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COINTEGRATION AND CAUSALITY BETWEEN MACROECONOMIC VARIABLES AND SHARE PRICES

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Abstract
The aim of this study is to investigate whether there is a relationship between some macroeconomic variables and share prices. In the analysis, covering the period of 1990: 01-2003: 12, variables of Istanbul Stock Exchange index, consumer price index, money supply, interest rate, exchange rate, trade balance, and industrial production index were used. Least squares estimation method, Johansen-Jeselius cointegration test, Granger causality test and variance decomposition results produced by VEC model were used in the study. These analysis shows that there is a long run relationship between some macroeconomic variables and share prices.

I. Introduction

In recent years, one of the most discussed issues in the economics and finance literature is the relationship between share prices and macroeconomic variables. During the 1980's and 1990's, unexpected share price volatility has occurred in European countries, Japan, and especially in the USA. Many researchers expressed that these volatilities may be caused by macroeconomic factors. Parallel to this, researchers started to investigate relationship between share prices and macroeconomic variables such as money supply, inflation, interest rate, industrial production, gross domestic product, trade balance, exchange rate, and petroleum prices. It has been observed that approaches and results about this subject have been changed. For example, some researchers say that these volatilities are caused by speculative movements rather than macroeconomic factors (Binswanger, 1999, 2004; Shiller, 2000). In spite of all these opinions, the relationship between share prices and macroeconomic variables has been studied since 1970's in economics and finance literature and the contradictory results made this subject contemporary.

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The most commonly used model in share valuation is present value or discounted cash flows model. In the model, share price (Pt) is the sum of present value of future dividends (Dt+i) and the expected final price of share (Pt+K) during K period. R expresses the expected yield rate (discount rate) which is the sum of risk free interest rate and risk premium.

$$P_t = E_t\left[\sum_{i=1}^{K}\left(\frac{1}{1+R_{t+1}}\right)^{i} D_{1+i}\right] + E_t\left[\left(\frac{1}{1+R_{t+K}}\right)^{K} P_{t+K}\right] \qquad (1)$$

Starting point of discussion about the relationship between share prices and macroeconomic factors has emerged from the effects of real economic variables on dividends and discount rate that have very important role in share valuation. In this subject, Leroy and Porter (1981), Shiller (1981) and Flannery and Protopapadakis (2002) stated that macroeconomic factors change the discount rate placed in discounted cash flow model and firms' ability to create cash.

This study consists of five sections. In the second section, a brief literature is presented. In the third section, we introduced the method and data used in the study, and the fourth section presents the findings of the applications. Results and evaluations are given in the fifth section.

II. Literature Review

In recent years, the studies focused on the relationship between macroeconomic variables and national stock markets have become one of the most important papers of economics and finance literature. Lots of macroeconomic variables are used in these studies but money supply, inflation rate, interest rate, exchange rate, trade balance and industrial production index are the most commonly used variables in the literature.

a) Relationship Between Money Supply and Share Prices

The first studies about relationship between money supply and share prices are prepared by Palmer (1970) and Spirinkel (1971). These researchers show that money supply changes affect the share prices. Then, Ho (1983), Smirlock and Yawitz (1985), Cook and Hahn (1988), Fung and Lie (1990), Malliaris and Urrutia (1991), Fosback (1991), Abdullah and Hayworth (1993), Lin (1993), Fitzpatrick (1994), Alexakis et al. (1996) and Thorbecke (1997) found similar results in their studies. The common view about this subject is that an increase in money supply causes an increase in share prices. Changes in money supply, first of all, affect the financial market via direct effects on general economy. If increase rate of money supply is high, market interest rates will decrease because of increase in the money amount of loanable fund.

Also, high increase rate in money supply can raise share prices by an increase in firms' operations and economic growth (Durukan, 1999).

On the other hand, the evidence that stock market leads money supply was found by Cooper (1974) and Rozeff (1974). More recently, this finding is supported by James et. al. (1985) and Thornton (1993).

Lee (1992), Rigobon and Sack (2001), Durham (2003) could not find a causality between money supply and share prices. But Kwon and Shin (1999) found a mutual causality between them for Korean economy.

In this study, as used in Abdullah and Hayworth (1993), Fitzpatrick (1994) and Alexakis et. al. (1996), M1 is used as a measure of money supply.

b) Relationship Between Inflation and Share Prices

Investor's level of protection against inflation is very important in stock markets. It is essential to investigate the effect of inflation on share prices especially in countries which have experienced high level of inflation as in Turkey.

The results of the previous studies are contradictory. The explanation for the negative correlation between stock returns and the level of inflation is fist introduced by Fama (1981), and later extended by Kaul (1987). This hypothesis, known as proxy hypothesis, states that the negative correlation is a result of the negative correlation between inflation and future output growth. Since stock prices reflect firms' future earnings potential, an economic downturn predicted by a rise in inflation will decrease stock prices and hence returns. Fama's argument has been supported by Lee's (1992) Granger causality test, by dynamic asset-pricing models of LeRoy (1984) and Marshall (1992).

Some researchers argue that the negative correlation between inflation and returns is a result of investors shifting from stocks to interest-bearing assets in inflationary periods. Modigliani and Cohn (1979) claim that investors mistakenly use nominal discount rate because of money illusion. In this subject, Mascaro and Meltzer (1993) argue that inflation uncertainty is positively correlated with the level of inflation, and an increase in inflation uncertainty raises demand for money, thereby reducing the demand for stocks. Feldstein(1980) and Feldstein and Summers (1979) state that a rise in unanticipated inflation decreases firm's equity values by reducing the real value of depreciation tax shield.

Opposite to these negative correlation arguments, Kessel (1956) claims that a rise in unanticipated inflation will increase the firm's equity values if the firm is a net debtor. Kessels's positive correlation argument has been supported by Fosback (1991), Abdullah and Hayworth (1993) and Graham (1996). These researchers supported Fisher's opinions. According to them, the observed negative relationship between inflation and share prices appears to contradict with Fisher's proposition that common stocks are hedges against inflation (Chopin and Zhong, 2000).

c) Relationship Between Interest Rate and Share Prices

Previous studies report evidence that there is a strong effect of interest rate on share prices. Interest rates are not only affective on share value but on stock demand by changing the value of bond which is an important alternative investment tool.

Interest rate changes can impact share prices through two channels. First, by affecting the rate at which the firm's cash flows will be capitalized, and second, by changing expectations of future cash flows. Central Bank's discount rates and target rate changes are often viewed by stock market participants as signals of the future direction of interest rates and inflation.

Cook and Hahn (1988) and Smirlock and Yawitz (1985) found evidence supporting the favorable signal sent by rate decreases and the unfavorable signal sent by rate increases as the short-term stock market reaction (announcement effect) is negative to rate increases and positive to rate decreases. These studies of short-term stock returns establish that the interest rate changes produce announcement effects in financial markets and the reaction to the rate changes is very rapid. However, Jensen and Johnson (1995) provide evidence suggesting that long-term stock market performance is correlated with changes in the Fed discount rate. The authors found that stock returns following discount rate decreases are higher and less volatile than returns following rate increases.

Saunders and Yourougou (1990) investigated the subject from the side of firm's assets and liabilities and show that differences in stock returns are partially explained by interest rate sensitivity of firm's assets and liabilities. In the study, they found that securities that claim on real assets (industrial firm stocks) are less sensitive to unexpected changes in nominal interest rates than those that are claims on monetary assets (financial intermediary stocks).

d) Relationship Between Exchange Rate and Share Prices

The exchange rate has often been used to analyze stock prices in the belief that corporate earnings are significantly affected by fluctuations in currency value.

The early researchers found no uniform pattern of reaction by the stock prices to exchange rate volatility. Its reason may be attributed to the fixed regime of the Bretton Wodds period when exchange rates hardly moved. However, the role of exchange rates in influencing domestic prices, including stock prices, has been heightened following the floating of major currencies after floating exchange rate regime in the beginning of 1973 and especially in recent years as the dollar exhibited unprecedented volatility in the 1980's. Also, the dramatic increases in the world trade and capital movements have made the currency value as one of the main determinants of corporate profitability and share prices. These developments helped considerable interests increase in the exchange rate-share price linkage (Kim, 2003).

There is no theoretical consensus on the relationship between stock prices and exchange rates. For example, while Aggarwal (1981) and Solnik

(1987) found a positive relationship between share prices and Exchange rate, Soenen and Hennigar (1988) found a negative relationship.

Aggarwal (1981) explored the relationship between the dollar exchange rates and changes in indices of stock prices. He used monthly U.S. stock price data and the effective exchange rate for the period of 1974-1978. The results, which were based on simple regressions, showed that share prices and the value of the U.S. dollar is positively related and this relationship is stronger in the short run than in the long run.

Solnik (1987) examined the impact of several variables such as exchange rates, interest rates and changes in inflationary expectation on share prices. In the study, he used monthly data from nine markets (U.S., Japan, Germany, U.K., France, Canada, Netherlands, Switzerland and Belgium). He concluded that exchange rates affect share prices positively in all countries except the U.S.

Soenen and Hennigar (1988) employed monthly data on share prices and effective exchange rates for the period of 1980-1986. In the study, it is concluded that there is a strong negative relationship between share prices changes and value of the U.S. dollar. However, when they analyzed the relationship for a different period, they found a statistically significant negative impact of revaluation on share prices.

Rittenberg (1993) examined the subject for Turkey and employed the Granger causality test to investigate the relationship between exchange rate changes and price level changes. Since the causality tests are sensitive to lag selection, he employed three different methods for optimal lag selection (an arbitrarily selected, Hsiao method and Kunst and Marin's SMAR model). In all cases, he found that causality runs from price level change to exchange rate changes.

e) Relationship Between Economic Activities and Share Prices

The results of the studies about the relationship between economic activities and share prices are controversial. Mahdavi and Sohrabian (1991) claimed that GNP, one of the most proper indications of economic activities, follows the trend of share prices. Abdullah and Hayworth (1993) stated that economic activities affect share prices due to its effect on firm earnings. Fama (1990), Shwert (1990), Abdullah and Hayworth (1993), Fitzpatrick (1994), Chatrath et. al. (1996) and Zhao (1999) used GNP as an indicator for economic activities in their studies. But Kwon and Shin (1999), Nasseh and Strauss (2000) and Binswanger (2004) used industrial production index as a proxy for economic activities. We used industrial production index in this study due to the absence of GNP monthly data.

f) Relationship Between Trade Balance and Share Prices

Although the relationship between trade balance and share prices is not much studied in the literature, this variable is included in the study because of openness of economy and its indirect relations with the other variables used. About this subject, Kwon and Shin (1999) especially stressed the bilateral causal relationship between trade balance and exchange rates for Korean economy.

III. Method and Data

In the study, the relationship between share prices and macroeconomic variables is investigated with both ordinary least squares and a Vector Error Correction model for mutual relationship between these variables. VEC model is introduced by Engle and Granger (1987) for examining cointegration i.e. for long-run relationship between variables in Vector Autoregression (VAR) model. A VEC model can be expressed as follow.

$$\Delta x_t = \alpha + B(L)\,\Delta x_{t-1} + d'(e_{t-1}) + \eta_t \qquad (2)$$

Where

Δx_t = n × t vectors of variables,
α = n ×1 vectors of constants,
B(L) = n × n matrices of the polynomial expression in the lag operator,
d' = n ×1 vectors of constants,
e_{t-1} = n ×t vectors of error correction terms,
η_t = n ×t vectors of residuals.

In the study, Granger causality test is applied for investigating direction of relationship between variables. This test is fist applied in the literature by Granger (1969). Then, Hamilton (1994) developed this test. In Granger causality, the direction of relationship is investigated between variables like x and y. If the current y value is better estimated with past values of x rather than current value of x, we can say that there is Granger causality from x to y variable (Cheremza and Deadman, 1993).

The study covers the period of 1990:01-2002:12 Turkish economic data. The monthly data are obtained from Turkish Republic Central Bank Electronic Data Delivery System. Eviews 3.0 was used for analysis. The symbols and names of variables are; ISE: Istanbul Stock Exchange Index, CPI: Consumer Price Index, IP: Industrial Production Index, M1: Money Supply, IR: Interest Rate, EX: Exchange Rate (expressed as USA dollar), and TB: Trade Balance.

IV. Findings

In the first step, stationarity of variables is investigated. Non-stationarity of series is probably possible when working with time series. It is possible to face spurious regression with non-stationary data. So, estimate results may give a spurious relationship between variables. In case that series are non-stationary at level values, they can be stabled by getting the difference between them. So, by eliminating spurious regression problem, it is possible to reach healthier results (MacKinnon, 1991).

For investigating stationarity of time series used in the study, Augmented Dickey-Fuller (ADF) unit root test is employed. In application of the unit root test, in the first step the test is performed with fixed and trend. In this stage, if stationarity is achieved, these values are used without passing to with fixed and no fixed process (Enders, 1995).

Table 1 shows Augmented Dickey-Fuller unit root test results. According to the results of the test, TB, IR, IP, and M1 are stationary at level values, EX, ISE and CPI are stationary at first difference values.

Table 1: ADF Unit Root Test *

Variables		Level	1st Difference
		Fixed and Trend	Fixed and Trend
EX		-1.797(1)	-7.792(1)[(a)]
TB		-3.441(2)[(b)]	-
IR		-4.700(2)[(a)]	-
ISE		-1.711(0)	-14.816(0)[(a)]
IP		-19.440(0)[(a)]	-
CPI		0.502(0)	-5.623(0)[(a)]
M1		-3.271(6)[(c)]	-
Critic Values	**a=%1**	-4.015	-4.015
	b=%5	-3.437	-3.437
	c=%10	-3.143	-3.143

* The numbers in parenthesis are the appropriate lag lengths and they are produced by using Akaike Information Criteria and Schwartz Bayesian Criteria.

OLS estimates results between ISE and selected macroeconomic variables are given in Table 2. According to the regression results, while the coefficients of EX, CPI, IR, TB, and M1 variables are found statistically significant, IP's coefficient is not. That is, while any variation in EX, CPI, IR, TB, and M1 variables affect ISE positively and significantly, any variation in IR and TB variables affect ISE negatively and significantly. IP variable doesn't affect the ISE variable.

