A Test of the Proxy-Effect Hypothesis: Evidence From the Pacific Basin Countries

Unro Lee
University of the Pacific

This study examines whether the proxy-effect hypothesis can adequately explain the negative relationship between stock returns and expected inflation for the four Pacific Basin countries—Hong Kong, Singapore, South Korea, and Taiwan. The extent to which inflation rate proxies expected economic activity and inflation variability in the stock return-inflation relationship is investigated. Both monthly and quarterly data are employed in this study. We find that negative and statistically significant relationship between stock returns and inflation rate still persists for these four countries even after the effects of expected economic activity and inflation variability have been explicitly incorporated. Therefore, the findings of this study do not support the proxy-effect hypothesis for the four Pacific Basin countries. Furthermore, these results are consistent with those of other studies that have also rejected the proxy-effect hypothesis for major industrialized countries.

Introduction

Many researchers have documented the negative relationship between stock returns and expected, unexpected, and changes in the inflation rate during the post-Korean War period. Bodie (1976), Nelson (1976), and Fama and Schwert (1977) study the United States, while other industrialized countries have been examined by Solnik (1983), Gultekin (1983), and Lee (1996). Kwok and Li (1993) and Kwok (1994) analyze the Pacific Basin countries.

These findings refute the traditionally held belief that common stocks are an effective hedge against unexpected inflation because stocks represent ownership of the perpetual income generated by real assets. Also these findings mitigate the validity of the *Fisher theory of interest rate* that posits positive relationships between stock returns and both expected inflation and changes in the expected inflation.

Fama (1981) was the first to introduce the *proxy-effect hypothesis* to explain the prevalence of negative stock return-inflation phenomenon. The central tenet of Fama's version of the proxy-effect hypothesis is that the observed negative relationship between inflation and stock returns is spurious because this relationship

0747-5535/98/1400-0040 $2.50

is induced by a positive relationship between stock returns and expected economic activity and an inverse relationship between expected economic activity and inflation. Inflation is merely serving as a proxy for expected economic activity in a statistical relationship between stock returns and inflation.

According to the theory of finance, stock prices reflect the present value of future dividends. Because dividends are determined by firms' earnings (that in turn depend on real economic activity), stock returns should be positively related to the expected economic activity.[1]

The *rational expectations* version of the quantity theory of money that Fama espoused in his 1981 paper predicts that a downward revision of the expected output leads to an immediate decrease in the demand for money. If both the money supply and interest rates are exogenous, excess supply of money would immediately trigger an increase in current and expected inflation. Geske and Roll (1983) and Kaul (1987, 1990) further demonstrate that if the central bank implements a (deficit-induced) counter-cyclical monetary policy, a downward revision of the expected output would still signal upward movements in current and expected inflation. This happens because a decline in real economic activity leads to increased deficits that, in turn, induce the central bank to *monetize* the federal debt. Increased money supply then triggers an increase in current and expected inflation.

Empirical evidence thus far does not lend overwhelming support for Fama's version of the proxy-effect hypothesis for both the United States and other industrialized countries. Although earlier studies (Fama, 1981; Geske and Roll, 1983; and Kaul, 1987) provide some support for the proxy-effect hypothesis, more recent empirical studies (McCarthy, Najand, and Seifert, 1990; Liu, Hsueh, and Clayton, 1993; Cochran and DeFina, 1993; and Lee, 1996) reject the hypothesis.

Kaul and Seyhun (1990) recently have offered another version of the proxy-effect hypothesis. They contend that the negative stock return-inflation relationship is a combined effect of a negative relationship between inflation variability and stock returns and a positive relationship between inflation variability and inflation rate. Kaul and Seyhun claim that inflation serves as a proxy for inflation variability in the stock return-inflation relationship.

Friedman (1977) argues that an increase in inflation variability engenders costly institutional changes, such as more frequent contracting, that would diminish the productive efficiency of the firms and therefore impede the real economic activity. Because stock returns and expected economic activity are positively correlated, inflation variability will have a deleterious effect on stock returns. Malkiel (1979)

[1] The positive relationship between stock returns and expected economic activity is well documented in the literature. For example, Fama (1981) finds that approximately 68 percent of the variations in the U.S. stock returns for the 1953-1977 period can be explained by future changes in economic activity.

also reports that a rise in inflation variability increases the riskiness of stocks which in turn reduces stock returns.

