$$S = -T \sum_{i=r+1}^n \ln(1 - \hat{\lambda}_i) \quad (4)$$

The asymptotic distribution of this statistic has been simulated in Johansen (1988) and extended in Osterwald-Lenum (1992).

The model in (3) admits a linear trend in the process and the cointegrating relation. To test for the presence of a linear trend in the system, the no-linear trend system is also estimated. Defining the new estimated coefficient with a "\*", the likelihood ratio statistic suggested by Johansen and Juselius (1990),

$$LR = -T \sum_{l=r+1}^n \ln \left[ \frac{1 - \lambda_l^*}{1 - \lambda_l} \right] \rightarrow \chi_{n-r}^2 \quad (5)$$

is used to determine the bindness of the no-linear trend restriction, given  $r$  cointegrating vectors.

As in Eun and Shim (1989), comovements between emerging stock markets and the major world stock exchanges are investigated by fitting a VAR(p) model to the index returns for the 16 markets. First, we test for the causal relation between a pair of countries  $i$  and  $j$ , that is, for the null hypothesis that country  $j$ 's market index returns do not Granger-cause those for market  $i$ , the following restricted (r) and unrestricted (u) regressions are estimated:

$$R_{i,t} = \alpha_0 + \alpha_{11} R_{i,t-1} + \alpha_{12} R_{i,t-2} + \dots + \alpha_{1p} R_{i,t-p} + e_{1,t} \quad (6)$$

$$R_{i,t} = \alpha_0 + \alpha_{11} R_{i,t-1} + \dots + \alpha_{1p} R_{i,t-p} + \alpha_{j1} R_{j,t-1} + \dots + \alpha_{jp} R_{j,t-p} + e_{1,t} \quad (7)$$

and the likelihood ratio  $R = (N-c)(\ln \sigma_r^2 - \ln \sigma_u^2)$ , distributed as a  $\chi_{u-r}^2$  is used to test the null, where  $\sigma_r^2$  and  $\sigma_u^2$  are the estimated variances of the residuals in the restricted and the unrestricted regressions respectively,  $N$  is the number of observations,  $u - r$  is the number of omitted variables in the restricted system, and  $c$  is a correction to improve small sample properties and set equal to the number of variables in each unrestricted equation.

The following VAR model is used to study the interdependence structure of the  $n$  markets:

$$R_t = C + \sum_{s=1}^p \Psi_s R_{t-s} + e_t \quad (8)$$

where  $R_t$  is a  $n \times 1$  column vector of the rates of return of the  $n$  stock markets,  $\Psi_s$  is a  $n \times n$  matrix of coefficients,  $p$  the lag-length, and  $e_t$  the  $n \times 1$  column vector of forecast errors or innovations. We follow Eun and Shim (1989) and Lee (1992) by examining the system's reaction to typical random shocks. Assume that  $R_t$  is a regular covariance stationary process having the following vector moving average representation by the Wold theorem:

$$R_t = \sum_{s=0}^{\infty} B_s e_{t-s} \quad (9)$$

where  $e_t$  is a one-step-ahead linear least squares forecasting error in predicting  $R_t$  as a linear function of  $R_{t-1}, R_{t-2}, \dots$ , and is serially uncorrelated. The  $i, j$ -th component of  $B_s$  shows the response of the  $i$ -th market in  $s$  periods after a unit random shock in the  $j$ -th market (Toronto or Paris market) and none in other markets. To take into account the contemporaneous correlation between components of  $e_t$ , and to decompose unambiguously the forecast error variances into components attributable to each innovation, we use a Cholesky decomposition to obtain orthogonalized innovations,  $u_t$ , from  $e_t = Vu_t$ , where  $V$  is a lower triangular matrix such that  $E(ee') = VV'$ . Then, the  $i, j$ -th component of  $C_s$ , with  $C_s = B_sV$ , states for the impulse response of the  $i$ -th market in  $s$  periods to a random shock of one standard error in the  $j$ -th market. Furthermore, to measure the relative importance of each market in generating the fluctuations in stock returns in its own and other markets, we proceed with the decomposition of forecast error variance by computing the component of forecast error variance in  $T + 1$  step-ahead forecast of market  $i$  which is accounted for by innovations in market  $j$ , i.e.  $f_{i,j}$  given by:

$$f_{i,j} = \frac{\sum_{s=0}^T C_s^2(i,j)}{\sum_{j=1}^n \sum_{s=0}^T C_s^2(i,j)} \quad (10)$$

