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Informational Efficiency in Developing Equity Markets

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Abstract

The issue of informational efficiency in the evolution of asset prices is examined using data on equity markets in Jordan, Turkey and Pakistan over the period 1986-93. The analysis is carried out in two steps. The parameters of agents' dynamic consumption and investment decisions are first estimated, and then the implied equity market price, based on market fundamentals, is compared with the actual evolution of equity market prices. While the informational efficiency of each of the three markets is found to be deficient, the causes of market inefficiency are varied. For Jordan it appears that a large negative shock to economic activity in the late 1980s caused agents to discount market fundamentals. For Turkey and Pakistan it is likely that institutional and legal rigidities in equity and banking markets resulted in these markets being illiquid, although this lack of market depth did reduce in severity for Turkey over the sample period, as liberalization of financial markets occurred.

JEL Classification Numbers:

C52, E21, G12

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Summary

The issue of informational efficiency in the evolution of asset prices is examined using data on equity markets in Jordan, Turkey, and Pakistan. These three markets are of interest because, for the sample periods analyzed, they represent a market that remained relatively open and where, in particular, no special regulations affected investment by nonresidents (Jordan during 1986-92); a market that underwent substantial liberalization, including the withdrawal of restrictions on foreign investors trading in listed equities (Turkey during 1988-93); and a market where restrictive financial market regulations remained in place, with foreign investors prohibited from trading in equities (Pakistan during 1986-91).

The analysis is carried out in two steps. First, the key determinants of agents' dynamic consumption and investment decisions (the coefficient of relative risk aversion and subjective discount rate) are estimated for each country. The values of these parameters, along with data on market fundamentals (current and future levels of aggregate private consumption and the stream of current and future dividends), are then used to construct an implied equity market price. This price is based, at any time t, on the assumption that consumers know the future time path of consumption, dividends, and the terminal price of the equities. Next, this implied price is compared with the actual evolution of equity market prices for each country to assess the informational efficiency of these markets by checking for the presence of excess volatility.

A key finding is that while the informational efficiency of each of the three markets is deficient, the causes of the inefficiency vary. While it appears that a large negative shock to economic activity in the late 1980s caused Jordanians to discount market fundamentals, in Turkey and Pakistan the illiquidity of equity markets influenced the determination of actual equity market prices. This result highlights the important role that financial market liberalization, and particularly the opening up of previously autarkic equity markets to foreign investors, can play in improving the efficiency with which equity markets mobilize resources for growth-enhancing investment.

I. Introduction

This paper undertakes an analysis of the informational efficiency of equity markets in Jordan, Turkey and Pakistan over the period 1986-93. In particular, we test for excess volatility in these three developing equity markets. Studying the information efficiency of equity prices is useful because theory suggests that these prices contain information about market expectations of future economic growth and interest rates; movements in equity prices can also have direct effects on consumption and investment expenditures via wealth and possible liquidity effects. The analysis is carried out in two steps. First we estimate a non-linear Euler equation from the consumption-based capital asset pricing model (C-CAPM) of Breeden (1979) and Grossman and Shiller (1981). Methods of estimating Euler equations were developed by Hansen (1982) and first applied to a C-CAPM by Hansen and Singleton (1982, 1984). The second step employs a form of Shiller's (1981) excess volatility test under the assumption of perfect foresight.

In a recent paper, El-Erian and Kumar (1994) study the equity markets of Jordan and Turkey in order to evaluate the informational efficiency of developing equity markets in the Middle East. 1/ Their study is important in that it provides a basis for exploring policies that would enhance the role of equity markets in stimulating investment and growth. However, their analysis is concerned with evaluating, using parametric and non-parametric techniques, the extent to which equity prices follow a random walk process. An alternative method, which is used here, is to analyze the issue of informational efficiency of stock markets by employing the C-CAPM as a baseline measure of available information. An advantage of our approach is that implications derived from economic theory can be tested directly by comparing them with actual outcomes.

An important contribution of this paper is to provide a methodology for statistically testing whether the estimated parameters from the C-CAPM are consistent with the efficient market hypothesis. To do this we take the estimated parameters and generate a time series of prices, P_t^*, called the ex-post rational asset price. This is the price at t under the assumption that consumers know the future time path of consumption, dividends and the terminal price, P_T. According to Shiller (1981) and Grossman and Shiller (1981), the variance of a certain ex-post asset price must be higher than the variance of the actual asset price, since the actual asset price is the consumer's optimal forecast (under rational expectations) of the ex-post asset price. When actual and ex-post asset prices follow random walks

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1/ See El-Erian and Tareq (1993) and Abisourour (1994) for analyses of major issues in economic reform in Arab countries, and in the development of emerging Arab capital markets, respectively. Feldman and Kumar (1994) also discuss more generally the role of equity markets in the development process.

(are integrated of order one), Shiller's variance ratio test for market efficiency is no longer valid since the unconditional variances do not exist. However, the procedure can be transformed by testing the hypothesis that ex-post and actual asset prices move together over time, that is, they are cointegrated with cointegrating vector (-1,1)--indicating no excess volatility.

A final issue analyzed here is whether the recent liberalization of financial markets in developing countries has affected the performance of these markets in terms of their informational efficiency. Our model provides a benchmark with which to assess the contribution of asset market liberalization to improving the functioning of equity markets, as it enables us to measure the extent to which markets use information efficiently to price equities. We expect that equity prices would be less influenced by regulatory distortions as impediments to market efficiency are gradually removed.

A key finding is that while the informational efficiency of each of the three markets is deficient, the causes of the inefficiency are varied. For Jordan it appears that a large negative shock to economic activity in the late 1980s caused agents to discount market fundamentals. For Turkey and Pakistan it is likely that institutional and legal rigidities in equity and banking markets resulted in these markets being illiquid, although this lack of market depth did reduce in severity for Turkey over the sample period.

The paper is organized as follows. Section II briefly reviews the development of equity markets in Jordan, Turkey and Pakistan. Section III sets out the C-CAPM and derives the test for excess volatility in equity markets. The data are described in Section IV, which is followed in Section V by the presentation and description of the estimates of the Euler equation for the C-CAPM, and the results of the excess volatility test. The cointegrating vectors of the excess volatility test will be estimated using Phillips and Hansen's (1990) fully modified (FM) cointegrating regression, and the hypothesis of no cointegration between ex-post and actual asset prices will be tested using the residual-based test of Phillips and Ouliaris (1990). The paper concludes with some reflections on the policy implications of our findings.

II. Equity Markets in Jordan, Turkey and Pakistan

Commercial banks have traditionally dominated the financial sectors of developing countries, with equity markets providing only a small share of total investible funds. This dominance of debt finance arose chiefly because of macroeconomic and regulatory policies which did not foster an economic environment in which the private sector was encouraged to play an active role in the economy. High fiscal deficits and inflation, overvalued exchange rates, the dominant role of the public sector, low (and often negative) real interest rates, and associated quantitative restrictions on the availability of credit were important constraints on the ability of the private sector in developing countries to contribute to national economic

growth. In addition, high taxes on dividends and capital gains, inadequate regulation and supervision of financial markets, the poor quality and dissemination of financial information, and barriers to inflows of foreign capital all contributed to a relatively weak demand for equity finance.