There is ample empirical evidence that justifies a positive relationship between inflation variability and expected inflation, unexpected inflation, and actual inflation (Cukierman, 1983).

Although Kaul and Seyhun's findings strongly support their version of the proxy-effect hypothesis for the United States, Lee (1996) finds that Kaul and Seyhun's version of the proxy-effect hypothesis is rejected for both U.S. and other industrialized countries.

These studies testing the proxy-effect hypothesis to date have focused on the industrialized countries. Given the inconclusiveness of the empirical results across major industrialized countries, it would be interesting to test both Fama's and Kaul and Seyhun's versions of the proxy-effect hypothesis for emerging stock markets as well.

In light of spectacular economic success enjoyed by Hong Kong, Singapore, South Korea, and Taiwan—four Pacific Basin countries collectively referred to as *four little tigers*—since the early 1980s, literature on the intertemporal behavior of these Pacific Basin stock markets is surprisingly scarce.

The main thrust of this study is to test both versions of the proxy-effect hypothesis for these four Pacific Basin countries. The extent to which inflation rate proxies expected economic activity and inflation variability in the stock return-inflation relationship is investigated.

Average annual (nominal) GNP growth was approximately 9 percent for these four countries between 1980 and 1989, compared to 3 percent, 4 percent, and 2 percent for the United States, Japan and the European Economic Community (EEC), respectively, during the same period. Due to the spectacular economic growth achieved by these four countries, stock market capitalization for these countries grew 59 percent per year on average during the 1980s. The aggregate (nominal) stock prices increased 92 percent, 124 percent, 750 percent, and 1600 percent between 1980 and 1989 for Hong Kong, Singapore, South Korea. and Taiwan, respectively (Kwok and Li, 1993).

Both monthly and quarterly data are employed in this study. We find that a negative and statistically significant relationship between stock returns and the inflation rate persists for these four countries even after the effects of expected economic activity and inflation variability have been explicitly incorporated. Therefore, the findings of this study do not support the proxy-effect hypothesis for the four Pacific Basin countries. These results are consistent with the findings of Cochran and Defina (1993), Lee (1996), Liu, Hsueh, and Clayton (1993), and McCarthy, Najand, and Seifert (1990). These other researchers also reject both

versions of the proxy-effect hypothesis for major industrialized countries such as Belgium, France, Germany, Japan, the United Kingdom, and the United States.[2]

Data and Methodology

Monthly data for stock prices, the consumer price index (CPI), and an index of industrial production are obtained for South Korea and Taiwan. Quarterly data for the same variables are obtained for Hong Kong, Singapore, South Korea, and Taiwan.[3] The period investigated with monthly data is January 1978 through June 1995 and January 1982 through June 1995 for South Korea and Taiwan, respectively. The period investigated with quarterly data is 1983:I to 1995:II, 1983:I to 1995:II, 1980:I to 1995:II, and 1982:I to 1995:II for Hong Kong, Singapore, South Korea, and Taiwan, respectively. The sample periods are incongruous, and the sample period is relatively short due to the dearth of financial data for the Pacific Basin countries.

The source of data for Hong Kong is *the Hong Kong Monthly Digest of Statistics*. The data for Singapore are from *the Singapore Monthly Digest of Statistics*. The data for South Korea are obtained from various issues of the *Monthly Statistical Bulletin* and the *Economic Statistics Yearbook* published by the Bank of Korea. The data for Taiwan are obtained from various issues of *the Financial Statistics Monthly (Taiwan District)* and *Monthly Statistics of the Republic of China*, both published by the Central Bank of China

Stock indexes used in this study are the Hang Seng Index for Hong Kong, the Strait Times Industrial Index for Singapore, the Korea Composite Stock Price Index for South Korea, and the Taiwan Stock Exchange Index for Taiwan. The Hang Seng Index is a value-weighted index of 33 stocks traded at the Stock Exchange of Hong Kong. These 33 stocks represent approximately 70 percent of the exchange's total market capitalization. The Strait Times Index of the Stock Exchange of Singapore is an unweighted index of 30 stocks. The Korea Composite Stock Price Index is a value-weighted index of all common stocks traded on the Korea Stock Exchange.