### 3. Results

Only results of tests conducted against an alternative consistent with fluctuations around a constant mean are presented, since the results are not qualitatively different when stationary fluctuations around a deterministic linear trend are introduced. Results presented in table 1 show unambiguously that the null hypothesis of a unit root in stock index levels cannot be rejected at 5% since all statistics are greater than their critical values. Moreover, the consistent rejection of the null of unit root in all the first differences of stock prices, leads to the conclusion that the stock prices in the nine emerging markets and the seven developed countries index series are integrated of first order, i.e., all series are  $I(1)$ . Using the Phillips-Ouliaris (PO) pairwise cointegration procedure, the investigation of the joint weak form efficiency of each stock market with the Toronto stock exchange and with the Paris stock market on the other hand reveals that the hypothesis of no-cointegration between the pairs of countries cannot be rejected, except the pair U.K.-France. In other words, stock prices of any studied market cannot be predicted using those of the Toronto or the Paris stock market price index series. Only the U.K. and the Paris stock price index series shows a significant pairwise comovements. In the last column of table 1, the ADF higher order cointegration test statistics are displayed, and the results provide no evidence that a stock price series of a single country can be predicted using the stock price series of all the other countries.

To look for the number of stochastic trends in the international stock markets, we consider five different systems of country-markets. The results of the Johansen's multivariate test based on VAR(4) for the presence of cointegrating vectors reveal that in system 1, composed by all the seven developed markets and the nine emerging markets, the trace statistic rejects  $r \leq 11$  but not  $r \leq 12$  or higher, providing evidence of 12 cointegrating vectors and four stochastic or common trends in the system. The tests reveal the evidence of three linearly independent cointegrating vectors in system 2 (North and Latin American plus Asian markets), and six common stochastic trends in system 3 (European, African and Asian markets). There are two cointegrating relations in the system of the nine emerging markets (system 4), while all linear combinations of the seven developed markets price series (system 5) are I(1). To assess whether the data contain a stochastic linear trend, the likelihood ratio (LR) statistics ( $\chi^2_{n-r}$  displayed in table 2) indicate that a null hypothesis of no linear trend cannot be rejected in system 1 (LR = 1.0 given  $r = 12$ ), system 3 (LR = 0.4 given  $r = 4$ ), system 4 (LR = 7.0 given  $r = 2$ ) and in system 5 (LR = 7.2 given  $r = 0$ ). Only system 2, composed by the North and Latin America plus Asia markets, rejects the no linear trend hypothesis since the LR statistic equals to 14.2 given  $r = 3$ , slightly higher than the 5% critical value of 14.07.

The Granger causality test results for the hypothesis that current stock price changes in a given market  $i$  is caused by past stock price changes in one (Toronto, Paris, U.S. or U.K) or a set of other markets (Toronto plus Paris, all others countries except the one being caused) are reported in table 3. The tests of the causal relations suggests that past stock price changes in the Toronto stock market significantly Granger-cause current stock price changes in the Mexico market and in all the others. The Paris stock market Granger-causes three emerging markets (Chile, Mexico and Zimbabwe) and Italy. The U.K. market, Granger-caused by the U.S. market and the joint Toronto and Paris markets, does not influence none of the other markets. Panels A, B and C of table 4 present the normalized impulse responses of the column markets to a unit shock (positive residuals of one standard deviation unit) in the U.S., the Toronto and the Paris markets, respectively, on the  $i$ -th month ( $i = 0$  to 3) after a shock occurred.<sup>1</sup> Examination of table 4 reveals that, unlike the nine emerging markets, the other developed markets respond most strongly to a U.S. shock on the month of the shock and the adjustment is almost completed by the next month. For example, the impulse responses of the Toronto and the Paris stock exchanges to a U.S. shock are respectively 0.71 and 0.46 on the month  $i = 0$ , and 0.14 and 0.26 on the following month. Among the emerging markets, only Mexico and Thailand display a noticeable response to a U.S. shock.