However, during the last decade many developing countries have recognized the attractive features of efficient equity markets in acting both as an instrument of portfolio diversification for savers and a relatively low-cost form of investment capital. Notwithstanding this recognition by policymakers, recent experience in developing equity markets indicates that appropriate macroeconomic policies are a necessary yet not a sufficient condition for financial markets to play a more important role in mobilizing resources for growth-enhancing investment. Policies which expand both the demand for, and supply of, equities are also of vital importance. These policies typically involve enhancing the flow of market information, improving the enforcement of private property rights, opening up markets to participation by non-residents, and strengthening institutions entrusted with the oversight of financial markets (El-Erian and Kumar 1994).

Such policies have recently been enacted in many developing countries, typically as part of a broad range of macroeconomic and structural adjustment policies designed to increase economic efficiency and enhance the pace and sustainability of the path of economic growth. Moreover, this process of liberalization of financial markets has also been prompted by developing countries themselves, which have become increasingly conscious of the important role for equity, bond and foreign exchange markets in both mobilizing and channelling (domestic and foreign) savings into productive investments (Feldman and Kumar 1994). 1/

The rapid increase in activity in developing equity markets since the mid-1980s has also occurred in Jordan, Turkey and Pakistan. An important influence on the process of financial intermediation in all three countries was the extent to which regulatory and institutional barriers to the flow of national and foreign savings into listed equities were lowered, and these country-specific experiences are briefly outlined below.

Relative to the size of its economy, the Jordanian equity market has traditionally been one of the most important in the Arab world. The Amman Stock Exchange (ASE), which was established in 1978, trades a wide variety of private equities and bonds, bank certificates of deposit and public bonds. The ASE also oversees regulations governing activity in Jordan's markets for equities and securities, and since 1986 the authorities have established an attractive institutional and regulatory environment for the operation of these markets. Individual investors (whether Jordanian or non-Jordanian) are exempted from income tax on all types of interest earned, and

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1/ Work by Atje and Jovanovic (1993) and Levine and Zervos (1995) indicates that financial services provided by equity markets are crucial factors underlying long-run economic growth.

capital gains, profits and dividends accruing from the holding of equities are tax free. There are no special regulatory impediments to foreigners trading on the ASE--Jordanian and non-Jordanian (including Arab and non-Arab) investors are treated equally; moreover, the repatriation of income from investment in equities by non-residents is free of restrictions (Abinsourour 1994, Toukan 1994).

Although securities have been traded in Turkey for over 120 years, activity in the Istanbul Stock Exchange (ISE) expanded greatly following the introduction of widespread financial and economic liberalization initiatives in the early 1980s. This process was also accelerated by the introduction of a privatization program in 1986, designed to divest the public sector of inefficient state agencies, and the subsequent trading of public sector shares on the ISE. Importantly, all restrictions on trading by foreign investors in listed Turkish equities were removed in 1989, and equity returns are free of withholding and capital gains taxes provided the foreign investors are registered with the proper public authorities (El-Erian and Kumar 1994).

The Karachi Stock Exchange (KSE), which was established in 1948, is the largest of the three equity markets in Pakistan (Karachi, Islamabad and Lahore), and accounts for about 90 percent of all Pakistani equities traded. As of mid-1993 there were some 630 companies listed on the KSE, having increased from 361 at end-1986. However, in contrast to Jordan and Turkey, the Pakistani equity market has remained relatively small and illiquid, due in large part to weak regulation and supervision of asset markets, and policies of financial repression which have dampened both the demand for, and supply of, traded equities. Trading on the KSE is confined to long positions in equities; short selling of equities and derivatives is not allowed. In addition, until 1992 the only mutual funds that were permitted to trade on the KSE were those administered by government-controlled agencies, mainly industrial and financial institutions providing long-term development finance and mutual fund services in Pakistan (Haque and Kardar 1993).

Key statistics relating to each of the three equity markets are presented in Table 1. They illustrate the traditionally greater relative importance of equity markets in Jordan, which compares favorably with the capitalization to GDP ratio of industrial equity markets in the United States, Japan and the United Kingdom. Expressed in U.S. dollar terms, equity markets in Turkey and Pakistan have increased in value by over thirty-fold between 1988-93 and by over four-fold between 1986-91, respectively. The spectacular growth in the value of equities traded on the ISE between 1988-93 also stands out, especially in comparison with Jordan (high capitalization) and Pakistan (low capitalization), where relatively little trading in equities occurs (Table 1). Table 2 reveals for the period 1988-93 the extent of cross-country correlation in movements in equity market prices for Jordan, Turkey, Pakistan, and equity markets in selected industrial countries (the United States, the United Kingdom and Japan). While Jordan exhibits positive correlation with all three

Table 1: Summary Statistics of Equity Markets in Jordan, Turkey, Pakistan and Selected Industrial Countries 1/

Country and Year	Market Capitalization: Millions of U.S. Dollars	Market Capitalization: Percent of GDP	No. of Listed Companies 2/	Value Traded (Millions of U.S. Dollars)	IFC Global Index: Price Level	IFC Global Index: Number of Stocks	IFC Global Index: % of Market Capitalization
Jordan (1986)	2,839	48.7	103	185	135.5 4/	10	44.3
Jordan (1992)	3,365	68.9	103	1,317	129.1 4/	27	59.1
Turkey (1988)	1,135	1.6 3/	50	101	123.5 5/	14	47.2
Turkey (1993)	37,496	29.8 3/	152	23,242	477.1 5/	36	63.3
Pakistan (1986)	1,710	5.5	361	155	122.3 4/	52	38.0
Pakistan (1991)	7,326	17.6	542	620	354.2 4/	54	53.6
United States (1986)	2,636,598	61.8	8,403	1,795,998	144.8 6/	..	..
United States (1993)	5,223,768	81.9	7,607	3,507,223	280.2 6/	..	..
United Kingdom (1986)	439,500	77.5	2,106	132,912	173.7 7/	..	..
United Kingdom (1993)	1,151,646	124.1	1,646	423,526	354.9 7/	..	..
Japan (1986)	1,841,785	87.6	1,866	1,145,615	237.2 8/	..	..
Japan (1993)	2,999,756	71.6	2,155	954,341	348.2 8/	..	..

Sources: IFC (1994) and International Monetary Fund, International Financial Statistics, various issues.

Notes: The IFC Global indexes are intended to represent the performance of the most active stocks in their respective stock markets, and to be the broadest possible indicator of market movements.

1/ End-of-period levels.

2/ Amman Financial Market for Jordan; Karachi Stock Exchange Ltd. for Pakistan; Istanbul Stock Exchange for Turkey; combined NASDAQ, New York and American Stock Exchange for the United States; London Stock Exchange for the United Kingdom; and combined Fukuoka, Hiroshima, Kyoto, Nagoya, Niigata, Osaka, Tokyo and Sapporo Stock Exchanges for Japan.

3/ Percent of GNP for Turkey.

4/ Base December 1984 = 100.

5/ Base December 1986 = 100.

6/ Standard and Poor's 500 Index, base 1984 = 100.

7/ Financial Times 100 Index, base 1984 = 100.

8/ Nikkei Index, base 1984 = 100.

Table 2. Jordan, Turkey, Pakistan and Selected Industrial Countries--Correlation among Stock Prices 1/
(December 1988 - December 1993)

	United States	United Kingdom	Japan	Jordan	Turkey	Pakistan
United States	1.00					
United Kingdom	0.59	1.00				
Japan	0.29	0.49	1.00			
Jordan (0.3)	0.22	0.27	0.21	1.00		
Turkey (2.4)	-0.23	-0.07	0.07	-0.07	1.00	
Pakistan (1.0)	0.04	0.17	-0.13	0.10	0.04	1.00

Source: Based on International Finance Corporation's (1994) Price Indexes, in U.S. dollar terms.