---

[2] Cochran and Defina (1993) employ both quarterly and annual U.S. data for the period 1947 to 1989 in their study. Lee (1996) uses monthly and quarterly data for Belgium, France, Germany, Japan, the United Kingdom, and the United States for the period 1976 to 1992. Liu, Hsueh, and Clayton (1993) use monthly and quarterly data for Canada, Germany, the United Kingdom, and the United States for the period 1974 to 1990. McCarthy, Najand and Seifert (1990) use monthly data for Germany, the United Kingdom, and the United States for the period 1962 to 1987. All four papers have tested the proxy-effect hypothesis by estimating an empirical model that is either identical or similar to equations (1) through (4) in this paper. All four papers apply the OLS technique to their models.

[3] Monthly data for stock prices, the consumer price index (CPI), and industrial production are not available for Hong Kong and Singapore.

The Taiwan Stock Exchange Index is a value-weighted index of all stocks traded on the Taiwan Stock Exchange (Martin and Wood, 1996).

The Stock Exchange of Hong Kong is one of the least regulated stock markets in the world. There are no government controls or restrictions on the terms of transactions. Although the Korean government recently has taken significant steps to deregulate Korean capital markets, the Korea Stock Exchange (KSE), the only stock exchange in Korea, is owned and managed by the Korean government. The Stock Exchange of Singapore is heavily regulated by the Singapore central bank—the Monetary Authority of Singapore. The Taiwan Stock Exchange is the second largest stock market in the Pacific-Basin, next to Japan. Although both Hong Kong and Singapore have aggressively attracted foreign capital recently by either removing or relaxing interest rate regulations and abolishing exchange rate controls, South Korea and Taiwan have imposed severe restrictions on foreign ownership of their securities (Ko and Lee, 1991).

Following Kaul and Seyhun (1990), the proxy-effect hypothesis is tested in this study by estimating the following set of equations:

(1) $ERSR_t = \alpha + \beta_1 EI_{t-1} + \varepsilon_t$

(2) $ERSR_t = \alpha + \beta_1 EI_{t-1} + \beta_2 ERA_{t+1} + \varepsilon_t$

(3) $ERSR_t = \alpha + \beta_1 EI_{t-1} + \beta_3 INFVAR_t + \varepsilon_t$

(4) $ERSR_t = \alpha + \beta_1 EI_{t-1} + \beta_2 ERA_{t+1} + \beta_3 INFVAR_t + \varepsilon_t$

where:

$ERSR_t$ = Expected real stock return for period t:

EI_{t-1} = Expected inflation formed at period t-1 for period t;

ERA_{t+1} = Expected economic activity for period t+1; and

$INFVAR_t$ = Inflation variability at time t.

β_1 in the equation (1) should be negative and statistically significant if there is a negative stock return-inflation relationship. If the Fama version of the proxy-effect hypothesis is valid, however, then β_1 should lose its statistical significance in equation (2) whereas β_2 should be statistically significant and positive. If Kaul and Seyhun's version of the proxy-effect hypothesis is valid, then β_1 should lose its statistical significance in equation (3) whereas β_3 should be statistically significant and negative. If both versions are valid, then β_1 should not be statistically significant in equation (4) if both ERA_{t+1} and $INFVAR_t$ are introduced.[4]

[4] Most of the empirical studies investigating the stock return-inflation relationship for both the U.S. and other countries have been conducted within a partial equilibrium framework such as the one employed in this study. The major drawback of this approach, however, is that it ignores the dynamic interaction of stock returns, the inflation rate, and real economic activity (Stulz, 1986). Utilizing a vector autoregressive (VAR) framework (James, Koreisha, and Partch, 1985; Lee, 1992) may circumvent this problem when analyzing the stock return-inflation relationship.

The VAR approach is not a proper empirical apparatus for testing the proxy-effect hypothesis across international stock markets for two reasons. First, VAR procedure is more

Nominal stock returns and inflation rate are calculated as the logarithmic difference of the stock prices and CPI, respectively, between two consecutive periods. Inflation rates are subtracted from nominal stock returns to obtain real stock returns. The rate of growth of the index of industrial production is used as a proxy for real economic activity.