Table 2: OLS Estimates Results between ISE and Selected Macroeconomic Variables

Variable Pairs	Estimate Result of Model	R^2	DW	F
ISE – EX	ISE = 981.2946 + 0.0079 EX t (3.4762) (19.8019)*	0.7026	0.1685	392.1158*
ISE – CPI	ISE = 1055.0170 + 0.0388 CPI t (3.8429) (20.2952)*	0.7128	0.1624	411.8955*
ISE – IR	ISE = 5914.373 – 25.3620 IR t (5.9147) (-1.7244)***	0.0176	0.0620	2.9736***
ISE – TB	ISE = 1920.410 – 2.5997 TB t (2.3388) (-3.3768)*	0.0643	0.0838	11.4029*
ISE – M1	ISE = 1968.760 + 0.00001 M1 t (4.3689) (8.2853)*	0.2926	0.5508	68.6468*
ISE - IP	ISE = 4339.391 + 3.0999 IP t (10.4533) (0.0639)	0.00002	0.0468	0.0041

Note: *,** and *** show significance levels of %1, %5 and %10 respectively.

After estimating OLS results, our next step is to investigate whether there is a long run relationship between variables by using Johansen-Juselius cointegration test. Because the variables used in the study are cointegrated in different degrees, this method is preferred to Engle-Granger two step cointegration test. Johansen-Juselius cointegration test is first performed for bilateral relationships with ISE, and then for multi-relationships.

Table 3: Johansen-Juselius Cointegration Tests of Two-Variable Model

Variable Pairs	Null Hypothesis	Alternative Hypothesis	Likelihood Ratio	%5 Critic Value	%1 Critic Value
EX(2)	r = o	r = 1	9.4728	15.41	20.04
ISE(1)	r ≤ 1	r = 2	1.2192	3.76	6.65
CPI(3)	r = o	r = 1	22.3108*	15.41	20.04
ISE(1)	r ≤ 1	r = 2	6.1511**	3.76	6.65
IR(2)	r = o	r = 1	24.0919*	15.41	20.04
ISE(1)	r ≤ 1	r = 2	0.0117	3.76	6.65
TB(2)	r = o	r = 1	13.6548	15.41	20.04
ISE(1)	r ≤ 1	r = 2	0.1591	3.76	6.65
M1(5)	r = o	r = 1	12.6518	15.41	20.04
ISE(1)	r ≤ 1	r = 2	0.0515	3.76	6.65
IP(4)	r = o	r = 1	75.9577*	15.41	20.04
ISE(1)	r ≤ 1	r = 2	0.2405	3.76	6.65

Notes: * and ** show significance levels of %1 and %5 respectively. The numbers in parenthesis are the appropriate lag lengths and they are produced by using Akaike Information Criteria and Schwartz Bayesian Criteria.

According to the two-variable cointegration test presented in Table 3, while there is a cointegration relationship between ISI and CPI, and IR and IP, a cointegration relationship can not be observed between ISE and EX, and TB and M1. So, we can say that there is a long run relationship between ISE and CPI, and IR and IP variables.

Table 4: Johansen-Juselius Multivariate Cointegration Test

Null Hypothesis	Alternative Hypothesis	Likelihood Ratio	%5 Critic Value	%1 Critic Value
r = 0	r = 1	258.1459*	94.25	103.18
r ≤ 1	r = 2	148.8039*	68.52	76.07
r ≤ 2	r = 3	72.8711*	47.21	54.46
r ≤ 3	r = 4	35.0843**	29.68	35.65
r ≤ 4	r = 5	12.6425	15.41	20.04
r ≤ 5	r = 6	1.5073	3.76	6.65

Note: * and ** show significance levels of %1 and %5 respectively.

Johansen-Juselius cointegration tests of multi-relationships are presented in Table 4. The result shows that there are four cointegration vectors between variables. It means that, there is a long run relationship between variables used in the study. So, at least one directional relationship must be expected between variables (Gujarati, 1995).

Table 5 presents Granger causality test results between ISE and six macroeconomic variables from VEC model. According to the test results, there is a bidirectional causality between ISE and EX and M1. That is, share prices anticipate changes in EX and M1 and vice versa. Also, we found causality from changes in CPI and IR to changes in ISE. This means inflation and interest rates anticipate changes in share prices and share prices appear to have no effect on inflation and interest rates. There is no causality between ISE and IP and TB variables.

Table 5: Granger Causality Test Results from VECM

Variables	Direction of Causality	F Statistics	Probability
ISE – EX(2)	→	19.2182*	0.0000
EX – ISE (1)	→	3.4397**	0.0346
ISE – CPI(3)	-	1.3514	0.2601
CPI – ISE (1)	→	5.4807*	0.0014
ISE – IP(4)	-	0.5722	0.6832
IP – ISE (1)	-	0.6674	0.6156
ISE – IR(2)	-	0.0986	0.9062
IR – ISE (1)	→	7.3731*	0.0009
ISE – M1(5)	→	3.0167*	0.0128
M1 – ISE (1)	→	3.9065*	0.0024
ISE – TB(2)	-	1.7558	0.1763
TB – ISE (1)	-	1.9564	0.1450

Notes: * and ** show significance levels of %1 and %5 respectively. The numbers in parenthesis are the appropriate lag lengths and they are produced by using Akaike Information Criteria and Schwartz Bayesian Criteria.

Variance decomposition results from VEC model are given in Table 6. Variance decomposition shows how much percentage of the total variance is explained by each component. According to the variance decomposition results, ISE variable is mostly affected by its shocks, and then by variables of interest rate, consumer price index, trade balance, money supply, exchange rate and industrial production index, respectively. At the beginning of the period, share prices are determined only by their shocks but at the end of the 24^{th} month, this ratio decreases to 75%. So, 25% percentage of change in ISE was determined by macroeconomic variables used in the study. Interest rate shows its impact after the fourth month and this ratio rises to 13% at the end of the period. Inflation affects ISE by 6% at the end of the fourth month and this ratio rises to 9% at the end of the 24^{th} month. Trade balance increasingly affects the share prices but the figure doesn't exceed the level of 2%. Money supply can explain only 1% of share price changes. The effect of EX and IP is rather small.

Table 6: Variance Decomposition of ISE from VECM (%)

Period	ISE	IR	EX	M1	CPI	TB	IP
1	100.000	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000
4	90.1983	2.2793	0.6982	0.6603	5.8919	0.0157	0.2563
8	81.5128	9.0281	0.3702	1.0891	7.5347	0.1562	0.1527
12	77.7276	11.7772	0.2540	1.0366	7.9390	1.1462	0.1195
16	76.3154	12.5199	0.2058	0.9623	8.1281	1.7748	0.0938
20	75.7239	12.6215	0.1859	0.9132	8.3278	2.1514	0.0764
24	75.4085	12.5174	0.1836	0.8818	8.5358	2.4086	0.0644

V. Conclusion

The aim of this study is to investigate the relationship between share prices and some macroeconomic variables. For this aim, OLS method, Johansen-Juselius cointegration method, Granger causality test and VEC model are used. OLS estimation results show that there is a significant relationship between share prices and consumer price index, exchange rate, interest rate, money supply and trade balance, but not with industrial production index.

By using Johansen-Juselius cointegration test, the long run relationship between share prices and macroeconomic variables is investigated for two variables and multivariable. The result of this test shows that, there is a long run relationship between share prices and consumer price index, interest rate and industrial production index. In the multivariate Johansen-Juselius cointegration test, four cointegration vectors are found. This result shows that there may be a long run relationship between variables, even with one direction.

Causality relationships between variable pairs are examined with Granger causality test. According to the Granger causality test results, there is a mutual relationship between share prices and money supply and exchange rate. Also, we found that there is one directional Granger causality relationship between share prices and variables of exchange rate and interest rate and these two variables lead the share prices. A relationship can not be found between share prices and the other macroeconomic variables of trade balance and industrial production index.

According to the share price variance decomposition results produced by VEC model, share prices are mostly affected by their own shocks, and then by variables of interest rate, consumer price index, trade balance, money supply, exchange rate, and industrial production index, respectively.

As a result, OLS estimates, Johansen-Juselius cointegration test, Granger causality test and variance decomposition results produced by VEC model show that even in various degrees, there is a relationship between share prices and macroeconomic variables for Turkish economy in the sample period.

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COMPARING THE CONVERGENCE BEHAVIOUR OF BINOMIAL AND TRINOMIAL MODELS

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Abstract
An American option differs from a European one by the early exercise possibility. An American option can be exercised at any time up to the maturity date. In general, there is unfortunately no analytical solution to the American option problem. Binomial and trinomial approximations are useful to solve this problem but using a lattice model introduces approximation error. Both models have the property of convergence to Black & Scholes prices thus; can be used alternatively to solve the Black & Scholes partial differential equation. This paper presents the suitable conditions under which the two models converge to Black & Scholes and compares the convergence behaviour in order to question their efficiency.

I. Introduction

Since Fischer Black and Myron Scholes published a paper in 1973 on the pricing of options and corporate liabilities in the Journal of Political Economy (Black and Scholes, 1973) and Robert Merton published a paper on the theory of option pricing in the Bell Journal of Economics and Management Science (Merton, 1973) the world of financial derivatives have grown up so fast. In contrast with the world of financial derivatives, the financial world has grown enormously both in dollar size and sophistication. The size of the contracts has grown with the same speed as the educational background of traders and the mathematics used in financial modelling (Wilmott, 1998). An advanced mathematical understanding is required as well as finance and economics background.

This paper presents the suitable conditions under which the two models converge to Black & Scholes and compares the convergence behaviour in order to question their efficiency. The next section gives insight about deriving the Black & Scholes partial differential equation and the difference on solving the partial differential equation analytically or numerically. Section three and four describes two numerical approaches such as binomial and trinomial model to solve the partial differential equation in order to calculate the fair price of the derivative. Section five compares two numerical approaches on their efficiency with focus on the deviation from the theoretical

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price. The results are also presented and discussed in section five and finally, the section six concludes the paper using the two models on preventing the arbitrage opportunities and gives directions for future research.

II. Black and Scholes Model

The first assumption Black & Scholes made to obtain the Black and Scholes partial differential equation is about the dynamics of the stock price. The stock price follows a log-normal process where W_t is a one dimensional Brownian motion. ($\mu \in \Re$ and $\sigma \in \Re+$ are constant deterministic numbers)

$$dS_t = \mu S_t dt + \sigma S_t dW_t \tag{1}$$

To understand the Black and Scholes model not only a very deep mathematical background, but also some basic insights is required. Clearly the value of a call option depends on the value of the underlying asset and time. ($C_t=F(S_t,t)$) Taking the derivatives by using the Ito's lemma, one can easily obtain the stochastic differential equation given below.

$$dC_t = \frac{\partial F}{\partial t}dt + \frac{\partial F}{\partial S_t}dS_t + \frac{1}{2}\frac{\partial^2 F}{\partial S_t^2}(dS_t)^2 \tag{2}$$

Replacing the term (1) in the stochastic differential equation (2) and rearranging terms the stochastic differential equation transforms into a simpler form.

$$dC_t = \frac{\partial F}{\partial t}dt + \frac{\partial F}{\partial S_t}(\mu S_t dt + \sigma S_t dW_t) + \frac{1}{2}\frac{\partial^2 F}{\partial S_t^2}(\mu S_t dt + \sigma S_t dW_t)^2$$

$$= \frac{\partial F}{\partial t}dt + \frac{\partial F}{\partial S_t}(\mu S_t dt + \sigma S_t dW_t) + \frac{1}{2}\frac{\partial^2 F}{\partial S_t^2}\sigma^2 S_t^2 dt$$

$$= \left(\frac{\partial F}{\partial t} + \mu S_t \frac{\partial F}{\partial S_t} + \frac{1}{2}\frac{\partial^2 F}{\partial S_t^2}\sigma^2 S_t^2\right)dt + \sigma S_t dW_t \tag{3}$$

The second important insight Black & Scholes put into action is to construct a special risk free portfolio. Since there is another risk free instrument in the economy, namely the money market account, both instruments should provide the same rate of return. ($\mu \equiv r$) The return of the money market account can be determined by an ordinary differential equation.

$$dF = rFdt \tag{4}$$

By using the no arbitrage arguments the equality of the return of the instruments should be maintained. This can be put into action by equalling the trend functions, the coefficient of the dt terms basically.

$$\begin{aligned} &\frac{\partial F}{\partial t} + rS_t \frac{\partial F}{\partial S_t} + \frac{1}{2}\frac{\partial^2 F}{\partial S_t^2}\sigma^2 S_t^2 = rF \\ &\frac{\partial F}{\partial t} + rS_t \frac{\partial F}{\partial S_t} + \frac{1}{2}\frac{\partial^2 F}{\partial S_t^2}\sigma^2 S_t^2 - rF = 0 \end{aligned} \tag{5}$$

This system is called the Black & Scholes Partial Differential Equation. Partial term comes from the partial derivatives of C with respect to different variables enter. It has many solutions, corresponding to all the different derivatives that can be defined with S as the underlying variable (Hull, 2003). Solving this system for function C is called the Cauchy problem. Mathematically, it is of the same type as the heat or diffusion equation, one of the most widely studied of partial differential equations (Wilmott, 1998). Option prices can be obtained as solutions of corresponding Cauchy problems under certain assumptions. If the corresponding Cauchy problem does not admit an explicit analytical solution then there remains the possibility of solving the Cauchy problem numerically (Korn and Korn, 2000). The Black & Scholes formula given below establishes the function C_t that solves partial differential equation for a European call option on a non-dividend paying stock where N(.) is the cumulative standard normal distribution function.

$$C(S_t, t) = S_t N(d_1) - Ke^{-r(T-t)}N(d_2) \tag{6}$$

where

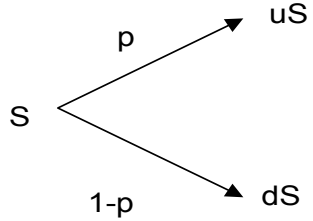
$$d_{1,2} = \frac{\ln(\frac{S_t}{K}) + (r \pm \frac{1}{2}\sigma^2)(T-t)}{\sigma\sqrt{T-t}}$$

This formula provides determining the value of a call option in a continuous-time framework. The option is said to be in-the money when the asset price is greater than the strike price for call options (S>K) and out of money when the asset price takes lower values than the strike price (S<K) at maturity. But regarding the complex mathematical techniques based on and the difficulties in implementing the model, a simpler approach was adopted (Cox et al, 1979) called the binomial model.

III. Binomial Model for Valuing Options

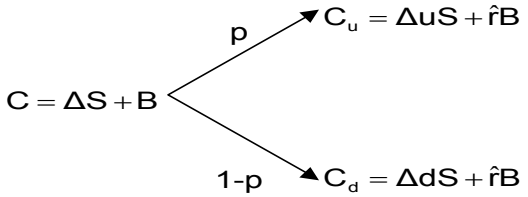
Cox, Ross and Rubinstein model relies on numerical computation instead of solving the partial differential equation analytically. The procedure followed by binomial model is to assume that, the stock price follows a discrete time process. The life of the option T-t is decomposed into n equal time steps of length (Δt=T-t/n). At each time interval (t_j=j.Δt , j=0,1,…,n), it is assumed that the rate of return on a stock over each period can have two possible values, u with probability p and d with probability 1-p. Thus, if the current stock price is S, it will be either uS or dS at the end of the period.