1/ Correlations are for five years of monthly returns beginning December 1988. Industrial country indexes (in U.S. dollar terms) are the Standard and Poor's 500 (for the United States); the Financial Times 100 (for the United Kingdom) and the Nikkei (for Japan).

Numbers in parentheses underneath country names show the percentage share of developing market capitalization (in the IFC's Composite Index of emerging markets) accounted for by each country at end-1993.

industrial markets, there is either negative or only slightly positive correlation between Turkey and the industrial markets and between Pakistan and the industrial markets (apart from its positive correlation with the United Kingdom). This suggests that investors in industrial countries may benefit from portfolio diversification in Turkish and Pakistani equity markets. Finally, there is little evidence of correlation in movements in equity prices across the three developing equity markets.

III. C-CAPM and Tests for Excess Volatility

This section sets out the C-CAPM, which is used to construct equity prices based on factors taken from economic theory. It then derives the test for excess volatility to be used in measuring informational efficiency in developing equity markets.

Consider an economy composed of a large number of similar, infinitely lived consumers, each maximizing:

$$E_t \sum_{i=0}^{\infty} \beta^i U(c_{t+i}). \tag{1}$$

Here c_t is the agent's consumption at time t, and β is the subjective discount factor ($0<\beta<1$) that reflects preference for current consumption over future consumption. 1/ Consumers have the choice of saving in any of N risky assets. Assume also that consumers are not interested in the decision-making aspects of production, and so the results of the production side of the economy appear as stochastic returns to the investments of consumers. Let x_t be the N-vector representing the quantity of all assets that the consumer holds between t and t+1. Furthermore, let P_t and d_t be N-vectors of prices and dividends of the N assets at time t.

The consumers' budget constraint can be written as:

$$c_t + P_t' x_t = (P_t + d_t)' x_{t-1}. \tag{2}$$

The Euler equations from maximizing (1) subject to (2) are:

1/ As $\beta \to 0$ then consumers are more and more indifferent to the time path of consumption; as $\beta \to 1$ then consumers exhibit stronger preferences for present consumption over future consumption.

$$E_t\left[\beta\frac{U'(c_{t+1})}{U'(c_t)}R_{it}\right]=1,\quad i=1,\ldots,N, \tag{3}$$

where $R_{it}=(P_{i,t+1}+d_{i,t+1})/P_{it}$ is the return on the asset and $\beta U'(c_{t+1})/U'(c_t)$ is the marginal rate of substitution between current and future consumption. An interpretation of equation (3) is that the weighted expectation of returns is the same for all assets, where the weights correspond to the marginal rates of substitution. The intuition is that returns which are delivered in periods of low marginal utility of consumption (high levels of consumption) have low weights because they provide little additional utility. The opposite is true for returns delivered in periods of high marginal utility (low levels of consumption), which have high weights because they provide a great deal of additional utility.

To estimate the model we need to make some assumptions on the form of the utility function, $U(\cdot)$. We shall work with a utility function of the constant relative risk aversion (CRRA) type, that is:

$$U(c)=\begin{cases}\dfrac{c^{1-\lambda}}{1-\lambda}, & if\ \lambda\neq 1\\ \log(c), & if\ \lambda=1,\end{cases} \tag{4}$$

where λ is the coefficient of relative risk aversion, which measures the concavity of the utility function. Using (4) in (1) yields the Euler equations:

$$E_t\left[\beta\left(\frac{c_{t+1}}{c_t}\right)^{-\lambda}R_{it}\right]-1=0,\quad i=1,\ldots,N. \tag{3'}$$

By iterating (3'), we find that the price of asset i at time $t<T$ is the expected present value of dividends and a terminal price P_{iT}, and is given by:

$$P_{it} = E_t\left[\sum_{j=1}^{T-t} \beta^j \left(\frac{c_{t+j}}{c_t}\right)^{-\lambda} d_{i,t+j} + \beta^{T-t}\left(\frac{c_T}{c_t}\right)^{-\lambda} P_{iT}\right], \tag{5}$$

where asset prices are determined by market fundamentals, which are assumed to be: current and future levels of economic activity (represented by movements in aggregate private consumption); and the stream of current and future dividends. It is also assumed that there are no liquidity premia arising from country-specific risk factors, which would then raise the price of equities above that determined solely by market fundamentals. The stream of dividends are discounted by both the subjective discount factor and changes in consumption patterns. That is, the overall discount factor varies with economic activity, and the larger the value of λ the greater is the variability of the discount factor.

The perfect foresight asset price, P^*_{it} is defined such that $P_{it}=E_t[P^*_{it}]$, that is P^*_{it} is the term inside the conditional expectation operator. Given that we have parameter estimates for λ and β, we can construct an estimate of the perfect foresight asset price using (5). Shiller (1981) shows that under rational expectations,

$$Var(P^*_{it}) \geq Var(P_{it}). \tag{6}$$

Defining the rational consumer's forecast errors to be $u_{it}=P^*_{it}-P_{it}$, we can transform (6) as follows:

$$\frac{Cov(P^*_{it},P_{it})}{Var(P_{it})} - 1 \geq \frac{-Var(u_{it})}{2Var(P_{it})}. \tag{7}$$

If P^*_{it} and P_{it} are integrated of order one (denoted $I(1)$) and the market is efficient, then u_{it} is weakly stationary and satisfies $Eu_{it} = 0$ and $Cov(R_{it},u_{it}) = 0$, otherwise u_{it} would contain unexploited information. Since u_{it} is a stationary linear combination of P^*_{it} and P_{it}, then by definition P^*_{it} and P_{it} are cointegrated of order (1,1). Hence, the conditions of market efficiency imply that the right-hand side of (7) will approach zero as the sample size increases, and that the first term on the left-hand side of (7), which is the slope of the coefficient from a regression of P^*_{it} on P_{it} (denoted henceforth by γ), must approach its upper bound of one. Consequently, the correct excess volatility test when P^*_{it} and P_{it} are $I(1)$ is to test whether the cointegrating vector is (-1,1), that is,

whether $\gamma = 1$; a necessary condition for this test to be valid is that P^*_{it} and P_{it} are cointegrated.

IV. The Data

Quarterly stock market data for Jordan (1986:1 to 1992:4), Turkey (1988:1 to 1993:4) and Pakistan (1986:1 to 1991:4) were taken from the International Finance Corporation's (IFC) Emerging Markets Database. For each country the following series were obtained from the IFC's global (IFCG) index: (i) the local stock exchange index; and (ii) the dividend yield corresponding to the stock price index. The IFCG index is constructed to take into account movements in domestic equity prices, regardless of the restrictions on foreign investors. Since the establishment of the Amman Stock Exchange in 1978 there have been no special restrictions affecting foreign investors in Jordan; such restrictions were relaxed considerably in Turkey in 1989, and in Pakistan in 1992. Data on the coverage of the index for these three countries is given in Table 1. Quarterly data on aggregate private consumption were interpolated from annual International Financial Statistics (IFS) data (line 96f), using a polynomial interpolation method. IFS data on the GDP deflator (line 99bi) and population (line 99z) were used to derive real (1990) per capita aggregate consumption for each of the three countries, matching the periods covered by the IFC data.

V. Empirical Results

The results of our model estimation and hypothesis tests are presented in this section. First, the parameters of agents' dynamic consumption and investment decisions are estimated. Second, the implied equity market price, based on market fundamentals and the abovementioned parameters, is compared with the actual evolution of equity market prices. This comparison is made to determine whether the markets can be deemed to display informational efficiency, through a test for the presence of excess volatility.

