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Cheng, Ai-ru and Jahan-Parvar, Mohammad R. and Rothman, Philip

University of California Santa Cruz, East Carolina University

5 February 2009

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Ai-Ru Cheng*
University of California, Santa Cruz
Mohammad R. Jahan-Parvar†
East Carolina University
Philip Rothman‡
East Carolina University §

February 3, 2009

Abstract

This paper studies excess market returns in the relatively understudied financial markets of nine Middle Eastern and North African (MENA) countries within the context of three variants of the Capital Asset Pricing Model: the static international CAPM; the constant-parameter intertemporal CAPM; and a Markov-switching intertemporal CAPM which allows for the degree of integration with international equity markets to be time-varying. On the whole we find that: (1) Israel and Turkey are most strongly integrated with world financial markets; (2) in most other MENA markets examined there is primarily local pricing of risk and evidence of a positive risk-return trade-off; and (3) there is substantial time variation in the weights on local and global pricing of risk for all of these markets. Our results suggest that investment in many of these markets over the sample studied would have provided returns uncorrelated with global markets, and thus would have served as financial instruments with which portfolio diversification could have been improved.

*Contact: 433 Engineering 2, Department of Economics, University of California, Santa Cruz, Santa Cruz, CA 95064, USA, Phone No: (831) 459-2318, e-mail: archeng@ucsc.edu

†Contact: Brewster A-426, Department of Economics, East Carolina University, Greenville, NC 27858-4353, USA, Phone No: (252) 328-4770, e-mail: jahanparvarm@ecu.edu

‡Contact: Brewster A-424, Department of Economics East, Carolina University, Greenville, NC 27858-4353, USA, Phone No: (252) 328-6151, e-mail: rothmanp@ecu.edu

§We would like to thank David Peel, Stuart Hyde, Basma Majerbi, John Hund, and seminar participants at Lancaster University, the Norges Bank, the Third ESRC Research Seminar on "Nonlinear Economics and Finance Research Community" held at Keele University, the University of Leeds, the University of Wisconsin-Milwaukee, Virginia Commonwealth University, the Western Economic Association 2007 Annual Meeting, and the 2008 Triangle Econometrics Conference for useful comments. Ming Lu provided research assistance.

1 Introduction

The financial literature is thin on the Middle East and North Africa (MENA) region. Since many MENA financial markets are rather new, this may not be surprising. But a gap in the literature exists, especially in the light of the superior performance of many of these markets in recent years. These financial markets have posted high returns and grown fast. For example, Saudi Arabia's stock exchange had a market capitalization larger than that of South Korea in 2004-2005.

Following the turn of the century, the MENA region experienced significant oil windfalls up to the middle of 2008. Further, many companies in the area have done well and expanded beyond their traditional markets. Saudi Basic Industries Corporation (SABIC) bought General Electric's GE Plastics for USD 11.6 billion on May 21, 2007. In 2006, SABIC ranked 331$^{\text{st}}$ in the *Fortune Global 500* list, with an estimated revenue of USD 20.9 billion and equity worth USD 16.6 billion, ranking 6$^{\text{th}}$ among international chemical producers.[1] Koc Holdings of Turkey, with over USD 18 billion in revenue in 2006, is ranked 358$^{\text{th}}$ in the *Fortune Global 500*. Many Israeli companies are world class leaders in hi-tech sectors. Dubai in the United Arab Emirates (UAE) is trying to position itself as an important financial center between Hong Kong and London. The Shaheen Business & Investment Group of Jordan is an international business conglomerate which operates globally; its activities benefit considerably from Jordan's free trade agreement with the US. Egypt's Orascom is an important telecom and construction player in the MENA sphere and South Asia. Even the non-profit world of academia is responding to financial developments in the MENA region. For example, in June of 2007 the Harvard Management Company, which is responsible for the university's endowment, announced a USD 1 billion investment in MENA Arabic financial markets in collaboration with Egypt's Hermes Funds.

It is worth noting that these high financial market returns have been realized while the MENA area has experienced major political and security instability, the War on Terror, civil war in Iraq, deteriorating relations with the West, and turmoil in world oil markets. From the macroeconomic point of view the MENA region is important not only because six out of the twenty major oil-producing countries are located there, or since the area contains the largest reserves of fossil fuels. As argued in a series of papers by Hamilton (1983, 1985, 1996, 2003), most major global recessions since the Second World War followed either oil price shocks or political instability in or originating from MENA.

While there are many studies dealing with equity markets, risk, and returns in emerging economies, only a small number of them examine the MENA region. Besides investigation of Israel during the hyperinflation period of the 1980s, very few researchers have studied other countries in this area. One of the best examples is Ghysels and Cherkaoui (2003), who conduct an in-depth

[1] Sources: the SABIC 2006 annual report and the 2006 *Fortune Global 500*.

analysis of trading costs in Morocco. Kim and Singal (2000) consider the level and volatility of returns in Jordan and Turkey around the opening up of their financial markets. Errunza (2001) focuses on issues pertaining to the liberalization and integration of financial markets in Egypt, Israel, Jordan, Morocco, and Turkey, but not the Persian Gulf region. Gulen and Mayhew (2000) include Israel in their study of stock market volatility before and after the introduction of equity-index futures trading in twenty-five countries.