Figure 1: Binomial Sock Price Tree in One Period Setting



The condition $d < \hat{r} < u$ is required, where $\hat{r}$, is one plus the risk free interest rate over that period. This inequality ensures that neither the stock nor the safe asset is dominated by the other. The same arguments with the Black & Scholes model, hedging portfolio and no arbitrage restrictions, have to be considered to obtain the value of the option. Constructing a special portfolio containing Δ shares of stocks and investing the amount B in the risk free asset one can easily obtain the replicating portfolio. At the end of the period the portfolio will become as given below.

Figure 2: An Illustration of the Replicating Portfolio



Considering no arbitrage restrictions the values of Δ and B should be determined appropriately. To preclude arbitrage, the value of the option must be replicated by the given portfolio. This requires the values of Δ and B to be given as follows.

$$\Delta = \frac{C_u - C_d}{S(u-d)} \qquad B = \frac{1}{\hat{r}}\left(\frac{uC_d - dC_u}{u-d}\right) \tag{7}$$

This particular portfolio is called the hedging portfolio. If there are no risk free arbitrage opportunities, the current value of the call, C, is equal to the current value of the hedging portfolio, ΔS+B.

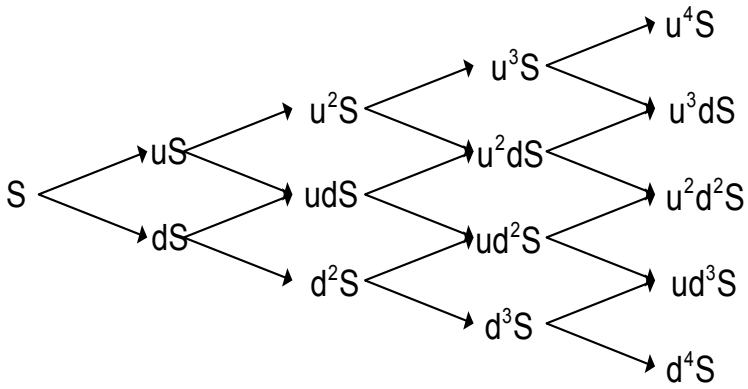
$$C = \frac{1}{\hat{r}}\left[\left(\frac{\hat{r}-d}{u-d}\right)C_u + \left(\frac{u-\hat{r}}{u-d}\right)C_d\right] \tag{8}$$

This formulation mimics the calculation technique if the world is risk-neutral. The terms in the brackets sum up to one and each of them are strictly positive. The value of the call does not depend on investors' attitudes toward risk. Binomial model assumes that an individual prefers more wealth to less and therefore has an incentive to take the advantage of profitable risk free arbitrage opportunities and the same formula whether the investors are risk-averse, risk-neutral or risk-preferring is obtained. The value of the call can be interpreted as the expectations of its discounted future value in a risk-neutral world where the risk-neural probability is, $q = (\hat{r} - d)/(u-d)$. The equation can be transformed as follows.

$$C = \frac{1}{\hat{r}}\left[qC_u + (1-q)C_d\right] \tag{9}$$

In order to approximate Black & Scholes differential equation by means of the Cox, Ross and Rubinstein approach, the probabilities p and q, as well as the rates u and d have to be chosen such that $\Delta t \to 0$. Considering this, the binomial model converges to a geometric Brownian motion. This convergence can be shown by using the binomial trees by increasing the number of steps in a given time period. Increasing the number of steps provides Δt to take closer values to zero. The argument can be extended to a multi-period setting.

Figure 3: An Illustration of Binomial Tree in a Multi-Period Setting



At the end of the each period an identical problem, solved in the previous period, is faced. Therefore, with a similar hedging argument and by using the risk-neutral probabilities, the value of the option can be summarized as follows.

$$C_u = \frac{1}{\hat{r}}(qC_{uu} + (1-q)C_{ud}) \tag{10}$$

$$C_d = \frac{1}{\hat{r}}(qC_{ud} + (1-q)C_{dd}) \tag{11}$$

This yields to express the current value of the option in terms of C_u and C_d. Therefore, working recursively backwards, one can easily obtain the value of the option for two steps.

$$C_u = \frac{1}{\hat{r}}(qC_u + (1-q)C_d) = \frac{1}{\hat{r}^2}(q^2C_{uu} + 2q(1-q)C_{ud} + (1-q)^2C_{dd}) \tag{12}$$

A similar recursive procedure for finding the value of a call with any number of periods remaining until expiration can be implemented for n number of steps. Starting from the expiration date and working backwards, the value of the option can be generalized as follows.

$$C = \frac{1}{\hat{r}^n}\left(\sum_{j=0}^{n}\frac{n!}{j!(n-j)!}q^j(1-q)^{n-j}\max(0, u^j d^{n-j}S - K)\right) \tag{13}$$

To achieve the convergence to geometric Brownian motion, one can increase the number of steps evaluated in a fixed time period by setting $n\to\infty$. Regarding the Cox, Ross and Rubinstein approach, the value of the call is adjusted to be in the money. Letting the parameter a to be the minimum number of upwards moves which the stock must take over the n periods for the call to finish in the money and furthermore denoting $q' \equiv (u/\hat{r})q$, the equation (13) can be transformed as given below.

$$C = S\left[\sum_{j=a}^{n} \frac{n!}{j!(n-j)!} q'^{j}(1-q')^{n-j}\right] - \frac{K}{\hat{r}^n}\left[\sum_{j=a}^{n} \frac{n!}{j!(n-j)!} q^{j}(1-q)^{n-j}\right] \quad (14)$$

Where

for all $j < a$, $\max(0, u^j d^{n-j} S - K) = 0$,

for all $j \geq a$, $\max(0, u^j d^{n-j} S - K) = u^j d^{n-j} S - K$

This transformation gives the binomial option pricing formula. The terms in the brackets are the complementary binomial distribution functions with probabilities q,1-q and the number of trials n, evaluated at a.

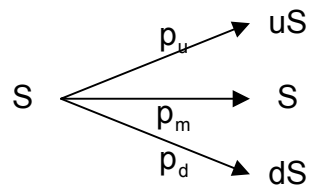
$$C = S\ \Phi(a;n;q') - \frac{K}{\hat{r}^n}\ \Phi(a;n;q) \quad (15)$$

The formula obtained by (15) seems to be interesting in sense of similarity with the formulae obtained by (6). Due to the increase in the number of steps, the binomial distribution changes to the normal distribution while other parameters remain the same.

IV. The Trinomial Model for Valuing Options

The binomial model is very intuitive but it has little flexibility to deal with more complex option problems (Figlewski and Gao, 1999). Trinomial trees can be used to model the stock price changes and valuing the options numerically as well as the binomial model. The trinomial model (Boyle, 1986) attempts to model the stock price movements better than the binomial method since the stock prices at each time point can change to three instead of two directions. It is assumed that the rate of return on a stock over each period can have one of three possible values, u with probability p_u, d with probability p_d and stable with probability p_m (p_m=1-p_u-p_d). Thus, if the current stock price is S, it will be uS, S or dS at the end of the period.

Figure 4: Trinomial Stock Price Tree in One Period Setting



As the name suggests, trinomial model uses a similar approach to the binomial one. But the hedging and replication arguments do not take place in constructing the trinomial trees. For a non-dividend paying stock, parameter values that match the mean and standard deviation of price changes are given below (Hull, 2003).

$$u = e^{\sigma\sqrt{3\Delta t}} \quad , \quad d = 1/u \quad (16)$$

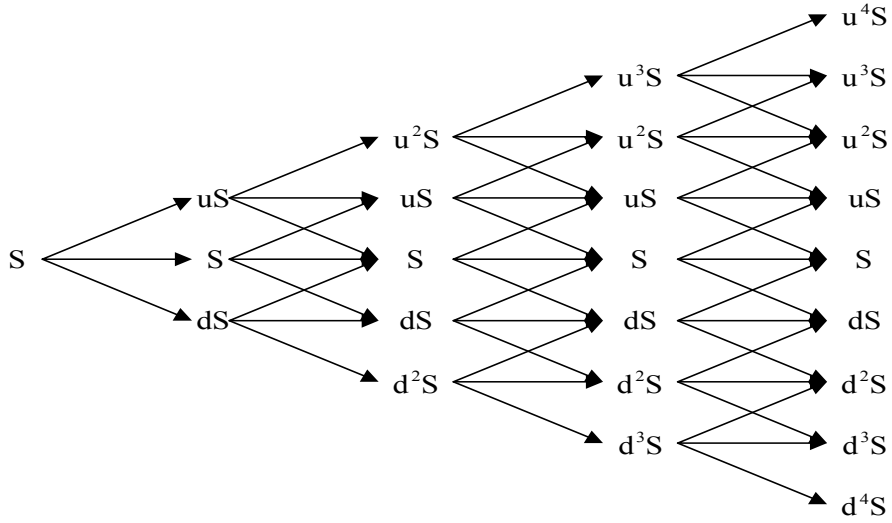
$$p_u = \sqrt{\frac{\Delta t}{12\sigma^2}}\left(r - \frac{1}{2}\sigma^2\right) + \frac{1}{6} \quad (17)$$

$$p_d = -\sqrt{\frac{\Delta t}{12\sigma^2}}\left(r - \frac{1}{2}\sigma^2\right) + \frac{1}{6} \quad (18)$$

$$p_m = 1 - p_u - p_d \quad (19)$$

Given the equations (16)-(19) and the backwards calculation approach used in Binomial model, one can easily obtain the price of an option by using trinomial model. An illustration of the trinomial stock price tree in multi-period setting is given below.

Figure 5: Trinomial Stock Price Tree in Multi-Period Setting



For many applications, the rate of converge is enhanced, if the tree is symmetrical (Figlewski and Gao, 1999). The convergence behaviour of binomial and trinomial models can be compared regarding the number of steps taken to reach an exact value or the computer time spent. Another interesting property of the trinomial model is that it can be shown to be equivalent to the explicit finite difference model (Hull and White, 1990).

V. Comparing the Convergence Behaviour of Binomial and Trinomial Model

In most analysis, finer time steps are used in a lattice to reduce the approximation error. Unfortunately, using finer steps increases the number of nodes dramatically. So, in many applications the number of calculations or the computer time spent has to be taken into consideration.

First of all, the number of nodes at a certain number of steps for each model can be compared. Binomial model has n+1 nodes at time step n, or $[(N+1)^2+ (N+1)]/2 = (N^2+3N+2)/2$ nodes in total. A trinomial tree has 2n+1 nodes at time step n and $(N+1)^2$ in total. Thus the computation time increases sharply as greater accuracy is required (Figlewski and Gao, 1999).

Table 1: Comparing the Number of Nodes for Various Number of Steps

Number of steps	Binomial Model		Trinomial Model	
	Nodes at each node	Total number of nodes	Nodes at each node	Total number of nodes
2	3	6	5	9
4	5	15	9	25
8	9	45	17	81
16	17	153	33	289
64	65	2145	129	4225
128	129	8385	257	16641
250	251	31626	501	63001
500	501	125751	1001	251001
999	1000	500500	1999	1000000
5000	5001	12507501	10001	25010001
9999	10000	50005000	19999	100000000

Considering the outputs given in Table 1, trinomial model brings about almost twice more nodes than Binomial model. But this claim is not sufficient to compare the efficiency of these two models. As pointed out earlier, there is a relationship between the binomial model and the Black & Scholes option pricing model. With a significant number of time nodes, the binomial model begins to converge and the convergence of this value ultimately becomes the value determined by using Black & Scholes formula.

An illustrative example can be used to compare the convergence speed of these two models. The parameters of the illustrative example taken into consideration are as follows. The Black & Scholes price can be calculated easily using the given parameters.

Table 2: The Value of an American Call option on a Non-dividend Paying Stock

Parameters	Values
Stock Price	230
Strike Price	210
Volatility	0.25
Interest Rate	0.04545
Time to Maturity	0.5

The Black & Scholes value for the derivative is determined as 30.7416 analytically. This value can be compared with the ones obtained by using the binomial and trinomial models. For example, at the 500th time node the value of the call is calculated as 30.74494286 with binomial model whereas 30.74392118 with the trinomial one. Compared with the Black & Scholes price, trinomial model seems to be more effective. These values are calculated on a spreadsheet by implementing a VBA code. The VBA code implemented for binomial and trinomial model are given in the appendix.

Table 3: Comparing the errors of Binomial and Trinomial model

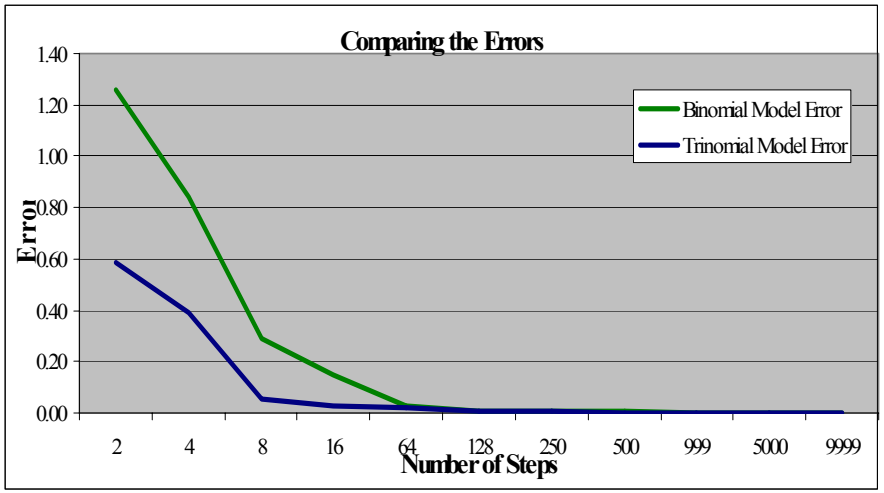
Number of Steps	Black& Scholes Price	Binomial Model Price	Trinomial Model Price	Binomial Model Error	Trinomial Model Error
2	30.74157465	32.00038517	31.32588806	1.25881051	0.58431341
4	30.74157465	31.58549904	31.13110699	0.84392439	0.38953233
8	30.74157465	31.03307427	30.68855260	0.29149962	0.05302205
16	30.74157465	30.59368251	30.71302621	0.14789214	0.02854844
64	30.74157465	30.71518215	30.76039417	0.02639251	0.01881952
128	30.74157465	30.73463131	30.75048269	0.00694335	0.00890803
250	30.74157465	30.73581799	30.74495486	0.00575666	0.00338021
500	30.74157465	30.74494286	30.74392118	0.00336821	0.00234653
999	30.74157465	30.74439400	30.74287974	0.00281935	0.00130509
5000	30.74157465	30.74175388	30.74098428	0.00017923	0.00059037
9999	30.74157465	30.74171491	30.74164703	0.00014026	0.00007238

Table 3 compares the performance of the binomial and trinomial models in computing the theoretical price. The theoretical price is calculated by using the Black & Scholes option pricing formula for a non-dividend paying stock and increasing the number of periods the value obtained by the two different lattice models are calculated. For each number of steps, binomial and trinomial error is determined by subtracting the theoretical value and the value obtained by the lattice one, respectively. Lattices until 9999 number of steps are examined and this value can be considered as sufficient regarding the size of the errors obtained.