The Box-Jenkins method is employed to extract estimates for expected inflation, expected real stock returns, and expected real economic activity from the fitted ARIMA models.[5] In all cases the estimated autoregressive-moving average parameters are statistically significant at the 5 percent level. Furthermore, the reported Q-statistics indicate that the null hypothesis of white noise residuals cannot be rejected at the 5 percent level for all fitted ARIMA models.[6]

The estimates of inflation variability are obtained from the GARCH (generalized autoregressive conditional heteroscedastic) procedure. The GARCH model proposed by Engle (1982) and later generalized by Bollerslev (1986) has been used extensively in the literature as a means to investigate the dynamic behavior of financial asset return volatilities over time. The GARCH model posits that the conditional variance of a variable is a linear function of both the past shocks (or innovations) and past conditional volatilities. The GARCH model captures the adaptive learning process of economic agents.

The conditional variance of the inflation rate extracted from the estimated GARCH model is used as a measure of inflation variability.[7] Engle (1982) also utilizes the GARCH procedure to extract inflation variability for the United Kingdom.[8]

---

appropriate as a means of conducting the Granger causality test because this procedure measures the impact of an innovation or *shock* in one variable on the forecast error variance of another variable. James, Koreisha, and Partch and Lee have focused on the causal relationships among these variables. Second, simultaneous estimation of four VAR models, one for each of the four countries investigated in this study, can be computationally burdensome if not impossible. James, Koreisha, and Partch and Lee employ U.S. data only.

[5] Fama and Gibbons (1984) have found that the interest-rate model provides a more accurate inflation rate forecasts than a ARIMA model for the United States during the period 1953-1977. Hafer and Hein (1990), using monthly data for the United States and five other industrialized countries, report that the ARIMA forecasts of inflation have equal or lower forecast errors than the interest-rate forecasts for the period 1967-1986.

[6] Details of the ARIMA specifications of the variables used in this study are available from the author upon request.

[7] The GARCH model of the inflation rate is estimated with the maximum likelihood estimation (MLE) method proposed by Bollerslev (1986). We find that the conditional variance equation of the inflation rate for all four countries can be best characterized as a GARCH (1,1) process. Details of the estimated GARCH models are available from the author.

[8] The GARCH procedure is chosen over the survey data as a means of extracting inflation variability in this study for two reasons. First, the GARCH method captures the changes in the

Empirical Results

Estimation of equations (1) through (4) presupposes that the variables in the regression equations are stationary processes. If any or all of these variables are nonstationary processes, however, then statistical inferences that can be drawn from these *spurious* regression equations would not be valid (Granger and Newbold, 1974). Therefore it is necessary to ascertain whether the variables in equations (1) through (4) are all stationary processes before we test for the proxy-effect hypothesis.[9]

The test of null hypothesis of first-order nonstationarity is performed using a statistical procedure proposed by Dickey and Fuller (1979, 1981). First, the following regression equation for each variable is estimated:

$$(5)\ \Delta Y_t = \alpha + \gamma t + \rho Y_{t-1} + \sum_{i=1}^{k} \beta_i \Delta Y_{t-i} + \xi_t,$$

where:

Y = The variable under consideration.

The order of k should be set large enough to ensure the estimated residual series, ξ_t, becomes white noise. The null hypothesis that Y is a first-order nonstationary variable is rejected if the estimated coefficient of the lagged variable, ρ, is significantly less than zero. The estimated t-statistic for ρ, t_ρ, is referred to as the *augmented Dickey-Fuller* (ADF) statistic.

Because the results of the Dickey-Fuller test can be sensitive to the choice of the lag length k (Thornton and Batten, 1985), the Akaike (1974) information criterion (AIC) is employed to determine the optimal lag-structure specification of equation (5). AIC searches for the lag length k that would minimize

$$(6)\ AIC(k) = N \ln (SSR(k)/N) + 2p,$$

where:

N = The number of observations;

SSR(k) = The sum of squared residuals of the regression with k as the lag length; and

p = The number of regressors (which is equal to k+1).

The appropriate order of the lag length in equation (5) is obtained by computing AIC(k) over the range of values for k from 1 to 12 for monthly data and from 1 to 4 for quarterly data. The kth lag that minimizes AIC, k*, is provided for each of the four variables for all four countries in Table 1.