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1. Note that the order the countries are displayed in table 4 is the orthogonalization order in the estimation.



A unit shock in the Toronto stock exchange has some responses in the emerging markets, especially in Argentina, Mexico and Zimbabwe on month  $i = 0$ , and in India and Chile on month  $i = 1$ . There is no noticeable responses in any emerging market to a shock in the Paris stock exchange. Among the developed markets, the most important reactions to a unit shock in the Toronto stock market are observed in the U.S., the U.K. and the Germany markets with one month delay (response of 0.23, 0.17 and 0.18 respectively at  $i = 1$ ), and in Paris market at  $i = 0$ . Only the Italy market responds noticeably to the unit shock in Paris market but with one and two month delay (response of 0.19 and 0.11 at  $i = 1$  and  $i = 2$  respectively), and the Germany market has the most spontaneous response (0.14 at  $i = 0$ ).

Table 5 displays the decomposition of 2 and 4 month ahead forecast error variances of each market returns into apportionments attributable to their own (diagonal) and to innovations in the other markets. Table 5 indicates the existence of interaction among different national markets since none of the market's own innovation fully accounts for its variance. However, the results show that much of the forecast variances in emerging markets are explained by their own innovations, while the U.S. market appears to be the most influential among the developed markets. The influences of the Toronto and the Paris stock markets are negligible since innovations in Toronto (Paris) explain at most 4.22% (3.89%) and 5.45% (4.38) respectively of the 2 and 4 month forecast variance of the other markets. The influences of the Toronto and the Paris markets are observed mostly in India and Italy, respectively.

## 4. Conclusion

Using a unit root and cointegration investigation techniques, this paper provides evidence of the weak-form individual and collective efficiency of the seven industrialized countries and nine emerging markets. The results of the Johansen (1988, 1992) tests indicate the presence of four common stochastic trends in the system of 16 countries. The investigation of the causal relations and dynamic interactions among stock returns based on a VAR analysis reveals that shocks in the Toronto stock exchange has responses mainly in the U.S., the U.K., Germany and in some emerging markets (Argentina, Chile, India, Mexico and Zimbabwe), while Germany and Italy markets react to shocks in Paris stock exchange. The influences of the Toronto and the Paris markets are observed mostly in India and Italy respectively but remain negligible. The study also confirms the results of Eun and Shim (1989) on the leading role of the U.S. stock market.

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TABLE 1

Results of the Unit Root Tests (URT) and cointegration analysis of monthly stock prices in nine emerging and seven major markets. Analysis are done from 1975:12 to 1992:12. The statistics displayed are the Augmented Dickey-Fuller (ADF), the Phillips-Perron (PP), and the Z-Phillips-Ouliaris (PO) test. Significance at the 10% and 5% levels are denoted by \* and \*\*, respectively.

Countries	URT: Price indices		URT: First differences of price indices		PO pairwise test of cointegration			ADF higher order cointegration
	ADF	PP	ADF	PP	Canada	France	Others	
Argentina	5.26	-0.23	-6.1**	-12.4**	-4.4	-6.3	-3.76	
Brazil	1.23	19.06	-5.4**	-0.3	36.5	29.4	-3.17	
Chile	1.99	2.75	-2.6*	-10.7**	1.5	-1.5	-2.95	
Greece	-1.68	-1.53	-3.8**	-12.4**	-7.4	-9.9	-3.13	
India	0.28	-0.55	-2.1	-12.2**	-4.3	-9.7	-4.89	
Korea	-0.69	-0.70	-3.1**	-16.6**	-8.0	-15.9	-3.81	
Mexico	0.18	2.77	-3.1**	-10.8**	2.9	0.9	-4.65	
Thailand	0.68	0.25	-2.7*	-11.4**	-0.7	-5.9	-3.59	
Zimbabwe	-1.13	-0.87	-3.7**	-10.2**	-8.0	-12.0	-3.39	
U.S.	1.63	0.83	-3.9**	-10.1**	1.0	-7.9	-3.17	
U.K.	0.38	-0.05	-4.4**	-10.8**	-7.9	-20.7**	-3.69	
Japan	-1.28	-1.10	-3.6**	-9.9**	-8.5	-3.6	-3.23	
Canada	-1.37	-1.41	-4.9**	-13.5**	-	-11.1	-3.55	
France	-0.46	-0.30	-4.2**	-11.1**	-8.9	-	-4.02	
Germany	-1.01	-1.10	-3.8**	-10.1**	-12.3	-8.7	-3.87	
Italy	-1.26	-1.17	-3.5**	-11.5**	-13.7	-4.5	-3.60	