An innovation of this paper is the use of a two-stage test for informational efficiency. The first stage is a variant of the excess volatility test typically used in the literature, which examines the significance of the parameter on the actual market price, P_t, in a regression of an equity market price based on market fundamentals, P^*_t, on P_t. The second stage examines whether P^*_t and P_t are cointegrated--whether they move together in response to shocks--which would again be evidence in favor of informational efficiency. 1/ We find evidence that all three equity markets suffer from informational inefficiencies. To determine the cause of this inefficiency, the assumption of no liquidity premia (that is,

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1/ While the cointegrated variables are linked by a long-run equilibrium relationship they may diverge from equilibrium in the short run. However, in the long run economic forces are expected to act so as to restore equilibrium.

an absence of country-specific risk) implicit in the above cointegration test is then relaxed. This enables us to gauge whether the financial liberalization which has occurred in equity markets in Jordan, Turkey and Pakistan since the mid-1980s has contributed to enhancing the informational efficiency of these markets.

1. Generalized method of moments (GMM) estimation

To estimate the parameters of equation (3') we use the GMM procedure. 1/ An important issue in GMM estimation is the choice of an optimal weighting matrix for efficient estimation. The optimal weighting matrix is given by the inverse of the asymptotic covariance matrix. To estimate the covariance matrix a kernel must be chosen, to ensure that the estimate is positive semi-definite, and the bandwidth parameter selected, to control the amount of serial dependence allowed in the estimate. For empirical work it is useful to have a means of choosing the bandwidth in an automatic way, so that comparisons between different studies can be made on the basis of a standardized method. In this paper we will use the automatic bandwidth selection procedure of Andrews (1991) and the Bartlett kernel, which are suitable for use in covariance matrix estimation. 2/

The estimation requires a set of instrumental variables, and those employed here are $Z_t=(1,g_{t-1},R_{t-1},\ldots,g_{t-\rho},R_{t-\rho})'$, where $g_t=c_{t+1}/c_t$ and R_t is defined as in equation (3). A range of lag lengths (ρ) were chosen to examine the robustness of the results (labelled 'instrument lags' in Table 3). In addition to the parameter estimates, an asymptotic chi-squared (χ^2) test of the over-identifying restrictions implied by the econometric model is provided, together with the degrees of freedom and significance level (sig) of the test. 3/ When $\lambda = 0$ in equation (4), this implies that the rate of return of an asset will be constant and equal to the subjective rate of time preference. However, when $\lambda \neq 0$, this implies that the rate of return may vary with consumption; a value of λ close to one indicates that consumers have a log utility function.

1/ Details of implementing the GMM estimator can be found in Hall (1993), Ogaki (1993) and Hamilton (1994).

2/ The Bartlett kernel takes the form $k(u)=(1-|u|).1(|u|\leq 1)$, where $1(\cdot)$ is the unit indicator function.

3/ In the GMM framework, to estimate k parameters we need at least k orthogonality conditions. When the number of orthogonality conditions exceeds the number of parameters, the model is overidentified. When the instruments Z_t are used, $2\rho+1$ orthogonality conditions exist. These orthogonality conditions are given by setting the population moments $E_t\ [(\beta R_t(1/g_t)^{\lambda} - 1)\ Z_t]$ to zero. The test of over-identifying restrictions determines whether the sample analogues of the population moments are close to zero, as would be expected. The degrees of freedom of the test are given by $2\rho+1-k$.

Table 3. GMM Estimation of the Subjective Discount Factor and Coefficient of Relative Risk Aversion: Jordan, Turkey and Pakistan

Country	Instrument lags (ρ)	λ	se(λ)	β	se(β)	χ^2	sig	AIC
Jordan	1	0.276	0.392	0.943	0.012*	2.086	0.149	-5.712
	2	0.274	0.423	0.943	0.012*	2.405	0.493	-5.764
	3	0.449	0.561	0.932	0.012*	4.387	0.495	-5.792
	4	0.429	0.339	0.930	0.011*	4.407	0.732	-6.109
	5	0.438	0.340	0.930	0.010*	4.486	0.877	-5.711
Turkey	1	-9.924	7.174	0.802	0.098*	0.100	0.752	-1.352
	2	-5.371	2.349*	0.822	0.052*	0.669	0.881	-2.000
	3	-6.887	3.110*	0.811	0.058*	1.988	0.851	-1.790
	4	-1.350	2.783	0.766	0.047*	3.484	0.837	-2.444
	5	-1.534	2.088	0.763	0.039*	3.517	0.940	-2.378
Pakistan	1	5.869	2.952*	0.905	0.018*	3.629	0.057	-3.325
	2	4.267	2.080*	0.910	0.023*	4.091	0.252	-3.282
	3	2.921	1.073*	0.913	0.022*	4.157	0.527	-3.200
	4	3.195	1.205*	0.921	0.023*	3.509	0.834	-3.193
	5	2.773	1.053*	0.921	0.025*	3.362	0.948	-3.151

Sources: International Finance Corporation (1994) and earlier issues; International Monetary Fund, International Financial Statistics, various issues.

Notes: The periods of analysis (using quarterly data) range between 1986:1 to 1992:4 (Jordan); between 1988:1 to 1993:4 (Turkey); and between 1986:1 to 1991:4 (Pakistan). The coefficient of relative risk aversion is denoted by λ, as given in equation (4). The subjective discount factor is denoted by β, as given in equation (1). The degrees of freedom for the χ^2 test of the over-identifying restrictions are 1, 3, 5, 7 and 9, for $\rho = 1$, 2, 3, 4 and 5, respectively. The significance level of the over-identifying restriction test is denoted by "sig". AIC denotes the Akaike Information Criterion, which determines the optimal lag length by minimizing the sum of the residual sum of squares plus a penalty term, where the penalty term increases with the number of lags. For all three countries the optimal lag length was found to be four.

* Indicates that the parameter is significantly different from zero at the 5 percent level.

2. Estimated parameters of the consumption Euler equation

Table 3 displays the parameter estimates of equation (3'). The estimated values of the relative risk aversion parameter (λ) for Jordan range from 0.274 to 0.449 for different instruments, but are all insignificant. The estimated values of λ for Turkey are all negative and are mostly insignificant. Values of $\lambda < 0$ are not economically sensible, since this would imply that consumers prefer a variable consumption path in aggregate. The estimated values of λ for Pakistan range from 2.773 to 5.869 for different instruments, and are always significant at the five percent level. The estimated values for the measure of impatience (β) are much more robust across instruments and countries than the estimated values of λ, with β ranging from 0.763 for relatively impatient Turkey to 0.943 for relatively patient Jordan. A test of the validity of the over-identifying restrictions used in our estimation was not rejected for any country at the five percent level of significance.

The Akaike Information Criterion (AIC) was used to determine which lag length is optimal, in the sense that it minimizes the sum of the residual sum of squares plus a penalty term, where the penalty term increases with the number of lags. For all three countries the optimal lag length was found to be four, and the results for other lag lengths are included as a check of the robustness of the AIC results (Table 3).

At a lag length of four the values of λ and β are: 0.429 (se = 0.339) and 0.930 (se = 0.011) for Jordan; -1.350 (se = 2.783) and 0.766 (se = 0.047) for Turkey; and 3.195 (1.205) and 0.921 (0.023) for Pakistan (Table 3). The parameter estimates of λ for Jordan imply a utility function with a low level of relative risk aversion (approximating the shape of the square root function). For Pakistan the implied utility function is more concave, meaning that Pakistan is more sensitive to varying returns. That is, the weights of the marginal rates of substitution affect the value of returns more in Pakistan than in Jordan. The estimated parameter of λ for Turkey is negative. As mentioned above this is not consistent with economic theory. Given that the estimate is not significantly different from zero we will assume that it is zero. Accordingly, this implies that the expected returns on all assets in Turkey must be equal in all time periods, irrespective of whether the relevant period is characterized by a high level of consumption or a low level of consumption.