---

volatility of inflation over time more precisely than the cross-sectional variance of the survey forecasts (Engle, 1982; Lahiri, Teigland, and Zaporowski, 1988). Second, monthly and quarterly survey data for inflation rate for the Pacific Basin countries are not available.

[9] The stationary process is a process whose joint and conditional distributions are both time invariant. If a series is stationary, then both mean and variance of the series should be stationary.

Table 1—Augmented Dickey-Fuller (ADF) Test Results for Expected Stock Return, Expected Inflation Rate, Real Economic Activity, and Inflation Variability for Hong Kong, Singapore, South Korea and Taiwan: Monthly and Quarterly Results

	Y	k*	tt(k*)
Panel A: Monthly Results			
South Korea	ERSR	1	-8.7560**
	EI	8	-4.1149**
	ERA	5	-6.2563**
	INFVAR	6	-3.9898**
Taiwan	ERSR	1	-7.7067**
	EI	1	-11.7512**
	ERA	10	-6.3861**
	INFVAR	6	-7.8639**
Panel B: Quarterly Results			
Hong Kong	ERSR	1	-4.5220**
	EI	3	-3.5239*
	ERA	3	-5.6552**
	INFVAR	3	-3.6149*
Singapore	ERSR	4	-4.5823**
	EI	3	-4.9895**
	ERA	3	-4.1892**
	INFVAR	3	-5.1782**
South Korea	ERSR	1	-4.5638**
	EI	2	-4.0873**
	ERA	1	-4.6672**
	INFVAR	3	-6.0816**
Taiwan	ERSR	1	-4.4162**
	EI	2	-5.8485**
	ERA	3	-4.2465**
	INFVAR	3	-4.6352**

Note: The null hypothesis of first-order nonstationarity for each of the four variables is tested with the ADF statistic, the estimated t-statistic for ρ in the regression equation:

$$\Delta Y_t = \alpha + \gamma t + \rho Y_{t-1} + \sum_{i=1}^{k} \beta_i \Delta Y_{t-i} + \xi_t,$$

where:

Y = The variable under investigation;
ERSR = Expected real stock returns;
EI = Expected inflation;
ERA = Expected real economic activity; and
INFVAR = Inflation variability

* and ** denote the significance at the 5 percent and 1 percent level, respectively

The critical values for the t statistic are -3.13 (10 percent), -3.43 (5 percent), and -3.99 (1 percent) (Fuller, 1976). k*, the optimal lag-length , is determined by the Akaike information criterion

Because the ADF statistic does not necessarily follow the usual Student t-distribution under the null hypothesis of first-order nonstationarity, the tabulated values provided in Fuller (1976) should be used.

Table 1 shows that the null hypothesis of a first-order nonstationarity process can be rejected for all four variables (i.e., expected real stock return, expected

inflation, expected real economic activity, and inflation variability) at the 1 percent significance level for both South Korea and Taiwan when monthly data are used.

The results obtained with quarterly data are similar to those obtained with monthly data for all four countries. The null hypothesis of a first-order nonstationarity process is rejected at the 1 percent significance level for all four variables for Singapore, South Korea, and Taiwan. For Hong Kong the null hypothesis is rejected at the 1 percent level for expected real stock returns and expected real economic activity and at the 5 percent level for expected inflation and inflation variability. We can conclude that all four variables in equations (1) through (4) are stationary processes for the four Pacific Basin countries, regardless of the frequency of data employed.

Tables 2 and 3 report monthly and quarterly OLS (ordinary least squares) estimates of the regression equations testing the proxy-effect hypothesis for the four Pacific Basin countries. The first equation for each country reports the estimates of the regression with expected inflation as the independent variable. The second equation reports the estimates with the expected economic activity variable added, and the third equation reports the estimates with inflation variability added. The fourth equation reports the estimates with both the expected economic activity and inflation variability variables included in the regression equation. Following McCarthy, Najand, and Seifert (1990), the dependent variable of each equation is expected real stock returns.