TABLE 2

Multivariate test for the presence of cointegrating vectors and linear trends. Determination of the number ( $r$ ) of cointegrating vectors present in five different systems. System 1 is composed with the seven developed markets and the nine emerging markets, System 2 with North and Latin American and Asian markets, System 3 with European, African and Asian markets, System 4 with the nine emerging markets, and System 5 with the seven developed markets. Monthly stock market index prices are from 1975:12 to 1992:12. Critical values are from Osterwald-Lenum (1992). Significance at the 10% and 5% levels are denoted by \* and \*\*, respectively.

$r$	System 1	System 2	System 3	System 4	System 5
0	906.78**	266.86**	290.89**	235.58**	114.37
1	765.90**	202.70**	226.36**	168.24**	70.19
2	643.55**	151.91*	170.97**	117.11	41.33
3	536.96**	113.75	121.70*	73.28	25.87
4	447.87**	83.57	76.70	46.16	11.82
5	367.32**	57.03	48.24	25.95	5.25
6	291.15**	35.93	29.77	12.84	0.03
7	225.50**	18.86	15.76	4.96	
8	178.16**	8.75	5.08	0.00	
9	135.16**	0.60	0.11		
10	100.49**				
11	68.53*				
12	43.12				
13	26.80				
14	13.15				
15	1.26				
$\chi_{\alpha,r}^2$	1.00	14.20**	0.40	7.00	7.20

TABLE 3

Likelihood ratios for the Granger causality tests based on VAR(4). The test statistics  $R = (N - c)(\ln \hat{\sigma}_r^2 - \ln \hat{\sigma}_u^2) \rightarrow \chi_{c-1}^2$  are displayed, where  $\hat{\sigma}_r^2$  and  $\hat{\sigma}_u^2$  are the restricted and unrestricted residual variances, N the number of observations, and c is a correction coefficient equal to the number of variables in each unrestricted equation. Significance at the 10% and 5% levels are denoted by \* and \*\*, respectively.

	Canada	France	Canada + France	U.S.	U.K.	All others
Argentina	1.85	2.55	5.05	4.30	2.17	50.61
Brazil	2.38	1.13	4.30	0.31	5.03	44.89
Chile	4.13	15.65**	16.85**	7.64	2.96	50.67
Greece	3.73	0.48	4.12	3.76	4.70	61.49
India	3.02	1.52	6.61	2.86	0.16	57.81
Korea	4.85	5.18	7.59	5.76	3.74	68.20
Mexico	12.39**	9.16*	21.77**	11.15**	5.14	65.22
Thailand	6.79	1.26	6.83	7.50	0.83	61.26
Zimbabwe	2.61	15.01**	14.82*	2.88	4.14	46.97
U.S.	14.14**	5.18	17.85**	-	0.86	51.05
U.K.	7.57	5.43	17.60**	12.77**	-	55.50
Japan	21.65**	8.88*	23.09**	7.44	2.31	71.06
Canada	-	4.91	-	5.42	0.54	38.27
France	8.95*	-	-	10.74**	3.87	65.08
Germany	16.69**	3.63	17.51**	12.69**	5.95	91.98**
Italy	8.46*	17.64**	20.83**	9.19*	2.68	50.96

TABLE 4

Impulse responses to the unit shock in the selected stock markets.  
 Sample period: 1976:1-1992:12.  $i^*$  is the  $i$ -th month after the shock.