It should also be noted that the interpolation method used to construct the data on aggregate private consumption will most likely introduce a degree of mismeasurement into the data, and so may consequently distort the correlations between consumption and returns. This in turn may explain why λ in Table 3 is estimated with large standard errors, or even why λ lies outside the admissible (concave) region of the parameter space in some instances (such as for Turkey).

3. Testing for excess volatility

The correct use of Shiller's excess volatility test when stock prices are $I(1)$ is to test for cointegration and also to test that the cointegrating vector is (-1,1), that is, whether $\gamma = 1$ (there is no excess volatility). The cointegrating regression can be estimated using Phillips and Hansen's (1990) FM method, which yields an asymptotically correct variance-covariance estimator when estimating cointegrating vectors in the presence of serial correlation and endogeneity. Hence, the hypothesis $\gamma = 1$ can be tested with standard asymptotic tests when the residual variance estimator is replaced with the 'long run' residual variance estimator. To apply Phillips and Hansen's (1990) method here, we regress a transformed p_t^* on p_t (where p_t^* and p_t are both logarithms of P_t^* and P_t). See Appendix A1 for details of the FM cointegrating estimator. The hypothesis of no cointegration can be tested using the residual-based test of Phillips and Ouliaris (1990). See Appendix A2 for details of the Phillips-Ouliaris test statistics.

It should be noted that ordinary least squares (OLS) estimation could be used to yield a consistent estimate of γ. However, OLS estimation is inefficient and yields non-standard distributions of the estimators. Consequently, standard tests of linear restrictions cannot be used in the OLS framework, while these problems are overcome in the FM method.

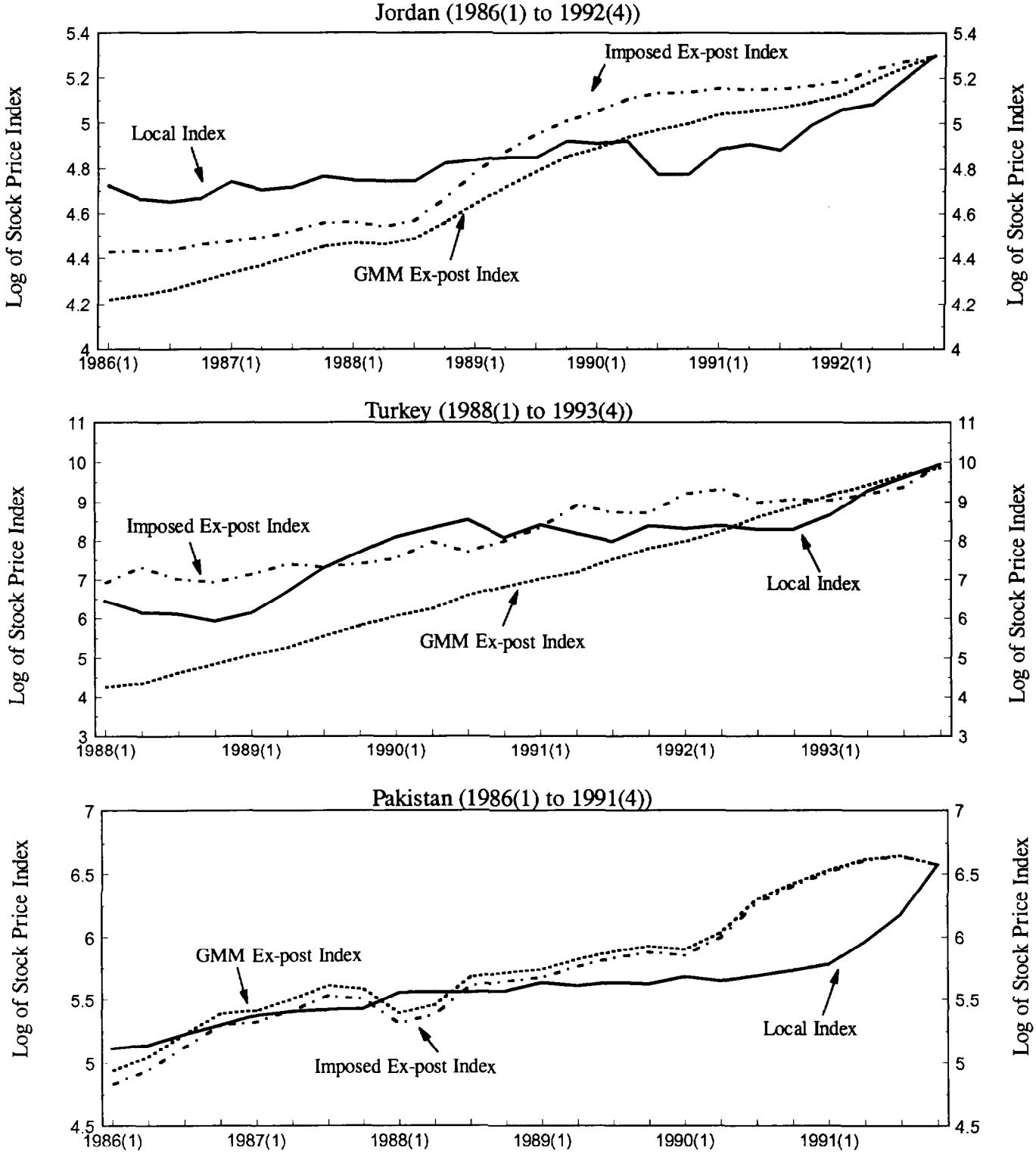
Table 4 displays the results of the test for informational efficiency. In addition to applying this test to the ex-post series generated using the GMM estimated Euler equation parameters (using up to five lagged instrumental variables), the test is also applied to an ex-post series generated when $\beta = 0.9$ and $\lambda = 6.0$ (denoted as 'Imposed ex post'). Parameter values of β were measured with a high degree of precision; a value of $\beta = 0.9$ was imposed as it was representative of the values reported in Table 3.

The value of $\lambda = 6.0$ was imposed because in this case p_t^* and p_t have very similar movements and thus have the greatest chance of passing the excess volatility test. The larger the value of λ the more variable is the discount factor, and thus the more variable is p_t^*. The estimated values of λ reported in Table 3 lack precision, and one important rationale for this result could be that agents are more risk averse than indicated in our estimates. To the extent that there is varying uncertainty in any given economy that affects returns to investment, rational agents will see the value in waiting until some of this uncertainty is resolved, but this will not be directly reflected in our model (Dixit and Pindyck 1994). Instead, we impose a higher value of λ (the parameter reflecting the extent of risk aversion) to indirectly account for this phenomenon, by increasing the variability of the discount factor.

Figure 1 shows for each country the movements in the actual stock price index (p_t) and the two ex-post stock price indexes (p_t^*): 'GMM ex post' and 'Imposed ex post'. For Turkey the imposed ex-post series for p_t^* is closely correlated with p_t, while the GMM ex-post series is not. Both measures of

Figure 1

Actual and Ex-Post Stock Prices - Jordan, Turkey and Pakistan



Sources: International Finance Corporation (1994) and earlier issues; Authors' calculations.