This is quite certain that the size of the approximation errors binomial model produced are almost twice bigger than the trinomial ones which make us think that the trinomial model is more efficient. The size of the error diminishes to 0.00007238 which is considerably a small number for 9999 time steps. So using 9999 steps can be sufficient in order to obtain close values to theoretical price.

In order to compare the number of steps had to be taken to reach a certain value can be taken into consideration. The trinomial model produces the error 0.00234653 at 500th number of steps while this size of error is obtained at 999th step by the binomial one. So approximating the theoretical price at this level, one must make twice more calculations. Figure 6 illustrates the results, explained so far. The figure asserts that at any number of steps, binomial model produces bigger errors compare with the trinomial one.

Figure 6: The Relation Between the Number of Steps and the Errors

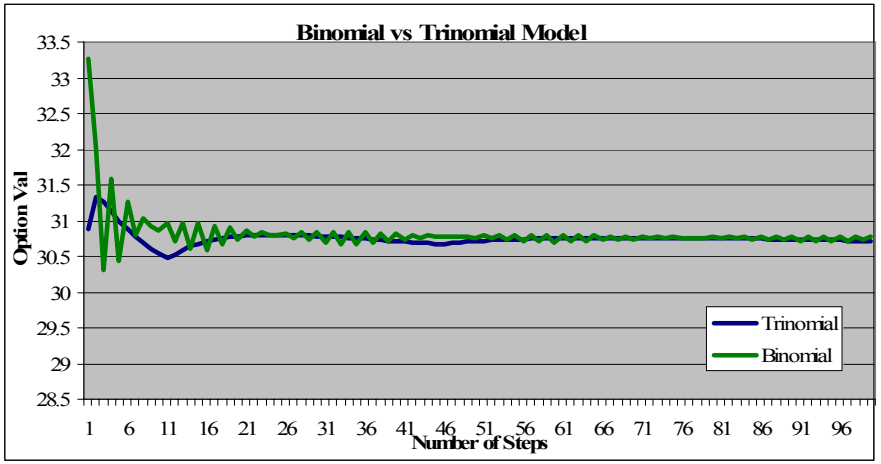


In Figure 6, the size of the errors produced by the two different models is compared in order to rank the efficiency of the lattice models. One can do the same analysis by calculating the option values obtained approximately at each number of steps and illustrating in a figure. Transposing the two lines together it is clear that the trinomial model converges much more rapidly than the binomial model.

Also the trinomial model obtains much closer values at smaller number of steps than binomial model. The size of the errors is dramatically bigger at the very beginning nodes for binomial model. Increasing the number of steps, closer evaluations can be achieved but this may be regarded as the most important shortcoming of the binomial model.

Basically, both models can be used to obtain the approximate values of the theoretical prices at higher nodes but, it is not appropriate to rely on the values obtained by the binomial model for considerably small number of steps. Figure 7 shows the option value obtained by the two lattice models as the number of steps increases.

Figure 7: Comparing the Option Values Calculated as Number of Steps Increases



VI. Conclusion

Although the Black & Scholes option pricing model provides a closed form solution to valuate the European options on a non-dividend paying stock, the model reaches an impasse with American options. Since, no closed form solution is derived, lattice based models as binomial model and trinomial model can be used in order to obtain approximate solutions.

In this paper in order to compare models efficiency, the theoretical price of a European call option is calculated with Black & Scholes option pricing formula and compared with the approximate values obtained with the lattice based models. To compare the efficiency of the models, the number of nodes at a certain number of steps for each model is compared and it is cited that trinomial model brings about almost twice more nodes than binomial model.

Number of nodes may effect the computer time spent to approximate the solution. For 5000 steps the binomial model executes the code in 37 seconds whereas the trinomial model executes in 61 seconds. Increasing the number of steps to 9999 affects the execution time to 161 seconds for the binomial model whereas 225 seconds for the trinomial model. Since, the trinomial model brings about almost twice more nodes than binomial model, the computer time spent to execute the source code for trinomial tree is longer than the binomial one. This conflicts with the size of the errors.

The second comparative study done was to compare the size of the errors, determined by subtracting the theoretical value and the value obtained by the lattice based models at each step. It is asserted that that at any number of steps, binomial model produces bigger errors compare with the trinomial one. The result of approximating the theoretical price at a certain level, one must make twice more calculations with binomial model is obtained which is a shortcoming for efficiency.

On an illustrative example it is shown that the trinomial model converges much faster to the theoretical price obtained by the Black & Scholes formula than the binomial method using the same input.

The price of the derivative obtained by using equation (6) can be regarded as the fair price of the contingent claim. Thus, trading on this value prevents all the arbitrage opportunities. Since, Black & Scholes option pricing model is established on the equilibrium on the market (no arbitrage opportunities are available) all investors have to be satisfied with bounded portfolios.

The notion of prices that do not allow for arbitrage is closely linked to the existence of state prices attached to these assets. If this condition is not satisfied the market is said to be incomplete. Therefore, to achieve the completeness of the market or no arbitrage opportunities in other words, every asset has to be priced truly.

When Black & Scholes option pricing formula is impossible to use due to the specific conditions of the contract, numerical techniques enables us to price contingent claims. Since, numerical techniques introduce approximation error; the model which allows investors to reach better approximation is robust. Therefore, comparison of the models is made focusing on achieving closer values to the Black & Scholes theoretical price.

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Appendix A: The VBA code for Binomial Model

```
Function Price(Asset As Double, Volatility As Double, IntRate As Double,
Strike As Double,   Expiry As Double, NoSteps As Double)

ReDim S(0 To NoSteps)
ReDim V(0 To NoSteps)

timestep = Expiry / NoSteps
DiscountFactor = Exp(-IntRate * timestep)

temp1 = Exp((IntRate + Volatility * Volatility) * timestep)
temp2 = 0.5 * (DiscountFactor + temp1)

u = temp2 + Sqr(temp2 * temp2 - 1)
d = 1 / u
p = (Exp(IntRate * timestep) - d) / (u - d)

S(0) = Asset

For n = 1 To NoSteps
   For j = n To 1 Step -1
      S(j) = u * S(j - 1)
   Next j
      S(0) = d * S(0)
Next n

For j = 0 To NoSteps
   V(j) = Payoff(S(j), Strike)
Next j

For n = NoSteps To 1 Step -1
   For j = 0 To n - 1

   V(j) = (p * V(j + 1) + (1 - p) * V(j)) * DiscountFactor

   Next j
Next n
Price = V(0)
End Function
Function Payoff(S, K)
Payoff = 0
If S > K Then Payoff = S - K
End Function
```

Appendix B: The VBA code for Trinomial Model

```
Function TPrice(Asset As Double, Volatility As Double, IntRate As Double,
Strike As Double, Expiry As Double, NoSteps As Double)

ReDim S(0 To NoSteps * 2) As Double
ReDim V(0 To NoSteps * 2) As Double

timestep = Expiry / NoSteps

DiscountFactor = Exp(-IntRate * timestep)

u = Exp(Volatility * (3 * timestep) ^ 0.5)
d = 1 / u

p_down = -1 * ((timestep / (12 * Volatility * Volatility))) ^ 0.5 * (IntRate -
0.5 * Volatility * Volatility) + 1 / 6
p_up = ((timestep / (12 * Volatility * Volatility))) ^ 0.5 * (IntRate - 0.5 *
Volatility * Volatility) + 1 / 6
p_middle = 2 / 3

For n = 0 To NoSteps * 2
        S(n) = Asset * u ^ (n - NoSteps)
Next n

For j = 0 To NoSteps * 2
        V(j) = Payoff(S(j), Strike)
Next j

For n = NoSteps - 1 To 0 Step -1
        For j = 0 To 2 * n

        V(j) = (p_down * V(j) + p_middle * V(j + 1) + p_up * V(j + 2)) *
DiscountFactor
        Next j
Next n

TPrice = V(0)

End Function
Function Payoff(S, K)
        Payoff = 0
        If S > K Then Payoff = S - K
End Function
```

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FOREIGN DIRECT INVESTMENT AND REAL EXCHANGE RATE: A CAUSALITY ANALYSIS

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Abstract
This paper empirically investigates the direction of a causal relationship between exchange rates and foreign direct investment (FDI) flows using quarterly data from Turkey for the period 1987-2000, by means of Granger non-causality testing procedure developed by Toda and Yamamoto (1995). The results indicate that Granger causality is unidirectional and is running from FDI to real effective exchange rates. The results of this article generally agree with the predictions by the portfolio model according to which financial and capital liberalization in Turkey leads to an increase in capital inflows, which in turn, a real exchange rate appreciation.

I. Introduction

Policy makers and researchers have been discussing the impact of exchange rate movements on trade and investment since the breakdown of the Bretton-Woods agreement. Motivated by growing trend toward floating exchange rates, one of the debated issues in the recent past has been on the nature of the causal relationship between exchange rates and foreign direct investment (FDI). There is, however, no consensus about the direction of such a relationship[1]. The determination of a causality relationship between exchange rates and FDI which can be defined as "an expansion of production of a corporation in some other host nation by establishing a subsidiary or purchasing a production plant abroad" (Seyidoğlu 2003) is important especially for developing countries because the determination of a causal pattern between exchange rates and FDI flows has important implications for policy makers' decisions about the appropriate growth and foreign trade policies to implement.

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[1] For a fuller discussion, see, Kosteletou and Liargovas (2000) and Lafrance and Tessier (2000).

Theoretically, the direction of a causal relationship between FDI and exchange rates is not clear. On the one hand, trade integrated model by Dornbusch (1973) and a portfolio model by Branson (1973) imply that the direction of causality runs from FDI to real exchange rates. On the other hand, some theoretical models such as imperfect capital markets model by Froot and Stein (1991) and the models by Cushman (1985) which stress the importance of input costs in analyzing the effect of the real exchange rate on FDI flows imply that causation runs from real exchange rate to FDI. In addition, most of the empirical literature examining the determinants of FDI implies that FDI inflows are driven by exchange rates (for example, Benassy-Quere, A., et.al, 2001 and Goldberg and Klein, 1997).

Few studies have dealt with the direction of causality between exchange rates and FDI. Using simultaneous equation model for 12 countries in Europe, the US and Japan, Kosteletou and Liargovas (2000) found that while causality runs from the real exchange rates to FDI in large countries with freely floating currencies, causality runs both ways in small countries whose currencies are directly connected with Euro. Lafrance and Tessier (2000) employed Toda and Yamamoto (1995)'s approach within a VAR framework to examine the relationship between the real exchange rate and investment in Canada and concluded that real exchange rates and their volatility have not had any detectable effect on investment activity (including FDI) in Canada.

This study aims to contribute the debate using quarterly data from Turkey for the period 1987-2000, by means of the Granger non-causality tests using the procedure developed by Toda and Yamamoto (1995). This paper departs from previous studies (e.g., Kosteletou and Liargovas, 2000; and Lafrance and Tessier, 2000) in two major ways. First, the previous studies examine the issue for developed countries. This study examines the issue using data from a developing country. Empirical studies show that the causality relationship between real exchange rate and FDI is valid even for countries which receive relatively low FDI[2]. Using Turkey's data makes the issue interesting since Turkey experienced several financial and currency crises during the last two decades. Even though financial and currency crises in Turkey created undesired macroeconomic environment from the investor's point of view (such as uncertain exchange rates and inflation rates, political instability, changes in the level of interest rates and taxation and increase in the degree of intervention), a notable increase in FDI inflows to Turkey has been observed especially after 1980 which is primarily explained by factors such as liberalized capital flows, privatization, mergers, joint ventures and GSM sales. Second, regarding

[2] For studies related to this topic, see, for example, Benassy-Quere, et.al., (2001) and Goldberg and Klein (1997).

econometric methodology, we apply recent advances in time series analysis: an augmented level VAR modelling proposed by Toda and Yamamoto (1995) which is particularly useful to identify the direction of the long-run relationship between real exchange rates and FDI irrespective of whether the series are integrated of different orders, non-cointegrated or cointegrated of an arbitrary order[3].

The remainder of the paper is organized as follows. The next section reviews theoretical arguments concerning the relationship between exchange rates and FDI. Section 3 describes the data set and presents the econometric methodology. Section 4 displays the empirical results. Section 5 concludes.

II. The Relation Between Exchange Rates and FDI

An ambiguity exists regarding the direction of causality between FDI and exchange rates from a theoretical point of view. First study implying the causality from FDI to real exchange rates was developed by Dornbusch (1973). He pointed out the importance of where foreign capital is directed in domestic economy to determine the effect of foreign capital inflows on real exchange rate movements. If foreign capital inflows are used to finance domestic spending, in the first phase, they will lead to a real exchange rate appreciation and a trade account deficit. In the second phase, trade deficit leads to nominal exchange rate depreciation, and thus, the final effect of the FDI inflows on real exchange rate is ambiguous. If capital inflows are used to finance capital accumulation in the traded goods sector, the supply of traded goods will increase, trade account will improve and the real exchange rate will tend to appreciate. On the other hand, if capital inflows are used to finance capital accumulation in the non-traded goods sector, the trade account will worsen and the real exchange rate will tend to depreciate. The same causal relationship was argued by Branson (1977). He developed a portfolio model in which financial and capital liberalization in a country with a shortage of capital and cheap labor leads to an increase in capital inflows, which in turn, leads to an increased inflation and a real exchange rate appreciation. Edwards (1985)'s empirical work confirms the predictions by Branson (1977) that liberalization of capital account led to a dramatic real appreciation of the Chilean currency between 1979 and 1981.