The OLS results obtained with monthly data show a negative relationship between expected inflation and expected real stock returns for both South Korea and Taiwan, although the posited relationship is not statistically significant for Taiwan. Inclusion of the expected economic activity variable does not alter the statistical relationship between expected stock returns and expected inflation rate for these two countries. Furthermore, the coefficient for the expected economic activity variable is not statistically significant for both countries and is negative for Taiwan. The coefficient for inflation variability is also not statistically significant for both countries and is negative only for Taiwan. Its inclusion in the regression equation does not alter the relationship between stock returns and expected inflation. Finally, the inclusion of both the expected economic activity and inflation variability variables does not affect the stock return-inflation relationship. Therefore, both the Fama and Kaul-Seyhun versions of the proxy-effect hypothesis are not valid for South Korea and Taiwan on the basis of monthly data.

The findings with quarterly data are not much different than those obtained with monthly data. Table 3 shows that the coefficient for expected inflation is negative and statistically significant for all four countries. This result is consistent with those of Kwok and Li (1993) and Kwok (1994) who also employ quarterly data. Adding either the expected economic activity variable and/or inflation variability variable does not diminish the statistical significance of expected inflation in the regression

Table 2—OLS (Ordinary Least Squares) Estimates of the Regression Equations Testing the Proxy-Effect Hypothesis for South Korea and Taiwan: Monthly Results

$ERSR_t = \alpha + \beta_1 EI_{t-1} + \beta_2 ERA_{t+1} + \beta_3 INFVAR_t + \varepsilon_t$

where:

$ERSR_t$ = Expected real stock return for period t;
EI_{t-1} = Expected inflation at period t-1 for period t;
ERA_{t+1} = Expected economic activity for period t+1; and
$INFVAR_t$ = Inflation variability at time t.

	α	β_1	β_2	β_3	R^2
Panel A: South Korea					
	.01556	-1.8531			.33
	(9.4272)*	(-10.0011)*			
	.01549	-1.8507	.0004		.34
	(9.2950)*	(-9.9589)*	(.3253)		
	.01465	-1.9831		32.1598	.33
	(7.6469)*	(-8.5693)*		(1.9380)	
	.01450	-1.9869	.0006	33.9159	.35
	(7.4553)*	(-8.5629)*	(.4371)	(1.9806)	
Panel B: Taiwan					
	.0037	-.5289			.01
	(1.1953)	(-1.6904)			
	.0039	-.6731	-.0001		.01
	(1.2515)	(-1.8626)	(-.9683)		
	.0118	-.4853		-85.4884	.03
	(1.6547)	(-1.6339)		(-1.2590)	
	.0118	-.6253	-.0001	-83.7369	.03
	(1.6556)	(-1.8013)	(-.9355)	(-1.2322)	

Note: Figures in the parentheses represent the values of t-statistic of the estimates
*Significant at the 5 percent level

equation for all four countries. The coefficient for expected economic activity is positive yet not significant for all countries. The coefficient for inflation variability is negative for all countries and is significant only for Singapore. Therefore, the quarterly results, as in the case of monthly results, do not support either Fama's or Kaul and Seyhun's version of the proxy-effect hypothesis.

This study finds that the proxy-effect hypothesis is not valid for all four Pacific Basin countries on the basis of both monthly and quarterly data. These results support the findings of Cochran and Defina (1993), Lee (1996), Liu, Hsueh, and Clayton (1993), and McCarthy, Najand, and Seifert (1990) which also reject the proxy-effect hypothesis for major industrialized countries using monthly, quarterly and annual data.

Conclusion

This study has examined whether the proxy-effect hypothesis can adequately explain the negative relationship between real stock returns and expected inflation

Table 3—OLS Estimate of the Regression Equations Testing the Proxy-Effect Hypothesis for Hong Kong, Singapore, South Korea, and Taiwan: Quarterly Results

$ERSR_t = \alpha + \beta_1 EI_{t-1} + \beta_2 ERA_{t+1} + \beta_3 INFVAR_t + \varepsilon_t$

where:

$ERSR_t$ = Expected real stock return for period t;
EI_{t-1} = Expected inflation at period t-1 for period t;
ERA_{t+1} = Expected economic activity for period t+1; and
$INFVAR_t$ = Inflation variability at time t.