Table 4. Excess Volatility Tests Using Cointegration Methods: Jordan, Turkey and Pakistan 1/

Country	Lags	γ	se(γ)	Z(α)
Jordan	1	1.005	0.012	-3.836
	2	1.005	0.012	-3.806
	3	0.969	0.018**	-2.206
	4	0.965	0.019**	-2.100
	5	0.966	0.019**	-2.130
	Imposed	0.991	0.017	-2.512
Turkey	1	1.044	0.024	-10.059
	2	1.037	0.023	-8.287
	3	1.033	0.024	-8.792
	4	0.912	0.051**	-1.878
	5	0.910	0.052**	-1.846
	Imposed	1.050	0.023	-5.475
Pakistan	1	0.864	0.008*	-0.381
	2	0.879	0.010*	1.213
	3	0.846	0.007*	-2.170
	4	0.879	0.010*	1.213
	5	0.889	0.010*	1.278
	Imposed	1.021	0.020	-5.638

Sources: International Finance Corporation (1994) and earlier issues; International Monetary Fund, International Financial Statistics, various issues.

Notes: "Lags" refers to the number of lags used as instruments. "Imposed" refers to the excess volatility test when the parameters are set to $\beta = 0.9$ and $\lambda = 6.0$, which yield an approximate correspondence between p_t^* and p_t, as explained in Section V. The periods of analysis (using quarterly data) range between 1986:1 to 1992:4 (Jordan); between 1988:1 to 1993:4 (Turkey); and between 1986:1 to 1991:4 (Pakistan).

* Indicates that γ is significantly less than unity (the null hypothesis of no excess volatility is rejected) at the 2.5 percent level.

** Indicates that γ is significantly less than unity (the null hypothesis of no excess volatility is rejected) at the 5 percent level.

1/ The 2.5 percent and 5 percent critical values for the Z(α) test are -18.883 and -15.638, respectively (Phillips and Ouliaris 1990). Computed values for the Z(α) statistic smaller (more negative) than these critical values indicate that the hypothesis of no cointegration is rejected, which is evidence in favor of informational efficiency. For the cointegration tests to be valid it is essential that both P_t^* and P_t are integrated of order one, I(1). This property of the prices can be ascertained from Figure 1. The results of formal unit root tests confirm that P_t^* and P_t are I(1) variables, and are available from the authors on request.

p_t^* are oversmoothed relative to the actual series for p_t for Jordan. The Pakistani data reveals that there is significant divergence between p_t and both measures of p_t^*, where the latter are biased upwards.

A test (reported in Table 4) of the null hypothesis of no excess volatility ($\gamma = 1$) against the alternative of excess volatility ($\gamma < 1$) finds that the null is rejected in all five cases for Pakistan (at the 2.5 percent level); in two of the five estimated cases for Turkey (lags = 4 and 5, at the 5 percent level); and in three of the five estimated cases for Jordan (lags = 3,4 and 5, at the 5 percent level of significance).

A value of the $Z(\alpha)$ statistic less than the 5% critical value of -15.638 (indicating a rejection of the no cointegration hypothesis with respect to p_t^* and p_t) is not reported for any country (Table 4). This implies that the difference between p_t^* and p_t contains a stochastic trend, which could be driven by market sentiment unrelated to fundamentals. Even when we used the imposed ex-post value as our measure for p_t^*, the stochastic trend remained for all three countries, although for Turkey it appears to make the imposed ex-post series for p_t^* more closely approximate the actual evolution of p_t (Figure 1).

Our key findings can be summarized as follows. First, the hypothesis of no excess volatility ($\gamma = 1$) is rejected for all lag lengths for Pakistan, and for several lag lengths for Jordan and Turkey. However, when values of β and λ are imposed which maximize the chance that the hypothesis of no excess volatility will not be rejected, it is indeed not rejected for any of the three countries. Second, the hypothesis of no cointegration between p_t^* and p_t (that both price series do not move together) is not rejected for any of the three countries, even when using the imposed values of the key parameters of the consumption Euler equation. Accordingly, we find that conditional on our model of asset price determination, all three equity markets suffer from informational inefficiency.

4. Testing for liquidity premia

An obvious question at this juncture is whether the omission of liquidity premia (an assumption implicit in equation (5)) might explain our inability to find that p_t^* and p_t are cointegrated. That is, it could be that our previous omission of country-specific risk as a determinant of equity prices is biasing our findings towards rejection of the hypothesis of no cointegration between p_t^* and p_t. To induce such a rejection, the nature of the liquidity premium would need to include a stochastic trend, indicating that it could move around without bound. Accordingly, we use Hansen's (1992) stability tests to examine whether the failure of cointegration was due to parameter instability arising from our omission of liquidity premia. In particular, we will examine whether there is an unstable liquidity premium, which can be represented by a constant term in the cointegrating regression of p_t^* on p_t.

To examine the hypothesis of parameter stability in the context of FM estimation of a cointegrating regression we will use the three tests suggested by Hansen (1992). The reason why we use a parameter stability test with a change point of unknown timing is to examine whether the many exogenous shocks and policy changes (for example, a revision of investment laws or the opening up of financial markets to foreigners) that significantly affect small economies such as Jordan, Turkey, and Pakistan are the potential cause of a structural change in the relationship between stock prices and their fundamentals. The timing of such a structural change is likely to be unknown, because there is not necessarily a one-to-one correspondence between potential causes of a structural change and its occurrence. A structural change can occur after an exogenous shock or policy change because it may take time for agents to learn about the new economic environment. Alternatively, a structural change can occur prior to an exogenous shock or policy change if agents' anticipate the economic consequences of a policy change.

The null hypothesis of the Hansen tests is that the parameters are constant, while the alternative hypothesis is that the parameters change at some unknown fraction of the sample. Let $\Pi \subset (0,1]$ be the set of all possible breakpoints searched over, where in this case $\Pi = [0.15, 0.85]$. When no information regarding the location of a change point π is available a natural choice of the possible set of change points Π is (0,1), where $\pi=0$ indicates that the change point is at the beginning of the sample, $\pi=1$ indicates that the change point is at the end of the sample, while values of π between zero and one indicate the fraction of the sample before the change point. Testing whether a sample has a change point at its extremities is not desirable because statistics tend to diverge to infinity since they cannot discriminate between true change points and boundary conditions. To overcome this problem the standard practice is to choose a restricted interval such as [0.15, 0.85]. That is, for our three samples the first five and last five observations are excluded from being possible breakpoints. See Appendix A3 for definitions of the test statistics used in Hansen's (1992) stability tests.

The results from the Hansen stability tests are as follows. For Jordan, while parameter stability is found using the Sup F test, this finding is not robust since the Lc and Mean F test results clearly reject parameter stability. For Turkey, while there is some evidence of parameter stability, this conclusion is not robust. In particular, at the optimal lag length of four (determined previously using the AIC), parameter stability is rejected by the Mean F and Sup F tests. For Pakistan, all three test statistics exceed their critical values, indicating parameter instability (Table 5). The key conclusion of these tests is that the failure to reject the hypothesis of no cointegration is indeed due to parameter instability, most likely arising from the omission of liquidity premia in our cointegration test. Our interpretation of these results can also be seen in the plots of p_t^* against p_t in Figure 1, which reveal that for Turkey a regime shift resulted in a closer relationship between the two

Table 5. Hansen Stability Tests: Jordan, Turkey and Pakistan 1/

Country	Lags	Lc	Mean F	Sup F
Jordan	1	0.800	7.601	9.939
	2	0.798	7.584	9.917
	3	0.777	7.407	9.949
	4	0.765	7.304	9.813
	5	0.770	7.341	9.864
	Imposed	1.164	11.196	15.213
Turkey	1	0.256	2.705	10.245
	2	0.261	3.368	17.036
	3	0.261	3.093	14.454
	4	0.302	4.937	28.674
	5	0.302	4.920	28.533
	Imposed	0.224	2.872	10.588
Pakistan	1	5.251	50.385	98.273
	2	5.176	47.056	92.451
	3	5.228	43.957	86.860
	4	5.165	45.497	89.430
	5	5.253	44.379	87.404
	Imposed	5.750	49.783	97.415

Sources: International Finance Corporation (1994) and earlier issues; International Monetary Fund, International Financial Statistics, various issues.