In contrast to these arguments, some studies implied the reverse causation. The relationship between FDI and exchange rates was discussed by dropping the assumption of perfect capital markets approach by Froot and Stein (1991). They argue that international

[3] Although Lafrance and Tessier (2000) used the same methodology, they focused on one way causation from exchange rate volatility to investment activity in Canada.

wealth, which is significantly influenced by exchange rate changes, has a systematic affect on FDI. A real depreciation of the domestic currency increases the relative wealth of foreign investors relative to that of domestic investors thus leads to an increase in foreign acquisitions of domestic assets and vice versa. By running OLS on US FDI inflows, they obtained results supporting their model that as the US dollar depreciates, FDI into the US rises.

Goldberg and Kolstad (1994) developed two-period model of the intertemporal decision-making of a producer and tested the effect of real exchange rate variability on bilateral FDI flows between the USA, the UK, Canada and Japan for the period 1978-1991. They found that exchange rate volatility had a statistically significant and positive effect on four of the six bilateral FDI shares; real exchange rate variability increased the share of total US investment capacity located in Canada and Japan and increased the share of Canadian and the United Kingdom investment located in the U.S. Even though they also found that exchange rate depreciations of the parent country lead to a reduction in investment flow shares to foreign markets, the coefficients were not statistically significant. Cushman (1985) also argues that causality runs from exchange rates to FDI. He develops a model in which a risk adjusted real foreign currency appreciation lowers the cost of foreign capital and thus stimulate direct investment inflows. However, if the cost of other inputs (such as labor) changes, it might offset the direct effect of real exchange rate changes on FDI. Cushman tested his theoretical models and found that the outward FDI from the US to five industrialized countries is significantly reduced when real value of these countries' currencies is increased.

Numerous empirical studies, which investigate the effects of factors that affect foreign inflows, also imply the causation from exchange rates to FDI. For example, Goldberg and Klein (1997) investigate the relationships among trade, FDI and real exchange rate between a set of South Asian and Latin American countries and both the United States and Japan. They conclude that real exchange rates significantly affect the FDI patterns for South East Asia. Dewenter (1995) investigated the relationship between FDI and exchange rate and found that a depreciating US Dollar is associated with higher levels of foreign acquisitions into the US[4]. The empirical studies typically added exchange rates as one of the explanatory variables to investigate the relationship.

[4] See also Benassy-Quere, A., et.al.'s (2001) paper which analyzes the effect of exchange rate volatility and the choice of an exchange-rate regime on FDI.

III. Data and Econometric Methodology

The data used in this study are quarterly real effective exchange rates (REER), real foreign direct investment and real GDP for the period 1987:1-2000:2 which is determined mainly by availability of data (all of them are transformed to natural logarithms)[5]. REER data are obtained from the OECD's Main Economic Indicators and FDI and GDP data are taken from the IMF's International Financial Statistics (IFS) both of which are deflated using the IFS GDP deflator (1995=100). Since GDP series exhibited strong seasonal patterns, we have seasonally adjusted the GDP data by using difference from moving average-additive method. Since FDI is defined in the US dollars, it is converted to Turkish Lira.

Traditionally, in order to test for the causal relationship between two variables, the literature uses the standard Granger (1969) methodology. The Granger causality test assumes that the two variables, *X* and *Y*, are stationary. However, many time series are found in empirical studies that they are non-stationary or integrated of order 1. If the variables are non-stationary, or, I(1), the estimates of models in levels will end up spurious regression results and the F-test is not valid unless the variables are co-integrated. If the variables are I(1) and co-integrated, granger causality tests should be performed using the error correction mechanism in which the first differenced variables are estimated with the one-period lagged value of the error term from the cointegrating regression (Engel and Granger, 1987).

Therefore, tests of stationary should precede tests of causality. In this study, the order of integration of the individual time series is determined using the Phillips-Perron (PP) test suggested by Phillips and Perron (1988) and the augmented Dickey-Fuller (ADF) test developed by Dickey and Fuller (1979).

When there is uncertainty concerning the order of integration/cointegration of variables, neither the standard Granger causality test nor the error correction mechanism due to Engle and Granger (1987) can be employed because standard Granger causality test can be applied only with stationary variables and the error correction model can be applied if all variables are I(1). Consequently, when the distinct order of integration of the variables exists, another methodology is required to test for the causality between two variables.

Following the literature, the non-causality test proposed by Toda and Yamamoto (1995) is particularly applicable to test for the causal relationship between two variables when the model includes a mixture

[5] The lower period limit is determined as 1987:1 due to the lack of quarterly GDP data before that period. The upper period limit is determined as 2000:2 due to negative value of FDI data at the 2000:3 period which causes missing observation when it is transformed to logarithm.

of stationary or non-stationary variables and avoids pre-test biases associated with prior tests for non-stationary and cointegration. Their method employs an augmented VAR model in levels thus the methodology has been considered as the long-run causality tests. The methodology is implemented in two steps. In the first step, the appropriate lag length of a VAR (k) and the maximum order of integration (d_{max}) of the variables in the system are determined. The number of lags in the system can be chosen using measures such as the Akaike Information Criterion (AIC), Akaike's final prediction error or Schwarz information criterion (SC). Having determined the lag structure of VAR (k) and the maximum order of integration (d_{max}), a VAR in level is estimated with a total of $p = [k + (d_{max})]$ lags. In the second step, the conventional Wald test is applied on the first k coefficient matrices using the standard χ^2-statistic. Toda and Yamamoto (1995) point out that so long as $k \geq d$ the procedure is valid. In estimation of the VAR and application of the Wald tests, SUR method provides efficient estimates of the VAR coefficients (Johnston and DiNardo; 1997 and Enders; 1995).

IV. Empirical Results

In order to determine the maximum order of integration (d_{max}) of the variables in the system the PP and ADF tests are performed with and without constant and with constant and trend variable. Table 1 and Table 2 display the results. The two tests yield the same results that with the exception of the real effective exchange rate series, which is I(1) in all specifications, the tests on the levels of FDI and real GDP series do not give a unique characterization. However, the tests indicate that all series are I(0) in their first differenced forms.

Table 1: The Phillips-Perron (PP) Tests for Unit Roots

	A. Levels			B. First Differences		
	LFDI	LREER	LGDP	ΔLFDI	ΔLREER	ΔLGDP
τ_t	-5.14[a]	-2.00	-4.05[b]	-11.34[a]	-6.17[a]	-8.12[a]
τ_t	-4.53[a]	-1.81	-0.45	-11.49[a]	-6.20[a]	-8.23[a]
τ	-1.00	0.55	2.27	-11.48[a]	-6.23[a]	-7.54[a]

Note: Critical values are taken from MacKinnon (1991). [a],[b] indicate rejection of the null hypothesis of a unit root at the 1% and 5% significance levels, respectively. τ_t, τ_t, τ are the PP Test Statistic with intercept and trend, with intercept and without intercept, respectively. Truncation lag is selected by the Newey-West automatic truncation lag selection.

Table 2: The Augmented Dickey-Fuller (ADF) Tests

	A. Levels			B. First Differences		
	LFDI	LREER	LGDP	ΔLFDI	ΔLREER	ΔLGDP
τ_γ	-5.04[a]	-2.22	-3.86[b]	-3.60[b]	-5.63[a]	-6.15[a]
τ_γ	-5.12[a]	-2.06	-0.22	-4.06[a]	-5.65[a]	-6.30[a]
τ	1.16	0.48	2.48	-4.16[a]	-5.65[a]	-5.50[a]

Note: Critical values are taken from MacKinnon (1991). [a],[b] indicates rejection of the null hypothesis of a unit root at the 1% and 5% significance levels, respectively. τ_t, τ_γ, τ are the ADF Test Statistic with intercept and trend, with intercept and without intercept, respectively. Lags are chosen by Schwarz criterion.

Since there is uncertainty concerning the order of integration/cointegration of variables, we employ the following lag-augmented VAR model proposed by Toda and Yamamoto (1995):

$$\text{LFDI}= \alpha + \sum_{i=1}^{k} \delta_i LFDI_{t-i} + \sum_{j=k+1}^{k+d\max} \delta_j' LFDI_{t-j} + \sum_{i=1}^{k} \phi_i LREER_{t-i} + \sum_{j=k+1}^{k+d\max} \phi_j' LREER_{t-j} + \gamma EX_t + \varepsilon_{1t} \quad (1)$$

$$\text{LREER}= \alpha + \sum_{i=1}^{k} \beta_i LREER_{t-i} + \sum_{j=k+1}^{k+d\max} \beta_j' LREER_{t-j} + \sum_{i=1}^{k} \lambda_i LFDI_{t-i} + \sum_{j=k+1}^{k+d\max} \lambda_j' LFDI_{t-j} + \gamma' EX_t + \varepsilon_{2t} \quad (2)$$

where FDI and REER are inward foreign direct investment and real effective exchange rates, respectively and ε_{1t}, ε_{2t} are uncorrelated white-noise disturbances and EX_t indicates a control variable which is related to both variables. In this case, GDP is included as a control variable since it is possible that both FDI and REER are related to GDP[6]. The two equations are jointly estimated as a seemingly unrelated regressions (SUR) model by maximum likelihood method. In order to test that REER 'Granger causes' FDI, we test the parameter restrictions $\phi_i \neq 0 \ \forall_i$ and to test that FDI 'Granger causes' REER we test the parameter restrictions $\beta_i \neq 0 \ \forall_i$ using the Wald statistic. The Wald

[6] The measures such as AIC, SC and Akaike's Final Predictor Error also provide that models include GDP give better goodness of fit.

statistic will be asymptotically distributed as a chi-square with degrees of freedom equal to the number of zero restrictions, irrespective of FDI and REER are I(0) or I(1), cointegrated of an arbitrary order or non-cointegrated.

Table 3 presents the results of the WALD test statistic as well as its *p*-values. The results suggest that while we fail to reject the null hypothesis of *Granger non-causality* from real effective exchange rate to inward foreign direct investment we can reject the null hypothesis of *Granger non-causality* from inward foreign direct investment to real effective exchange rate. The results also reveal that an increase in the inward foreign direct investment leads to the real effective exchange rate appreciation.[7]

Table 3. Results of the Toda and Yamamoto Wald Test

Null Hypothesis	Wald Statistics	*p*-value
Real effective exchange rate does not Granger-cause FDI	0.15	0.693
FDI does not Granger-cause real effective exchange rate	3.15	0.075

Note: The order of *k* was chosen to be 1 based on sequential modified LR test statistic, Final Prediction Error, Akaike information criterion, Schwarz information criterion and Hannan-Quinn information criterion. The order of d_{max} was chosen to be 1 since each variable contains a maximum of one unit root. A linear deterministic trend is included in equations (5) and (6) to capture the deterministic trend. GDP was added as a control variable.

V. Conclusions

In this article, the direction of a causal relationship between inward FDI and real exchange rates has been analyzed using quarterly data from Turkey for the period 1987-2000. An ambiguity exists in the theoretical literature regarding the direction of such a relationship. In order to determine appropriate methodological procedure to test for Granger causality, the relevant data were examined and the augmented level VAR model proposed by Toda and Yamamoto (1995) was used. The procedure can be applied when series are integrated of different order, non-cointegrated or cointegrated of an arbitrary order.

The empirical results indicate that Granger causality is unidirectional and is running from FDI to real effective exchange rates. In addition, an increase in the inward FDI leads to an appreciation of the

[7] These test results are not reported here to focus on the direction of causality, but are vailable from the authors.

real effective exchange rates. The results of this article would generally agree with the predictions from theoretical models by Dornbusch (1973) and Branson (1977) among which the causal relationship is seemed to be explained better by Branson's (1977) model. It can thus be inferred that financial and capital liberalization in Turkey with a shortage of capital and cheap labor leads to an increase in capital inflows, which in turn, a real exchange rate appreciation.

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GLOBAL CAPITAL MARKETS

The global economic continued to exceed expectations due to benign financial market conditions and accommodative macroeconomic policies. Despite higher oil prices and natural diasters, activity in the second half of 2005 was stronger than earlier projected, particularly in China, India and Russia. The global GDP growth is projected at 4.9 percent in 2006, which is 0.6 percentage point higher than expected last September.

Global financial market conditions remain very favorable characterized by unusually low risk premiums and volatility. Global short-term interest rates have continued to rise, led by the United States. In the emerging markets financing conditions are very favourable in part reflecting improved economic fundamentals. Accompanied by buoyany inflows to emerging markets equity prices have risen significantly, particularly outside the United States.

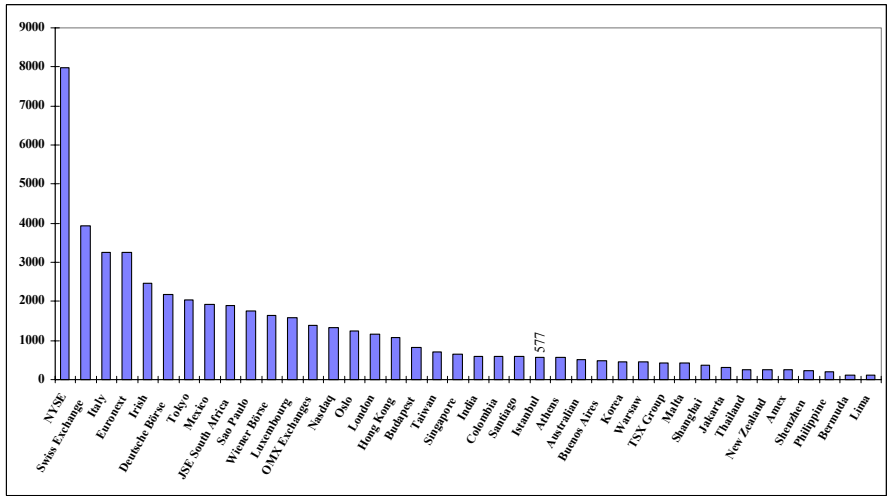
The performances of some developed stock markets with respect to indices indicated that DJIA, FTSE-100, Nikkei-225 and DAX changed by 3.66 %, 6.15 %, 5.89 % and 10.39 % respectively at March 31, 2006 in comparison with the December 30, 2005. When US $ based returns of some emerging markets are compared in the same period, the best performer markets were: Venezuela (60.9 %), Russia (31.9 %), Peru (28.8 %), Brazil (27.6 %), Indonesia (26.2 %), India (26.1 %) and Pakistan (21.4%). In the same period, the lowest return markets were: Taiwan (3.0 %), Israel (3.7 %) and Saudi Arabia (5.7 %). The performances of emerging markets with respect to P/E ratios as of end-March 2006 indicated that the highest rates were obtained in Jordan (52.1), Russia (29.3), India (23.1), Czech Rep. (22.3) and Taiwan (22.3) and the lowest rates in S. Africa (6.2), Venezuela (7.3), Thailand (10.0), Brazil (11.0) and Peru (11.9). In the same period Istanbul Stock Exchange's P/E ratio was 17.2.