	α	β_1	β_2	β_3	R^2
Panel A: Hong Kong					
	.1191	-4.0831			.18
	(4.2975)*	(-2.9324)*			
	.1187	-4.0649	.0002		.18
	(4.2060)*	(-2.8736)*	(.1631)		
	.1788	-3.9266		-1344.608	.23
	(3.6548)*	(-2.8531)*	(-1.4703)		
	.1793	-3.8882	.0002	-1374.437	.23
	(3.6187)*	(-2.7809)*	(.3151)	(-1.4773)	
Panel B: Singapore					
	.0116	-.0732			.13
	(1.4928)	(-2.1945)*			
	.0081	-.8283	.0005		.15
	(.8722)	(-2.6550)*	(1.2273)		
	0014	-.8629		-418.779	.21
	(.0707)	(-2.6591)*		(-2.5183)*	
	-.0038	-.8883	.0007	-196.011	.26
	(.1824)	(-2.6806)*	(1.1202)	(-2.2363)*	
Panel C: South Korea					
	.0111	-.6819			.11
	(1.5381)	(-2.6054)*			
	.0074	-.6467	.0001		.12
	(.9195)	(-2.4243)*	(.8062)		
	.0146	-.5849		-48.3288	.13
	(1.9401)	(-2.1597)*		(-1.2940)	
	.0116	-.5597	.0002	-45.6599	.14
	(1.3328)	(-2.0382)*	(.6788)	(-1.2102)	
Panel D: Taiwan					
	.0916	-4.2517			.23
	(3.7064)*	(-2.4342)*			
	.0888	-4.1052.0001			.24
	3.3766)*	(-2.2582)*	(.3663)		
	.0993	-4.2857		-47.9468	.24
	(1.5555)	(-2.3866)*		(-.1314)	
	.0851	-4.0785	-.0008	-21.4942	.26
	(1.1008)	(-2.1197)*	(-.3395)	(-.0508)	

Note: Figures in the parentheses represent the values of t-statistic of the estimates
*Significant at the 5 percent level

for four Pacific Basin countries: Hong Kong, Singapore, South Korea, and Taiwan. We find that the negative and statistically significant relationship between stock returns and inflation persists even after the effects of expected economic activity and inflation variability have been incorporated. Therefore, the findings of this study do not support the proxy-effect hypothesis for the four Pacific Basin countries. These results are also consistent with other recent studies that have rejected the proxy-effect hypothesis for major industrialized countries.

References

1. Akaike, H., "A New Look at Statistical Model Identification," *IEEE Tran. Automatic Control AC-19* (1974), pp. 716-722.

2. Bodie, Z., "Common Stocks as a Hedge Against Inflation," *Journal of Finance*, 31 (May 1976), pp. 459-470.

3. Bollerslev, T., "Generalized Autoregressive Conditional Heteroscedasticity," *Journal of Econometrics*, 31 (April 1986), pp. 307-327.

4. Cochran, S.J., and R.H. Defina, "Inflation's Negative Effects on Real Stock Prices: New Evidence and a Test of the Proxy-Effect Hypothesis," *Applied Economics*, 25 (January 1993), pp. 263-274.

5. Cukierman, A., "Relative Price Variability and Inflation: A Survey and Further Results," *Carnegie-Rochester Conference Series on Public Policy*, 19 (1983), pp. 103-157.

6. Dickey, D.A., and W.A. Fuller, "Distribution of the Estimators for Autoregressive Time Series with a Unit Root," *Journal of the American Statistical Association*, 74 (June 1979), pp. 427-431.

7. Dickey, D.A., "Likelihood Ratio Statistics for Autoregressive Time Series with a Unit Root," *Econometrica*, 49 (July 1981), pp. 1057-1072.

8. Engle, R.F., "Autoregressive Conditional Heteroskedasticity with Estimates of the Variance of U.K Inflation," *Econometrica*, 50 (July 1982), pp. 987-1008.

9. Fama, E.F., "Stock Returns, Real Activity, Inflation and Money," *American Economic Review*, 71 (September 1981), pp. 545-565.

10. Fama, E.F., and M. Gibbons, "A Comparison of Inflation Forecasts," *Journal of Monetary Economics*, 13 (May 1984), pp. 327-348.

11. Fama, E.F., and G.W. Schwert, "Asset Returns and Inflation," *Journal of Financial Economics*, 5 (November 1977), pp. 115-146.

12. Friedman, M., "Nobel Lecture: Inflation and Unemployment," *Journal of Political Economy*, 85 (June 1977), pp. 451-472.