Notes: "Lags" refers to the number of lags used as instruments. "Imposed" refers to the value of the stability test when the parameters are set to $\beta = 0.9$ and $\lambda = 6.0$, which yield an approximate correspondence between p_t^* and p_t, as explained in Section V. The periods of analysis (using quarterly data) range between 1986:1 to 1992:4 (Jordan); between 1988:1 to 1993:4 (Turkey); and between 1986:1 to 1991:4 (Pakistan).

1/ The critical values for the Lc test, the Mean F test, and the Sup F test are 0.58, 4.47, and 12.40, respectively (Hansen 1992). Computed values for the stability tests greater than these critical values indicate parameter instability.

price series towards the end of the sample period; while in the case of Pakistan, the opposite occurred.

The above results echo El-Erian and Kumar (1994) in indicating that equity markets in all three countries are not solely driven by market fundamentals, and do exhibit informational inefficiency. These problems remained, except for Turkey, even when we imposed a higher value of λ (which reflects agents' aversion to uncertainty) to allow movements in p_t^* to be more sensitive to the fundamentals. Accordingly, various country-specific effects appear to underlie our findings of informational inefficiency in each of the three equity markets. 1/

We find informational inefficiency in Jordan chiefly because the net present value of the return on Jordanian stocks did not decline during 1989-90, whereas the actual price of equities in the Jordanian market did decline over this period. Market sentiment, in light of a large negative shock to real GDP of over 20 percent in 1989, appears to have caused the economic fundamentals to be discounted. 2/ The results for Turkey indicate that despite the gradual deregulation of financial and equity markets since the mid-1980s, there may still be a mismatch between actual equity market prices and the corresponding price series which might be expected from the market fundamentals. Interestingly, the relaxation of restrictions on equity market investment by foreigners in 1989 corresponded with an apparent regime shift, which increased the liquidity of the market and reduced the extent of informational inefficiency.

We were unable to reconcile the actual evolution of Pakistani equity market prices with the evolution of the price series implied by market fundamentals. Accordingly, our finding of excess volatility in Pakistan most likely arises from the market's small and illiquid nature. Key explanations for the market's illiquid nature include: weak regulation and supervision of the market; policies of financial repression which encourage sponsors/owners of companies to lock away a large share of the equity; the exclusion of foreign investors; and public institutional investors which assist in the financing of projects by favored individual promoters (Haque and Kardar 1993). These distortions created a divergence between the ex-post price and the actual price, which appears to have begun in 1991

1/ Informational inefficiencies could also arise from common factors, in addition to country specific causes. Illiquid trading in all three equity markets provides limited (if not ill-signaled) information about current inflation expectations and real rates of return on all assets. As a consequence, the influence of market fundamentals could be distorted, reflecting valuation problems in actual stock prices. In addition, it can be difficult to assess risk/return prospects in economies where large-scale structural and financial reforms are taking place.

2/ Such a finding could also arise to the extent that, as noted above, private consumption is an inadequate proxy for economic activity in our model of asset price determination (equation (5)).

(Figure 1). This divergence biased upward our measure of p_T and, in turn, induced an inflated measure of p_t^* (and an apparent regime shift) from 1989 onwards.

Our results for the developing equity markets of Jordan, Turkey and Pakistan can be compared with those of McDermott (1994), who applied the testing strategy outlined in this paper to annual U.S. data on the Standard and Poor's stock price index for the period 1889-1985. Cointegration between actual stock prices and ex-post stock prices was found only when: (i) the ex-post prices were constructed from the imposed parameters β=0.85 and λ=4; and (ii) the GMM estimated parameters with five lags were used as instrumental variables (one to six lags were tried). McDermott (1994) found only very weak evidence in favor of informational efficiency. However, this conclusion was not robust to the use of different instrumental variables. In this paper we reject this hypothesis for Jordan, Turkey and Pakistan. On the other hand, when a priori sensible parameter values for β and λ were imposed, then a finding of informational efficiency was readily attained for the United States. This result contrasts with our findings for Jordan, Turkey and Pakistan, where the imposition of such sensible parameters still resulted in a finding of informational inefficiency for each of the equity markets.

VI. Conclusion

In this paper we have examined the issue of informational efficiency in three developing equity markets: Jordan, Turkey and Pakistan. These three markets are of interest, as for the sample periods analyzed they include: a market which remained relatively open and where, in particular, there were no special regulations affecting investment by non-residents (Jordan); a market which underwent substantial liberalization, including the withdrawal of restrictions on foreign investors trading in listed equities (Turkey); and a market where restrictive financial market regulations remained in place, with foreign investors prohibited from trading in equities (Pakistan).

The analysis was conducted by first estimating the parameters of agents' preference for risk and preference for current consumption over future consumption, in the context of a model of dynamic consumption and investment decisions. The implications of the model were then examined further by performing excess volatility tests. The tests were performed by constructing an equity market price based on the estimated model and the market fundamentals. The implied equity market price was then compared with the actual evolution of equity prices.

Conditional on our model of asset price determination, the informational efficiency of each of the three markets is found to be deficient. That is, the evolution of actual market prices in these equity markets is influenced by factors other than current and future levels of economic activity and current and future dividends, which can be regarded as the market fundamentals. While it appears that the large negative shock to

economic activity in the late 1980s was an important factor which caused Jordanians to discount market fundamentals, in Turkey and Pakistan the illiquid nature of these equity markets influenced the determination of actual equity market prices. This result highlights the important role that financial market liberalization, and particularly the opening up of previously autarkic equity markets to foreign investors, can play in improving the efficiency with which equity markets mobilize resources for growth-enhancing investment. The results presented here also reveal that there were regime shifts during the sample period in the Turkish and Pakistani equity markets. The former regime shift most likely reflects the opening up of the Turkish equity market to foreigners, which increased the liquidity of the market. The latter regime shift reflects the development of a divergence between the ex-post price of equities determined by market fundamentals, and the actual price of equities in the Pakistani market. This divergence arose in an economic, regulatory and institutional environment which diminished both the demand for, and supply of, traded equities.

This Appendix outlines the various estimators and tests used in this paper.