Market Capitalization (USD Million, 1986-2005)

	Global	Developed Markets	Emerging Markets	ISE
1986	6,514,199	6,275,582	238,617	938
1987	7,830,778	7,511,072	319,706	3,125
1988	9,728,493	9,245,358	483,135	1,128
1989	11,712,673	10,967,395	745,278	6,756
1990	9,398,391	8,784,770	613,621	18,737
1991	11,342,089	10,434,218	907,871	15,564
1992	10,923,343	9,923,024	1,000,319	9,922
1993	14,016,023	12,327,242	1,688,781	37,824
1994	15,124,051	13,210,778	1,913,273	21,785
1995	17,788,071	15,859,021	1,929,050	20,782
1996	20,412,135	17,982,088	2,272,184	30,797
1997	23,087,006	20,923,911	2,163,095	61,348
1998	26,964,463	25,065,373	1,899,090	33,473
1999	36,030,810	32,956,939	3,073,871	112,276
2000	32,260,433	29,520,707	2,691,452	69,659
2001	27,818,618	25,246,554	2,572,064	47,150
2002	23,391,914	20,955,876	2,436,038	33,958
2003	31,947,703	28,290,981	3,656,722	68,379
2004	38,904,018	34,173,600	4,730,418	98,299
2005	43,642,048	36,538,248	7,103,800	161,537

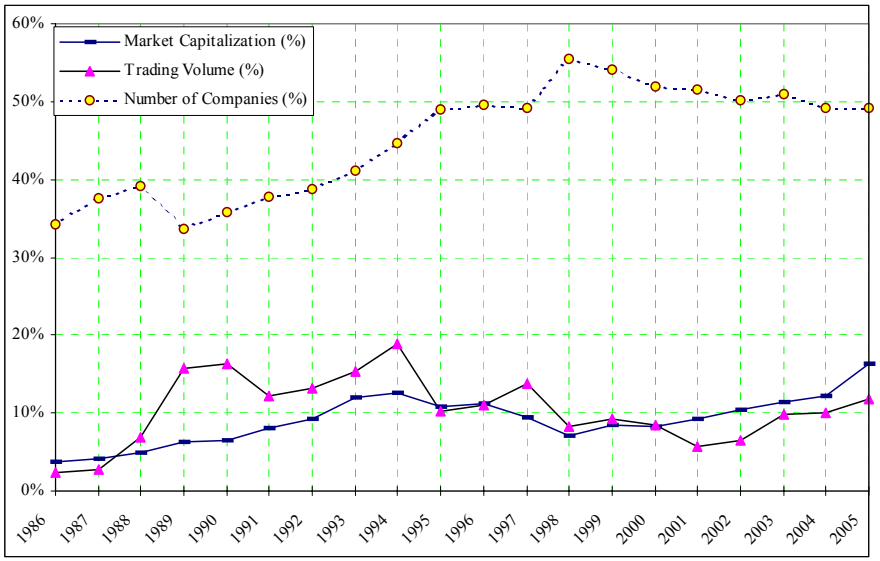
Source: Standard & Poor's Global Stock Markets Factbook, 2006.

Comparison of Average Market Capitalization Per Company (USD Million, March 2006)



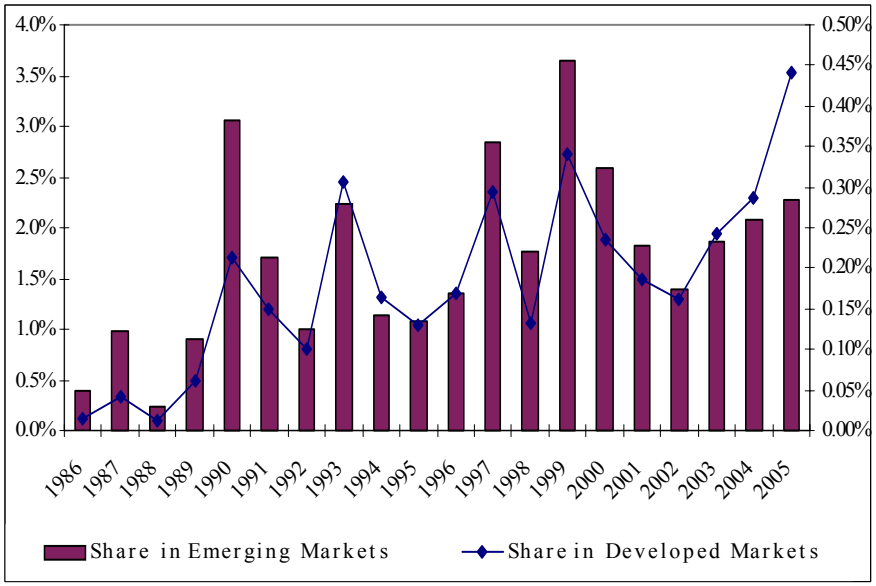
Source: FIBV, Monthly Statistics, March 2006.

Worldwide Share of Emerging Capital Markets (1986-2005)



Source: Standard & Poor's Global Stock Markets Factbook, 2006.

Share of ISE's Market Capitalization in World Markets (1986-2005)



Source: Standard & Poor's Global Stock Markets Factbook, 2006.

Main Indicators of Capital Markets (March 2006)

	Market	Monthly Turnover Velocity (March 2006) (%)	Market	Value of Share Trading (millions, US$) Up to Year Total (2006/1-2006/3)	Market	Market Cap. of Share of Domestic Companies (millions US$) March 2006
1	Nasdaq	255.71	NYSE	5,332,691	NYSE	14,072,267
2	Korea	221.31	Nasdaq	3,139,508	Tokyo	4,774,550
3	Italy	163.18	London	1,903,592	Nasdaq	3,786,552
4	Istanbul	158.67	Tokyo	1,643,368	London	3,253,372
5	Spanish Exchanges (BME)	158.53	Euronext	917,816	Osaka	3,146,920
6	Deutsche Börse	157.94	Deutsche Börse	668,012	Euronext	3,115,780
7	Shenzhen	145.94	Korea	438,534	TSX Group	1,620,894
8	Taiwan	139.83	Spanish (BME)	421,663	Deutsche Börse	1,399,555
9	OMX Exchanges	127.47	Swiss Exchange	381,990	Hong Kong	1,213,418
10	Oslo	126.70	Italy	353,089	Spanish (BME)	1,094,120
11	Tokyo	125.16	OMX Exchanges	336,815	Swiss Exchange	1,015,937
12	Swiss Exchange	122.90	TSX Group	319,476	OMX Exchanges	917,774
13	Euronext	117.03	Hong Kong	196,804	Italy	896,506
14	London	114.21	Australian	193,216	Australian	858,389
15	NYSE	103.53	Taiwan	179,910	Korea	751,687
16	Shanghai	93.52	Amex	170,812	Bombay	678,153
17	Australian	85.76	India	111,410	JSE South Africa	665,626
18	Thailand	78.25	Oslo	96,314	India	631,258
19	Budapest	75.30	Shanghai	88,093	Sao Paulo	583,354
20	India	74.86	JSE-South Africa	84,819	Taiwan	490,433
21	TSX Group	70.79	Osaka	76,061	Shanghai	313,483
22	Irish	59.44	Istanbul	73,395	Singapore	293,961
23	Athens	53.29	Sao Paulo	66,625	Mexico	251,611
24	Hong Kong	53.08	Bombay	60,459	Oslo	238,434
25	Singapore	52.19	Shenzhen	52,983	Malaysia	195,397
26	JSE South Africa	46.92	Singapore	44,578	Istanbul	176,561
27	Jakarta	46.08	Thailand	32,304	Athens	169,827
28	New Zealand	45.87	Athens	29,918	Wiener Börse	149,829
29	Tel-Aviv	45.46	Mexico	22,295	Santiago	143,251
30	Wiener Börse	43.72	Irish	20,085	Thailand	135,276
31	Sao Paulo	42.78	Wiener Börse	18,968	Tel-Aviv	132,338
32	Warsaw	38.98	Tel-Aviv	16,110	Irish	127,498
33	Bombay	35.99	Malaysia	15,438	Shenzhen	127,341
34	Colombia	28.93	Warsaw	12,393	Amex	122,365
35	Mexico	27.42	Jakarta	9,607	Warsaw	104,855
36	Malaysia	27.15	Budapest	7,161	Jakarta	100,204
37	Philippine	18.00	Santiago	5,741	Luxembourg	59,115
38	Colombo	17.70	New Zealand	5,288	Colombia	57,947
39	Santiago	15.84	Colombia	4,724	Buenos Aires	46,962
40	Tehran	15.57	Philippine	2,392	Philippine	45,360
41	Ljubljana	14.53	Buenos Aires	1,668	New Zealand	38,711
42	Lima	13.72	Lima	1,334	Budapest	34,719
43	Buenos Aires	10.14	Tehran	881	Tehran	33,660
44	Osaka	8.80	Ljubljana	383	Lima	23,310
45	Malta	6.02	Colombo	260	Ljubljana	8,192

Source: FIBV, Monthly Statistics, March 2006.

Trading Volume (USD millions, 1986-2005)

	Global	Developed	Emerging	ISE	Emerging / Global (%)	ISE/Emerging (%)
1986	3,573,570	3,490,718	82,852	13	2.32	0.02
1987	5,846,864	5,682,143	164,721	118	2.82	0.07
1988	5,997,321	5,588,694	408,627	115	6.81	0.03
1989	7,467,997	6,298,778	1,169,219	773	15.66	0.07
1990	5,514,706	4,614,786	899,920	5,854	16.32	0.65
1991	5,019,596	4,403,631	615,965	8,502	12.27	1.38
1992	4,782,850	4,151,662	631,188	8,567	13.20	1.36
1993	7,194,675	6,090,929	1,103,746	21,770	15.34	1.97
1994	8,821,845	7,156,704	1,665,141	23,203	18.88	1.39
1995	10,218,748	9,176,451	1,042,297	52,357	10.20	5.02
1996	13,616,070	12,105,541	1,510,529	37,737	11.09	2.50
1997	19,484,814	16,818,167	2,666,647	59,105	13.69	2.18
1998	22,874,320	20,917,462	1,909,510	68,646	8.55	3.60
1999	31,021,065	28,154,198	2,866,867	81,277	9.24	2.86
2000	47,869,886	43,817,893	4,051,905	179,209	8.46	4.42
2001	42,076,862	39,676,018	2,400,844	77,937	5.71	3.25
2002	38,645,472	36,098,731	2,546,742	70,667	6.59	2.77
2003	29,639,297	26,743,153	2,896,144	99,611	9.77	3.44
2004	39,309,589	35,341,782	3,967,806	147,426	10.09	3.72
2005	47,319,584	41,715,492	5,604,092	201,258	11.84	3.59

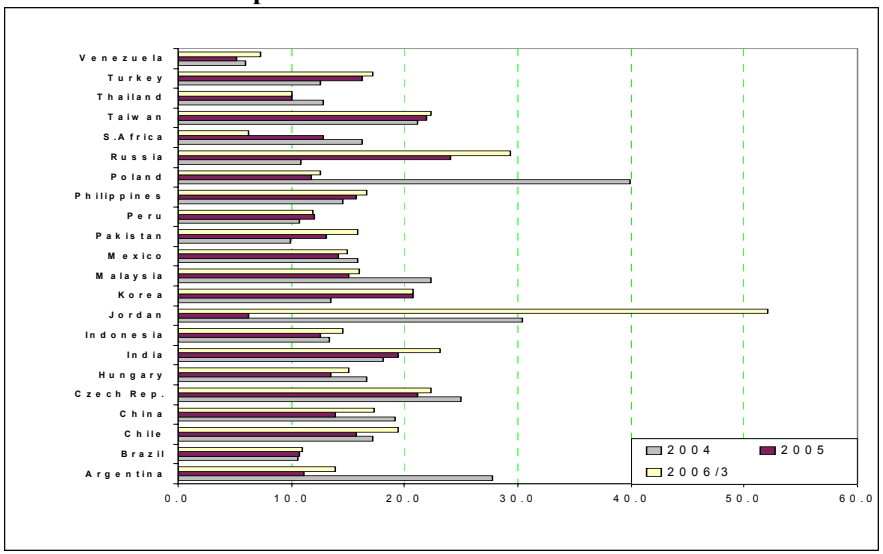
Source: Standard & Poor's Global Stock Markets Factbook, 2006.

Number of Trading Companies (1986-2005)

	Global	Developed Markets	Emerging Markets	ISE	Emerging / Global (%)	ISE/Emerging (%)
1986	28,173	18,555	9,618	80	34.14	0.83
1987	29,278	18,265	11,013	82	37.62	0.74
1988	29,270	17,805	11,465	79	39.17	0.69
1989	25,925	17,216	8,709	76	33.59	0.87
1990	25,424	16,323	9,101	110	35.80	1.21
1991	26,093	16,239	9,854	134	37.76	1.36
1992	27,706	16,976	10,730	145	38.73	1.35
1993	28,895	17,012	11,883	160	41.12	1.35
1994	33,473	18,505	14,968	176	44.72	1.18
1995	36,602	18,648	17,954	205	49.05	1.14
1996	40,191	20,242	19,949	228	49.64	1.14
1997	40,880	20,805	20,075	258	49.11	1.29
1998	47,465	21,111	26,354	277	55.52	1.05
1999	48,557	22,277	26,280	285	54.12	1.08
2000	49,933	23,996	25,937	315	51.94	1.21
2001	48,220	23,340	24,880	310	51.60	1.25
2002	48,375	24,099	24,276	288	50.18	1.19
2003	49,855	24,414	25,441	284	51.03	1.12
2004	48,806	24,824	23,982	296	49.14	1.23
2005	49,946	25,337	24,609	302	49.27	1.23

Source: Standard & Poor's Global Stock Markets Factbook, 2006.