13. Fuller, W.A., *Introduction to Statistical Time Series* (New York: Wiley, 1976).

14. Geske, R., and R. Roll, "The Fiscal and Monetary Linkage between Stock Returns and Inflation," *Journal of Finance*, 38 (March 1983), pp. 1-33.

15. Granger, C., and P. Newbold, "Spurious Regressions in Econometrics," *Journal of Econometrics*, 2 (July 1974), pp. 1045-1066.

16. Gultekin, N.B., "Stock Market Returns and Inflation: Evidence from Other Countries," *Journal of Finance*, 38 (March 1983), pp. 49-65.

17. Hafer, R.W., and S.E. Hein, "Forecasting Inflation Using Interest-Rate and Time-Series Models: Some International Evidence," *Journal of Business*, 63 (January 1990), pp. 1-17.

18. James, C., S. Koreisha, and M. Partch, "A VARMA Analysis of the Causal Relations among Stock Returns, Real Output, and Nominal Interest Rates," *Journal of Finance*, 40 (December 1985), pp. 1375-1384.

19. Kaul, G., "Stock Returns and Inflation: The Role of the Monetary Sector," *Journal of Financial Economics*, 18 (June 1987), pp. 253-276.

20. Kaul, G., "Monetary Regimes and the Relation between Stock Returns and Inflationary Expectations," *Journal of Financial and Quantative Analysis*, 25 (September 1990), pp. 307-321.

21. Kaul, G., and H.N. Seyhun, "Relative Price Variability, Real Shocks, and the Stock Market," *Journal of Finance*, 45 (June 1990), pp. 479-496.

22. Ko, K., and S. Lee, "A Comparative Analysis of The Daily Behavior of Stock Returns: Japan, The US and The Asian NICs," *Journal of Business Finance and Accounting*, 18 (January 1991), pp. 219-234.

23. Kwok, R., "Inter-relationship between Stock Returns and Real Economic Variables: The Experience of South Korea," *Research in International Business and Finance*, 11 (1994), pp. 217-226.

24. Kwok, R., and K. Li, "The Interrelationship between Macroeconomic Variables and Stock Prices in the Newly Industrialized Economies of Asia," *Research in International Business and Finance*, 10 (1993), pp. 221-228.

25. Lahiri, K., K. Teigland, and M. Zaporowski, "Interest Rates and the Subjective Probability Distribution of Inflation Forecasts," *Journal of Money, Credit, and Banking*, 20 (May 1988), pp. 233-248.

26. Lee, B.S., "Causal Relations among Stock Returns, Interest Rates, Real Activity, and Inflation," *Journal of Finance*, 47 (September 1992), pp. 1591-1603.

27. Lee, U., "Further Empirical Test of the Proxy-Effect Hypothesis: Some International Evidence," *Journal of International Financial Markets, Institutions and Money*, 6 (1996), pp. 35-46.

28. Liu, Y.A., L.P. Hsueh, and R.J. Clayton, "A Re-Examination of the Proxy-Effect Hypothesis," *Journal of Financial Research*, 16 (Fall 1993), pp. 261-268.

29. Malkiel, B., "The Capital Formation Problem in the United States," *Journal of Finance*, 34 (May 1979), pp. 291-306.

30. Martin, D., and B. Wood, Jr., "Market Microstructure and Index Returns: Evidence from The Pacific Rim," *Journal of International Financial Markets, Institutions and Money*, 6 (1996), pp. 45-58.

31. McCarthy, J., M. Najand, and B. Seifert, "Empirical Tests of the Proxy Hypothesis," *Financial Review*, 25 (May 1990), pp. 251-263.

32. Nelson, C.R., "Inflation and Rates of Return on Common Stocks," *Journal of Finance*, 31 (May 1976), pp. 471-483.

33. Solnik, B., "The Relation between Stock Prices and Inflationary Expectations: The International Evidence," *Journal of Finance*, 38 (March 1983), pp. 35-48.

34. Stulz, R.M., "Asset Pricing and Expected Inflation," *Journal of Finance*, 41 (March 1986), pp. 209-223.

35. Thornton, D.L., and D.S. Batten, "Lag-Length Selection and Tests of Granger Causality between Money and Income," *Journal of Money, Credit, and Banking*, 17 (May 1985), pp. 164-178.