Appendix A1: The Phillips-Hansen (1990) FM Cointegrating Vector

The FM cointegrating estimator of Phillips and Hansen (1990) is defined as: $\hat{\gamma} = (X'X)^{-1}(X'Y^{+} - T\hat{\delta}^{+})$, where $X = (p_1,\ldots,p_T)$, $Y^{+} = (y_1^{+},\ldots,y_T^{+})'$, $y_t^{+} = p_t^{*} - \hat{\omega}_{12}\Delta p_t/\hat{\omega}_{22}$, $\hat{\delta}^{+} = \hat{\Gamma}(1,-\hat{\omega}_{12}/\hat{\omega}_{22})'$, $\hat{\Gamma} = \hat{\Sigma} + \hat{\Lambda}$, $\hat{\Omega} = \hat{\Gamma} + \hat{\Lambda}'$, $\hat{\Sigma} = T^{-1}\Sigma_{t=1}^{T}\hat{\nu}_t\hat{\nu}_t'$, $\hat{\Lambda} = T^{-1}\Sigma_{\tau=1}^{b}k(\tau/(b+1))\Sigma_{t=\tau+1}^{T}\hat{\nu}_t\hat{\nu}_{t-\tau}'$, $\hat{\omega}_{ij}$ is the ij-th element of $\hat{\Omega}$, $\hat{\nu} = (\hat{e},\Delta p_t)'$, and $\hat{e}$ are the fitted residuals from a 'first step' cointegrating regression of P_t^{*} on P_t. $k(\cdot)$ is the Bartlett kernel window and b is the data dependent bandwidth parameter. The 'long-run' variance estimator required for the test of the restriction that $\gamma = 1$ is given by $\hat{\sigma} = \hat{\omega}_{11} - \hat{\omega}_{12}^{2}/\hat{\omega}_{22}$.

Appendix A2: The Phillips-Ouliaris (1990) Test Statistics

To implement the Phillips and Ouliaris (1990) method here we regress $\hat{e}_t = \hat{\alpha}\hat{e}_{t-1} + \eta_t$ and compute $Z(\alpha) = T(\hat{\alpha}-1) - (1/2)(S_{Tb}^{2} - S_{\eta}^{2})(T^{-2}\Sigma_{t=2}^{T}\hat{e}_{t-1}^{2})^{-1}$, where $S_{\eta}^{2} = T^{-1}\Sigma_{t=1}^{T}\hat{\eta}_t^{2}$, $S_{Tb}^{2} = T^{-1}\Sigma_{t=1}^{T}\hat{\eta}_t^{2} + 2T^{-1}\Sigma_{\tau=1}^{b}k(\tau/b+1))\Sigma_{t=\tau+1}^{T}\hat{\eta}_t\hat{\eta}_{t-\tau}$, $k(\cdot)$ is the Bartlett kernel window and b is the data dependent bandwidth parameter. We then compare the computed value of $Z(\alpha)$ against the critical value in Phillips and Ouliaris (1990, Table Ia, p.189).

Appendix A3: The Hansen (1992) Stability Test

The Hansen (1992) test statistics are defined as:

$$Sup\ F = sup_{\pi \in \Pi} F_{\pi},$$

$$Mean\ F = (T^{*})^{-1} \sum\nolimits^{\prime} F_{\pi},$$

$$Lc = tr \left[M_{\pi}^{-1} \sum_{t=1}^{T} S_{\pi} S_{\pi}^{\prime} / \hat{\sigma} \right],$$

where $\sum^{\prime}$ *denotes the sum over all possible break points in* Π;

$$T^{*} = \sum\nolimits^{\prime} 1;\ F_{\pi} = tr\,(S_{\pi}^{\prime} V_{\pi}^{-1} S_{\pi} / \hat{\sigma});\ S_{\pi} = \sum_{t=1}^{\pi T} \hat{s}_{t};\ \hat{s}_{t} = (p_{t} \hat{e}_{t}^{\prime} - T \hat{\delta}^{+});$$

$$V_{\pi} = M_{\pi} - M_{\pi} M_{\pi}^{-1} M_{\pi};\ and\ M_{\pi} = \sum_{t=1}^{\pi T} p_{t}^{2}.$$

References

Abisourour, A. (1994), "The Emerging Arab Capital Markets: Status, Role, and Development Prospects" in S. El-Naggar, ed., Financial Policies and Capital Markets in Arab Countries, International Monetary Fund, 54-87.

Andrews, D.W.K. (1991), "Heteroskedasticity and Autocorrelation Consistent Covariance Matrix Estimation", Econometrica 59, 817-58.

Atje, R. and B. Jovanovic (1993), "Stock Markets and Development", European Economic Review 37, 632-40.

Breeden, D.T. (1979), "An Intertemporal Asset Pricing Model with Stochastic Consumption and Investment Opportunities", Journal of Financial Economics 7, 265-96.

Dixit, A.K. and R.S. Pindyck (1994), Investment Under Uncertainty, Princeton, Princeton University Press.

El-Erian, M.A. and M.S. Kumar (1994), "Emerging Equity Markets in Middle Eastern Countries", IMF Working Paper 94/103 (Washington: International Monetary Fund, September 1994).

El-Erian, M.A. and S. Tareq (1993), "Economic Reform in Arab Countries: A Review of Structural Issues", in S. El-Naggar (ed.), Economic Development of the Arab Countries: Selected Issues, Washington: International Monetary Fund, 26-50.

Feldman, R.A. and M.S. Kumar (1994), "Emerging Equity Markets: Growth, Benefits, and Policy Considerations", PPAA/94/7 (Washington: International Monetary Fund, March 1994).

Grossman, S.J. and R.J. Shiller (1981), "The Determinants of the Variability of the Stock Market Prices", American Economic Review 71, 222-27.

Hall, A. (1993), "Some Aspects of Generalized Method of Moments Estimation", in C.R. Rao, G.S. Maddala, and H.D. Vinod, eds., Handbook of Statistics, Vol. 11, Econometrics, Amsterdam, North-Holland.

Hamilton, J.D. (1994), Time Series Analysis, Princeton, Princeton University Press.

Hansen, B.E. (1992), "Tests for Parameter Instability in Regressions with I(1) Processes", Journal of Business and Economic Statistics 10, 321-35.

Hansen, L.P. (1982), "Large Sample Properties of Generalized Method of Moments Estimators", Econometrica 50, 1029-54.

_____ and K.J. Singleton (1982), "Generalized Instrumental Variables Estimators of Nonlinear Rational Expectations Models", Econometrica 50, 1269-86.

_____ (1984), "Errata", Econometrica 52, 267-68.

Haque, N. and S. Kardar (1993), "The Development of Financial Markets in Pakistan", mimeo, International Monetary Fund.

International Finance Corporation (1994), Emerging Stock Markets Factbook, Washington D.C.: International Finance Corporation.

Levine, R. and S. Zervos (1995), "Policy, Stock Market Development and Long-Run Growth", paper presented at the World Bank Conference on Stock Markets, Corporate Finance, and Economic Growth, February, Washington, D.C.

McDermott, C.J. (1994), "Excess Volatility: A Testing Strategy", Economics Letters 44, 35-41.

Ogaki, M. (1993), "Generalized Method of Moments: Econometric Applications", in C.R. Rao, G.S. Maddala, and H.D. Vinod, eds., Handbook of Statistics, Vol. 11, Econometrics, Amsterdam, North-Holland.

Phillips, P.C.B. and S. Ouliaris (1990), "Asymptotic Properties of Residual Based Tests for Cointegration", Econometrica 58, 165-93.

Phillips, P.C.B. and B.E. Hansen (1990), "Statistical Inference in Instrumental Variable Regression with I(1) Processes", Review of Economic Studies 57, 99-125.

Shiller, R.J. (1981), "Do Stock Prices Move Too Much to be Justified by Subsequent Changes in Dividends?", American Economic Review 71, 421-35.

_____ (1989), Market Volatility (MIT Press, Cambridge, MA).

Toukan, U.S. (1994), "Jordan as an Emerging Market: Can the Boom Continue?", paper presented at the Royal Institute of International Affairs Conference on Financial Markets in the Middle East, April 28-29, London.