Comparison of P/E Ratios Performances



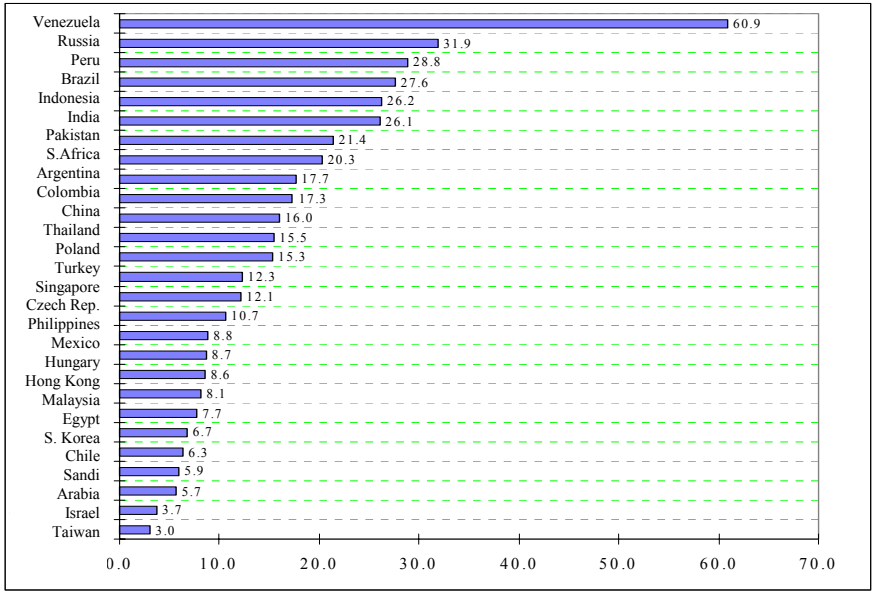
Source: IFC Factbook 2001. Standard & Poor's, Emerging Stock Markets Review, March 2006.

Price-Earnings Ratios in Emerging Markets

	1997	1998	1999	2000	2001	2002	2003	2004	2005	2006/3
Argentina	16.3	13.4	39.4	-889.9	32.6	-1.4	21.1	27.7	11.1	13.9
Brazil	12.4	7.0	23.5	11.5	8.8	13.5	10.0	10.6	10.7	11.0
Chile	14.7	15.1	35.0	24.9	16.2	16.3	24.8	17.2	15.7	19.4
China	34.5	23.8	47.8	50.0	22.2	21.6	28.6	19.1	13.9	17.3
Czech Rep.	37.1	-11.3	-14.9	-16.4	5.8	11.2	10.8	25.0	21.1	22.3
Hungary	27.4	17.0	18.1	14.3	13.4	14.6	12.3	16.6	13.5	15.0
India	15.2	13.5	25.5	16.8	12.8	15.0	20.9	18.1	19.4	23.1
Indonesia	10.5	-106.2	-7.4	-5.4	-7.7	22.0	39.5	13.3	12.6	14.5
Jordan	14.4	15.9	14.1	13.9	18.8	11.4	20.7	30.4	6.2	52.1
Korea	17.9	-47.1	-33.5	17.7	28.7	21.6	30.2	13.5	20.8	20.8
Malaysia	9.5	21.1	-18.0	91.5	50.6	21.3	30.1	22.4	15	16
Mexico	19.2	23.9	14.1	13.0	13.7	15.4	17.6	15.9	14.2	14.9
Pakistan	14.8	7.6	13.2	-117.4	7.5	10.0	9.5	9.9	13.1	15.8
Peru	14.0	21.1	25.7	11.6	21.3	12.8	13.7	10.7	12.0	11.9
Philippines	10.9	15.0	22.2	26.2	45.9	21.8	21.1	14.6	15.7	16.6
Poland	11.4	10.7	22.0	19.4	6.1	88.6	-353.0	39.9	11.7	12.5
Russia	8.1	3.7	-71.2	3.8	5.6	12.4	19.9	10.8	24.1	29.3
S.Africa	10.8	10.1	17.4	10.7	11.7	10.1	11.5	16.2	12.8	6.2
Taiwan	28.9	21.7	52.5	13.9	29.4	20.0	55.7	21.2	21.9	22.3
Thailand	-32.8	-3.6	-12.2	-6.9	163.8	16.4	16.6	12.8	10.0	10.0
Turkey	20.1	7.8	34.6	15.4	72.5	37.9	14.9	12.5	16.2	17.2
Venezuela	12.8	5.6	10.8	30.5	-347.6	-11.9	14.4	6.0	5.1	7.3

Source: IFC Factbook, 2004; Standard&Poor's, Emerging Stock Markets Review, March 2006
Note : Figures are taken from S&P/IFCG Index Profile.

Comparison of Market Returns in USD (30/12/2005-05/04/2006)



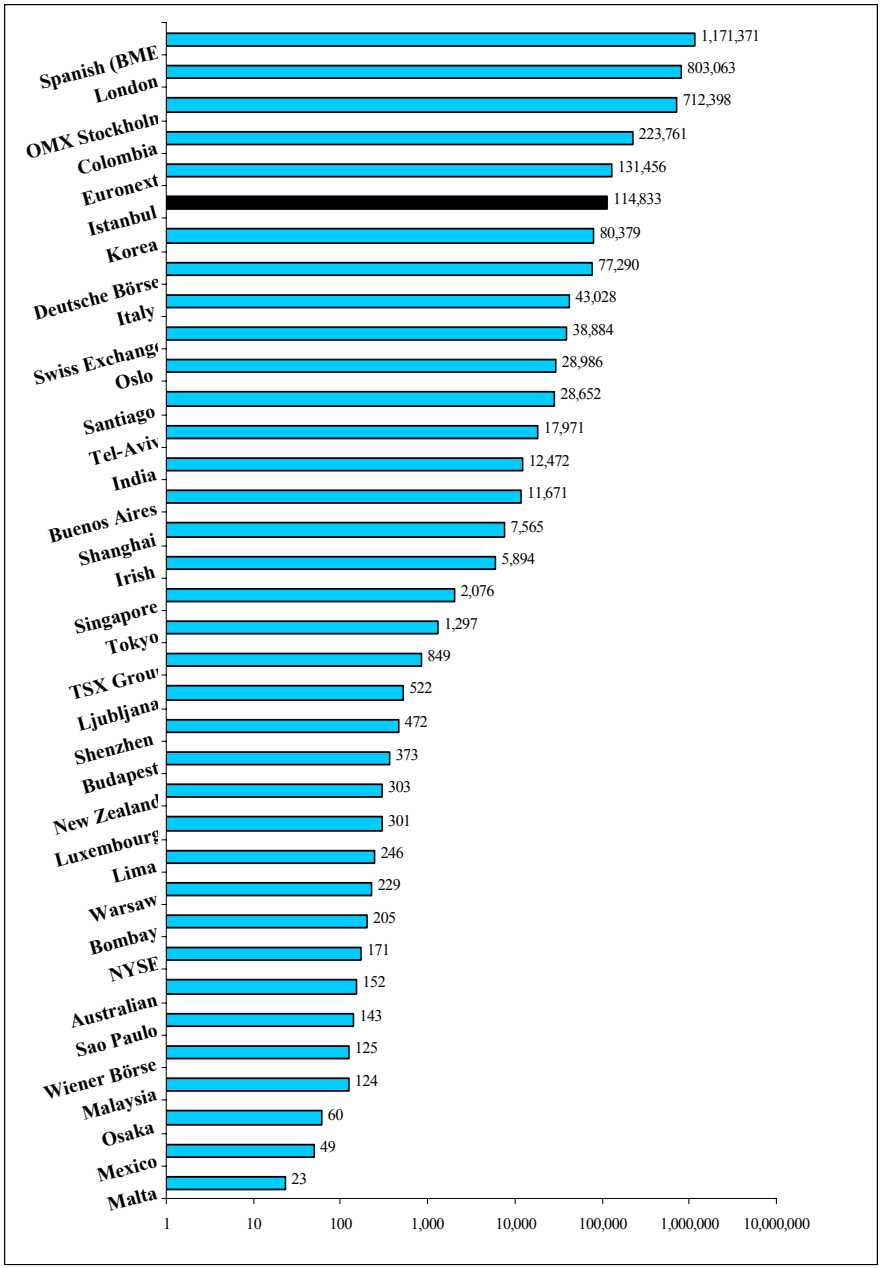
Source: The Economist, April 8th 2006.

Market Value/Book Value Ratios

	1997	1998	1999	2000	2001	2002	2003	2004	2005	2006/3
Argentina	1.8	1.3	1.5	0.9	0.6	0.8	2.0	2.2	2.5	3.1
Brazil	1.0	0.6	1.6	1.4	1.2	1.3	1.8	1.9	2.2	2.4
Chile	1.6	1.1	1.7	1.4	1.4	1.3	1.9	0.6	1.9	2.0
China	3.9	2.1	3.0	3.6	2.3	1.9	2.6	2.0	1.8	2.3
Czech Rep.	0.8	0.7	0.9	1.0	0.8	0.8	1.0	1.6	2.4	2.5
Hungary	4.2	3.2	3.6	2.4	1.8	1.8	2.0	2.8	3.1	3.4
India	2.3	1.8	3.3	2.6	1.9	2.0	3.5	3.3	5.2	6.1
Indonesia	1.4	1.5	3.0	1.7	1.7	1.0	1.6	2.8	2.5	2.9
Jordan	1.8	1.8	1.5	1.2	1.5	1.3	2.1	3.0	2.2	5.7
Korea	0.5	0.9	2.0	0.8	1.2	1.1	1.6	1.3	2.0	2.0
Malaysia	1.4	1.3	1.9	1.5	1.2	1.3	1.7	1.9	1.7	1.8
Mexico	2.3	1.4	2.2	1.7	1.7	1.5	2.0	2.5	2.9	3.1
Pakistan	2.3	0.9	1.4	1.4	0.9	1.9	2.3	2.6	3.5	4.2
Peru	2.0	1.6	1.5	1.1	1.4	1.2	1.8	1.6	2.2	2.6
Philippines	1.3	1.3	1.4	1.0	0.9	0.8	1.1	1.4	1.7	1.8
Poland	1.7	1.5	2.0	2.2	1.4	1.3	1.8	2.0	2.5	2.7
Russia	0.5	0.3	1.2	0.6	1.1	0.9	1.2	1.2	2.2	2.6
S.Africa	1.6	1.5	2.7	2.1	2.1	1.9	2.1	2.5	3.0	2.6
Taiwan	3.1	2.6	3.4	1.7	2.1	1.6	2.2	1.9	1.9	2.0
Thailand	0.8	1.2	2.1	1.3	1.3	1.5	2.8	2.0	2.1	2.0
Turkey	6.8	2.7	8.9	3.1	3.8	2.8	2.6	1.7	2.1	2.3
Venezuela	1.2	0.5	0.4	0.6	0.5	0.5	1.1	1.2	0.7	1.0

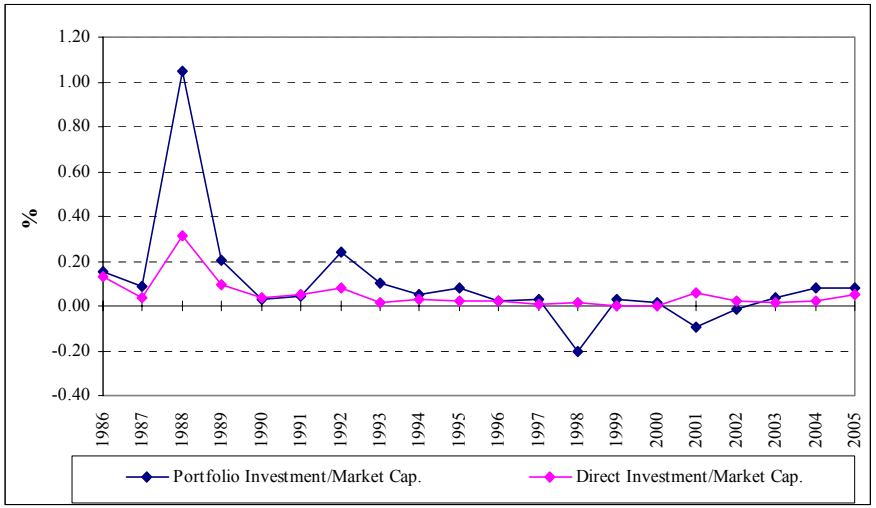
Source: IFC Factbook, 2004; Standard & Poor's, Emerging Stock Markets Review, March 2006.
Note: Figures are taken from S&P/IFCG Index Profile.

Value of Bond Trading (Million USD Jan. 2006-March 2006)



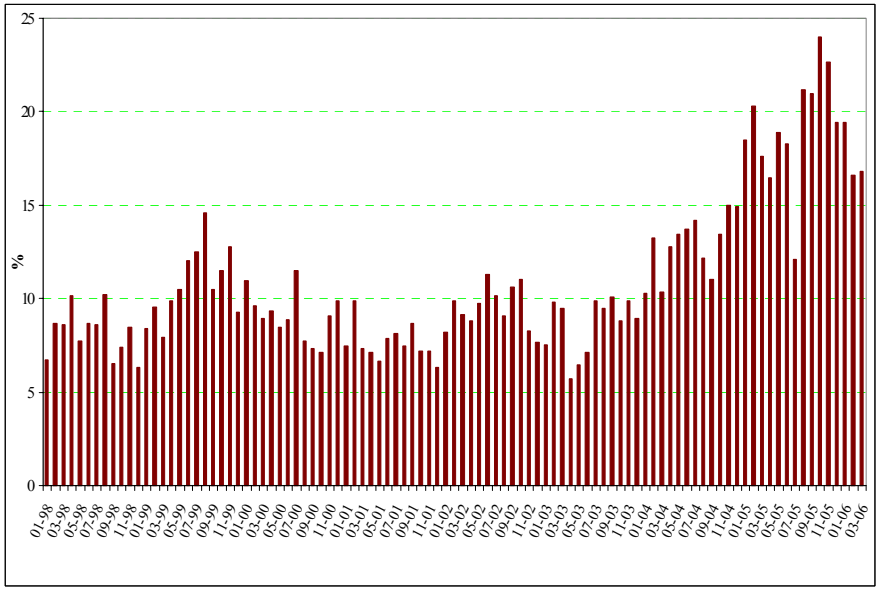
Source: FIBV, Monthly Statistics, March 2006.