$i^*$	U.S.	U.K.	JAP.	CAN.	FRA.	GER.	ITA.	ARG.	BRA.	CHI.	GRE.	IND.	KOR.	MEX.	THA.	ZIM.
Panel A: Impulse responses to the unit shock in the U.S. market																
0	1.00	0.49	0.32	0.71	0.46	0.41	0.27	0.05	0.07	0.06	0.15	-0.05	0.11	0.27	0.19	-0.01
1	0.39	0.42	0.22	0.14	0.26	0.42	0.17	0.14	-0.01	-0.02	0.15	0.06	-0.15	0.31	0.12	0.01
2	0.00	-0.06	0.12	-0.12	-0.12	0.08	0.03	0.07	0.00	-0.07	-0.05	0.09	-0.11	0.04	0.01	-0.05
3	-0.09	-0.11	-0.02	0.05	0.00	-0.01	0.04	-0.03	-0.06	-0.18	0.13	0.00	0.08	-0.06	-0.04	0.05
Panel B: Impulse responses to the unit shock in the Toronto stock market																
0	0.00	0.00	0.00	0.70	0.10	-0.04	-0.01	0.10	-0.13	0.06	-0.05	0.08	0.08	0.12	-0.02	0.11
1	0.23	0.17	0.04	-0.08	0.08	0.18	0.05	-0.12	-0.07	0.14	-0.02	0.21	-0.12	0.09	0.04	0.12
2	-0.05	0.04	0.07	0.03	-0.02	0.06	0.02	0.07	0.07	0.02	-0.12	-0.14	0.02	0.08	-0.01	0.08
3	-0.13	-0.02	0.03	0.00	0.05	0.07	0.07	-0.05	-0.13	0.09	-0.14	-0.06	0.05	-0.03	-0.20	0.07
Panel C: Impulse responses to the unit shock in the Paris stock market																
0	0.00	0.00	0.00	0.00	0.87	0.14	-0.10	0.00	0.04	-0.07	0.03	0.08	-0.12	0.03	0.02	-0.04
1	0.03	-0.03	0.06	-0.01	-0.04	-0.04	0.19	-0.09	0.00	0.01	-0.11	-0.14	-0.13	0.01	-0.05	0.03
2	-0.10	-0.07	-0.17	-0.04	-0.10	-0.02	0.11	0.01	-0.06	-0.06	-0.02	-0.04	0.09	0.06	0.08	-0.02
3	0.08	0.02	-0.03	0.10	-0.03	-0.02	-0.01	0.15	0.06	-0.08	-0.01	0.07	0.04	-0.07	-0.02	0.21

**TABLE 5**  
 Decomposition of forecast error variances: Percentage of 2- and 4-month-ahead forecast error variance of the row stock market returns explained by innovations in the column markets. Study based on VAR(4) from 1976:1 to 1992:12.