Foreign Investments as a Percentage of Market Capitalization in Turkey (1986-2005)



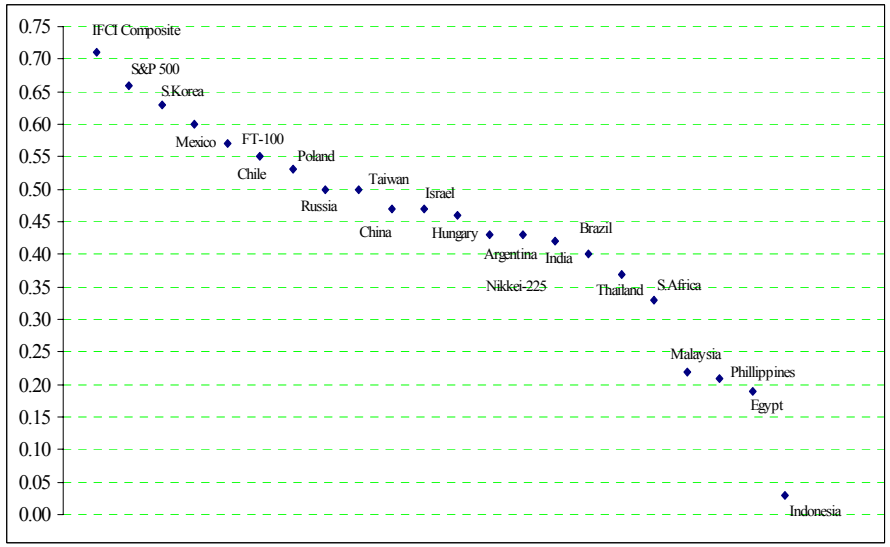
Source: ISE Data. CBTR Databank.

Foreigners' Share in the Trading Volume of the ISE (Jan. 1998-March 2006)



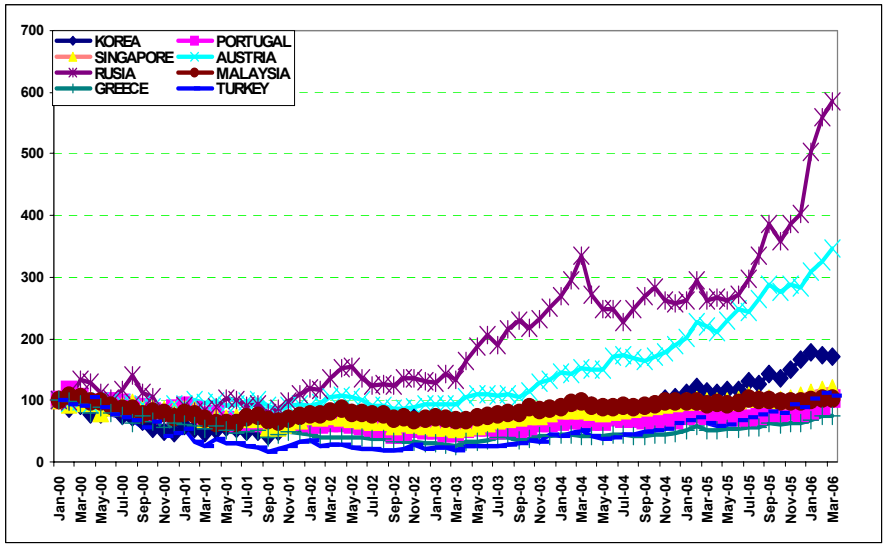
Source: ISE Data.

Price Correlations of the ISE (March 2001- March 2006)



Source: Standard & Poor's, Emerging Stock Markets Review, March 2006.
Notes: The correlation coefficient is between -1 and +1. If it is zero. for the given period. it is implied that there is no relation between two serious of returns.

Comparison of Market Indices (31 Jan. 2000=100)



Source: Reuters.
Note: Comparisons are in US$.

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ISE
Market Indicators

STOCK MARKET

	Number of Companies	Traded Value				Market Value		Dividend Yield	P/E Ratios		
		Total		Daily Average							
		(YTL Million)	(US$ Million)	(YTL Million)	(US$ Million)	(YTL Million)	(US$ Million)	(%)	YTL (1)	YTL (2)	US$
1986	80	0,01	13	---	---	0,71	938	9,15	5,07	---	---
1987	82	0,11	118	---	---	3	3.125	2,82	15,86	---	---
1988	79	0,15	115	---	---	2	1.128	10,48	4,97	---	---
1989	76	2	773	0,01	3	16	6.756	3,44	15,74	---	---
1990	110	15	5.854	0,06	24	55	18.737	2,62	23,97	---	---
1991	134	35	8.502	0,14	34	79	15.564	3,95	15,88	---	---
1992	145	56	8.567	0,22	34	85	9.922	6,43	11,39	---	---
1993	160	255	21.770	1	88	546	37.824	1,65	25,75	20,72	14,86
1994	176	651	23.203	3	92	836	21.785	2,78	24,83	16,70	10,97
1995	205	2.374	52.357	9	209	1.265	20.782	3,56	9,23	7,67	5,48
1996	228	3.031	37.737	12	153	3.275	30.797	2,87	12,15	10,86	7,72
1997	258	9.049	58.104	36	231	12.654	61.879	1,56	24,39	19,45	13,28
1998	277	18.030	70.396	73	284	10.612	33.975	3,37	8,84	8,11	6,36
1999	285	36.877	84.034	156	356	61.137	114.271	0,72	37,52	34,08	24,95
2000	315	111.165	181.934	452	740	46.692	69.507	1,29	16,82	16,11	14,05
2001	310	93.119	80.400	375	324	68.603	47.689	0,95	108,33	824,42	411,64
2002	288	106.302	70.756	422	281	56.370	34.402	1,20	195,92	26,98	23,78
2003	285	146.645	100.165	596	407	96.073	69.003	0,94	14,54	12,29	13,19
2004	297	208.423	147.755	837	593	132.556	98.073	1,37	14,18	13,27	13,96
2005	304	269.931	201.763	1.063	794	218.318	162.814	1,71	17,19	19,38	19,33
2006	306	98.235	74.208	1.637	1.237	237.563	177.180	1,81	23,27	21,63	21,59
2006/Q1	306	98.235	74.208	1.637	1.237	237.563	177.180	1,81	23,27	21,63	21,59

Q: Quarter

Note:

* Between 1986-1992, the price earnings ratios were calculated on the basis of the companies' previous year-end net profits. As from 1993,
TL(1) = Total Market Capitalization / Sum of Last two six-month profits
T(2) = Total Market Capitalization / Sum of Last four three-month profits.
US$ = US$ based Total Market Capitilization / Sum of Last four US$ based three-month profits.

* Companies which are temporarily de-listed and will be traded off the Exchange under the decison of ISE's Board of Directors are not included in the calculations.

* ETF's data are taken into account only in the calculation of Traded Value.

Closing Values of the ISE Price Indices

YTL Based

	NATIONAL -100 (Jan. 1986=1)	NATIONAL - INDUSTRIALS (Dec. 31.90=33)	NATIONAL - SERVICES (Dec. 27,96=1046)	NATIONAL - FINANCIALS (Dec.31.90=33)	NATIONAL - TECHNOLOGY (Jun.30.2000=14.466,12)	'INVESTMENT TRUSTS (Dec 27,1996=976)	'SECOND NATIONAL (Dec.27,1996=976)	'NEW ECONOMY (Sept. 02, 2004 =20525,92)
1986	1,71	---	---	---	---	---	---	---
1987	6,73	---	---	---	---	---	---	---
1988	3,74	---	---	---	---	---	---	---
1989	22,18	---	---	---	---	---	---	---
1990	32,56	---	---	---	---	---	---	---
1991	43,69	49,63	---	33,55	---	---	---	---
1992	40,04	49,15	---	24,34	---	---	---	---
1993	206,83	222,88	---	191,90	---	---	---	---
1994	272,57	304,74	---	229,64	---	---	---	---
1995	400,25	462,47	---	300,04	---	---	---	---
1996	975,89	1.045,91	---	914,47	---	---	---	---
1997	3.451,--	2.660,--	3.593,--	4.522,--	---	2.934,--	2.761,--	---
1998	2.597,91	1.943,67	3.697,10	3.269,58	---	1.579,24	5.390,43	---
1999	15.208,78	9.945,75	13.194,40	21.180,77	---	6.812,65	13.450,36	---
2000	9.437,21	6.954,99	7.224,01	12.837,92	10.586,58	6.219,00	15.718,65	---
2001	13.782,76	11.413,44	9.261,82	18.234,65	9.236,16	7.943,60	20.664,11	---
2002	10.369,92	9.888,71	6.897,30	12.902,34	7.260,84	5.452,10	28.305,78	---
2003	18.625,02	16.299,23	9.923,02	25.594,77	8.368,72	10.897,76	32.521,26	---
2004	24.971,68	20.885,47	13.914,12	35.487,77	7.539,16	17.114,91	23.415,86	39.240,73
2005	39.777,70	31.140,59	18.085,71	62.800,64	13.669,97	23.037,86	28.474,96	29.820,90
2006	42.911,32	32.961,91	20.018,86	68.636,50	15.558,11	22.089,86	30.205,93	31.556,79
2006/Q1	42.911,32	32.961,91	20.018,86	68.636,50	15.558,11	22.089,86	30.205,93	31.556,79

US $ Based

	NATIONAL - 100 (Jan. 1986=100)	NATIONAL - INDUSTRIALS (Dec. 31, 90=643)	NATIONAL - SERVICES (Dec. 27, 96=572)	NATIONAL - FINANCIALS (Dec.31, 90=643)	NATIONAL - TECHNOLOGY (Jun. 30,2000=1.360.92)	'INVESTMENT TRUSTS (Dec. 27, 96=534)	'SECOND NATIONAL (Dec. 27, 96=534)	'NEW ECONOMY (Sept. 02, 2004 =796,46)	EURO Based NATIONAL - 100 (Dec. 31, 98=484)
1988	119,82	---	---	---	---	---	---	---	---
1989	560,57	---	---	---	---	---	---	---	---
1990	642,63	---	---	---	---	---	---	---	---
1991	501,50	569,63	---	385,14	---	---	---	---	---
1992	272,61	334,59	---	165,68	---	---	---	---	---
1993	833,28	897,96	---	773,13	---	---	---	---	---
1994	413,27	462,03	---	348,18	---	---	---	---	---
1995	382,62	442,11	---	286,83	---	---	---	---	---
1996	534,01	572,33	---	500,40	---	---	---	---	---
1997	981,99	756,91	1.022,40	1.286,75	---	834,88	785,65	---	---
36160	484,--	362,--	689,--	609,--	---	294,--	1.004,--	---	---
1999	1.654,17	1.081,74	1.435,08	2.303,71	---	740,97	1.462,92	---	1.912,46
2000	817,49	602,47	625,78	1.112,08	917,06	538,72	1.361,62	---	1.045,57
2001	557,52	461,68	374,65	737,61	373,61	321,33	835,88	---	741,24
2002	368,26	351,17	244,94	458,20	257,85	193,62	1.005,21	---	411,72
2003	778,43	681,22	414,73	1.069,73	349,77	455,47	1.359,22	---	723,25
2004	1.075,12	899,19	599,05	1.527,87	324,59	736,86	1.008,13	1.689,45	924,87
2005	1.726,23	1.351,41	784,87	2.725,36	593,24	999,77	1.235,73	1.294,14	1.710,04
2006	1.862,36	1.430,55	868,82	2.978,84	675,23	958,70	1.310,94	1.369,57	1.801,72
2006/Q1	1.862,36	1.430,55	868,82	2.978,84	675,23	958,70	1.310,94	1.369,57	1.801,72

Q: Quarter

BONDS AND BILLS MARKET

Traded Value

Outright Purchases and Sales Market

	Total		Daily Average	
	(YTL Million)	(US$ Million)	(YTL Million)	(US$ Million)
1991	1	312	0,01	2
1992	18	2.406	0,07	10
1993	123	10.728	0,50	44
1994	270	8.832	1	35
1995	740	16.509	3	66
1996	2.711	32.737	11	130
1997	5.504	35.472	22	141
1998	17.996	68.399	72	274
1999	35.430	83.842	143	338
2000	166.336	262.941	663	1.048
2001	39.777	37.297	158	149
2002	102.095	67.256	404	266
2003	213.098	144.422	852	578
2004	372.670	262.596	1.479	1.042
2005	480.723	359.371	1.893	1.415
2006	113.677	85.883	1.864	1.408
2006/Q1	113.677	85.883	1.864	1.408

Repo-Reverse Repo Market

Repo-Reverse Repo Market

	Total		Daily Average	
	(Y TL Million)	(US$ Million)	(Y TL Million)	(US$ Million)
1995	5.782	123.254	23	489
1996	18.340	221.405	73	879
1997	58.192	374.384	231	1.486
36160	97.278	372.201	389	1.489
1999	250.724	589.267	1.011	2.376
2000	554.121	886.732	2.208	3.533
2001	696.339	627.244	2.774	2.499
2002	736.426	480.725	2.911	1.900
2003	1.040.533	701.545	4.162	2.806
2004	1.551.410	1.090.477	6.156	4.327
2005	1.859.714	1.387.221	7.322	5.461
2006	509.928	384.999	8.359	6.311
2006/Q1	509.928	384.999	8.359	6.311

Q: Quarter

ISE GDS Price Indices (January 02, 2001 = 100)

YTL Based

	3 Months (91 Days)	6 Months (182 Days)	9 Months (273 Days)	12 Months (365 Days)	15 Months (456 Days)	General
2001	102,87	101,49	97,37	91,61	85,16	101,49
2002	105,69	106,91	104,87	100,57	95,00	104,62
2003	110,42	118,04	123,22	126,33	127,63	121,77
2004	112,03	121,24	127,86	132,22	134,48	122,70
2005	113,14	123,96	132,67	139,50	144,47	129,14
2006	113,29	124,28	133,14	140,12	145,22	127,71
2006/Q1	113,29	124,28	133,14	140,12	145,22	127,71

ISE GDS Performance Indices (January 02, 2001 = 100)

YTL Based

	3 Months (91 Days)	6 Months (182 Days)	9 Months (273 Days)	12 Months (365 Days)	15 Months (456 Days)
2001	195,18	179,24	190,48	159,05	150,00
2002	314,24	305,57	347,66	276,59	255,90
2003	450,50	457,60	558,19	438,13	464,98
2004	555,45	574,60	712,26	552,85	610,42
2005	644,37	670,54	839,82	665,76	735,10
2006	666,36	693,42	868,48	690,34	762,65
2006/Q1	666,36	693,42	868,48	690,34	762,65

ISE GDS Portfolio Performance Indices (December 31, 2003 = 100)

YTL Based

	Equal Weighted Indices (YTL Based)			Market Value Weighted Indices			
	EA180-	EA180-	EA GENERAL	PDA180-	PDA180+	PDA GENERAL	REPO
2004	125,81	130,40	128,11	125,91	130,25	128,09	118,86
2005	147,29	160,29	153,55	147,51	160,36	154,25	133,63
2006	152,41	166,19	159,05	152,67	166,25	159,80	137,51
2006/Q1	152,41	166,19	159,05	152,67	166,25	159,80	137,51

Q: Quarter