	U.S.	U.K.	JAP.	CAN.	FRA.	GER.	ITA.	ARG.	BRA.	CHI.	GRE.	IND.	KOR.	MEX.	THAI.	ZIM.
U.S.	2 89.19	0.00	0.03	4.12	0.05	0.28	0.11	1.18	0.48	0.72	0.04	0.20	0.22	2.68	0.01	0.69
	4 78.48	0.01	0.29	4.86	1.23	0.46	0.60	2.14	1.64	0.71	2.11	2.84	0.29	2.66	0.76	0.91
U.K.	2 31.23	58.76	3.57	2.13	0.05	0.02	0.23	0.37	0.12	0.21	0.12	0.02	0.06	1.12	1.90	0.09
	4 29.20	53.83	3.70	2.05	0.43	0.32	0.21	0.70	0.54	1.05	2.23	0.68	0.23	2.08	2.54	0.22
JAP.	2 11.97	3.37	71.53	0.13	0.30	2.01	2.18	0.48	0.28	0.67	0.22	2.73	3.37	0.01	0.03	0.73
	4 10.92	3.20	60.15	0.53	2.15	2.01	2.05	0.52	5.41	0.66	2.37	2.58	5.02	0.60	1.21	0.64
CAN.	2 49.63	0.02	0.63	45.97	0.00	0.05	0.42	0.19	0.00	1.21	0.67	0.21	0.01	0.60	0.06	0.33
	4 45.88	0.03	1.22	41.22	0.90	1.19	0.67	0.59	0.73	1.64	2.81	0.91	0.30	0.94	0.30	0.67
FRA.	2 24.18	0.75	1.13	1.34	64.93	0.24	0.43	0.20	0.08	0.97	0.33	0.01	2.94	0.38	1.64	0.46
	4 21.38	1.79	1.12	1.34	55.99	1.58	0.94	1.35	1.52	0.82	0.90	0.28	2.86	3.18	2.23	2.91
GER.	2 23.64	4.24	1.02	2.21	1.40	52.37	0.29	0.16	3.74	0.24	0.05	1.04	1.77	5.84	2.00	0.00
	4 20.42	4.77	1.89	2.39	1.22	46.79	0.49	1.60	4.50	3.21	1.30	1.49	1.70	5.17	3.05	0.01
ITA.	2 9.19	2.90	1.08	0.25	3.89	1.68	77.06	0.00	0.40	0.17	1.66	0.37	0.32	0.14	0.15	0.73
	4 8.29	3.98	2.24	0.68	4.38	1.64	68.04	0.42	0.92	0.53	3.63	0.84	1.82	0.13	0.66	1.78
ARG.	2 1.94	0.13	0.37	2.30	0.73	0.01	1.11	91.88	0.13	0.00	0.34	0.48	0.03	0.34	0.04	0.15
	4 1.93	0.35	1.94	2.32	2.25	0.54	3.84	74.81	4.78	0.06	1.54	1.83	0.25	1.28	0.39	1.87
BRA.	2 0.48	0.55	0.23	1.92	0.15	1.90	0.09	3.44	89.52	0.30	0.11	0.13	0.14	0.41	0.55	1.13
	4 0.66	1.06	1.34	3.45	0.72	1.68	2.42	3.85	80.18	0.93	1.10	0.14	0.41	0.55	1.13	0.37
CHI.	2 0.42	2.19	0.01	2.06	0.42	1.36	1.63	0.10	0.92	88.87	0.04	0.36	0.16	1.22	0.13	0.10
	4 3.06	1.94	0.98	2.26	1.04	1.64	2.59	0.71	1.24	77.53	1.42	2.11	1.28	1.19	0.51	0.50
GRE.	2 3.89	1.44	2.39	0.27	1.02	1.94	0.32	0.39	0.94	2.00	79.01	0.02	0.27	2.21	3.39	0.49
	4 4.56	3.69	2.17	2.69	0.86	2.31	1.57	0.66	1.09	1.70	64.28	3.78	0.56	5.98	3.44	0.68
IND.	2 0.55	0.10	2.81	4.22	2.12	1.46	3.19	1.16	3.07	0.23	0.44	77.47	0.86	1.36	0.85	0.00
	4 1.08	1.63	2.64	5.45	2.25	1.37	3.39	2.31	4.62	0.22	0.90	67.77	1.30	2.57	1.52	1.00
KOR.	2 2.91	3.56	3.52	1.90	2.73	0.65	1.20	3.42	0.36	0.70	1.16	1.03	73.96	0.05	0.85	2.01
	4 3.63	4.64	4.24	1.72	2.84	0.76	5.16	5.13	2.19	0.70	2.39	2.43	60.87	0.62	1.08	1.69
MEX.	2 13.04	2.10	0.59	1.69	0.08	0.03	2.41	1.23	2.08	0.15	0.38	0.23	0.12	74.34	0.17	1.36
	4 11.65	2.61	1.01	1.93	0.65	0.98	2.38	1.07	2.24	0.51	5.62	1.74	0.32	64.49	0.65	2.13
THAI.	2 4.53	0.06	5.38	0.15	0.27	0.85	1.83	1.40	0.28	1.23	3.06	1.15	2.72	5.72	72.34	0.12
	4 3.76	0.26	7.43	3.19	0.87	2.15	3.32	1.40	0.24	1.73	3.27	3.84	3.08	4.60	60.34	0.71
ZIM.	2 0.02	0.04	0.67	2.48	0.28	0.82	2.60	1.17	0.39	5.70	1.01	0.68	0.65	0.50	3.41	79.58
	4 0.38	1.69	2.17	2.87	3.69	1.00	3.08	1.59	0.41	5.95	1.94	2.31	0.93	1.03	3.05	67.90