Our goal is to answer the following questions. First, is there evidence of static international CAPM efficiency in MENA markets and are these financial markets integrated with or segmented from global equity markets? By static we mean a framework based upon Sharpe (1964) and Lintner (1965), such that it is assumed the set of investment opportunities is constant, and our use of international in this context means the relevant CAPM market portfolio is given by the "world market" portfolio as measured by, for example, the Dow Jones World Index (DJW). Our check of CAPM efficiency and financial market integration in this setup is based upon examination of estimates of, respectively, alpha and beta.[2] We also investigate whether it is useful to augment the static international CAPM by addition of a select number of additional factors. Included in the group of such factors we use are significant event-periods extracted from the MENA data using the methodology of Hinich and Serletis (2007). Second, is there a significantly positive risk-return trade-off in these markets? Such a trade-off is implied by the intertemporal CAPM of Merton (1973). We address this issue by modeling the excess returns in the MENA markets through a GARCH-in-Mean (GARCH-M) approach. Third, is there time variation in the extent to which these markets are segmented from or integrated in the world financial system? Our static international CAPM results lead us to conclude that one of two polar extremes applies: the market is either segmented or integrated. Bekaert and Harvey (1995) developed a more flexible model which allows the degree of integration with world capital markets to vary across time and we estimate such models for the MENA markets; this methodology also allows us to consider the existence of a positive risk-return trade-off for the MENA financial markets. The answers to these questions have important implications for asset pricing, portfolio selection, and risk management for investors interested in opportunities available in these markets, as well as for scholars who study international aspects of finance theory and practice.

In Section 2 we introduce the data used in our research. Section 3 discusses our static international CAPM and factor model analysis. We present the results of GARCH-M modeling of the conditional mean of expected returns in MENA financial markets in Section 4, and we explore the time-varying nature of integration versus segmentation in these markets in Section 5. Section 6 concludes.

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[2]If a market is completely segmented, the covariance of its excess return with the excess return on the world market portfolio will be zero, such that the beta from its static international CAPM will also be zero.

2 Data

We use financial data for the MENA region from Thomson Financial's *Datastream* data bank. We collected data for nine countries based on the availability and length of data sets maintained by *Datastream*. The countries we included are Bahrain, Egypt, Israel, Jordan, Kuwait, Morocco, Oman, Saudi Arabia, and Turkey. Bahrain, Kuwait, Oman, and Saudi Arabia represent oil exporting and rich Persian Gulf basin markets. Israel and Turkey are the most advanced and globalized economies in the MENA region. Jordan and Egypt, while not oil exporters themselves, have strong trade and financial ties to the Persian Gulf region oil exporters. Morocco is a representative North African country, but in many ways its market is more integrated with Europe than with the rest of the MENA countries.

In order to maintain uniformity of results, we use US dollar denominated returns for all the markets. The data are sampled at daily frequency. We could not get higher frequency data for the Arabic countries. The length of the data samples are not uniform. For Egypt, Jordan, Israel, Morocco, and Turkey our sample spans July 7, 1997 to February 15, 2008. Data for Bahrain, Kuwait, and Saudi Arabia run from March 1, 2000 to February 15, 2008. Oman has the shortest data span, July 17, 2000 to February 15, 2008. It would have been optimal to include more countries, but we were quite constrained by data availability. For example, available USD market returns for Lebanon, Qatar, and the UAE start only in 2005.

Total market return index data for Bahrain, Egypt, Jordan, Morocco, Oman, and Saudi Arabia are reported by Standard and Poor's/IFCG. The same data for Israel and our proxy for the world financial market index, the DJW series, are reported by Dow Jones. Turkey's data are from FTSE World and Kuwait's data are reported by KIC. We use daily returns, computed as the log differences of market total return indices.

Our proxy for the risk-free rate is the daily 3-month secondary market US T-bill rate from Federal Reserve Bank of St. Louis FRED II database. Many Arabic countries do not have an active debt market. Moreover, the monetary authorities in these countries typically do not act independently. Some countries such as Saudi Arabia adhere to a strict reading of Islamic Shari'a law that in effect prohibits charging interest on deposits. The posted interest rates are not calculated through familiar machinations of financial and money markets, but through an ad-hoc Shari'a-based formula. Hence: interest rates across Islamic countries are not compatible with their free-market counterparts; these rates may not reflect the true cost of capital in at least some of the countries in our sample; and many investors look at the international market for assessing their opportunity costs. For these investors, the true benchmark is either a US T-bill or LIBOR rate. We chose a T-Bill rate.

Summary statistics for the excess returns series we use are presented in Table 1. The following properties of the data are worth noting. First, the sample mean of the excess returns in each

MENA market is an order of magnitude larger than that for the DJW. This, along with the MENA estimated unconditional second moments being of the same order of magnitude as for the DJW, is the sense in which we refer to the superior performance in these markets above. Second, none of the excess returns series exhibits heavy unconditional skewness. Third, as is common for financial market returns, the MENA series are highly leptokurtotic and thus non-Gaussian.

We carried out extensive stationarity time series tests on the available data. The empirical evidence reveals that the index data are non-stationary at the logarithmic level, while the unit-root null can be rejected at conventional significance levels for the returns data.

3 Static International CAPM and Factor Models

We are interested in testing market efficiency in our sample of MENA stock exchanges. The workhorse model of modern equity pricing since the 1960s has been the CAPM. It comes in many flavors and our initial choice is the Sharpe (1964) and Lintner (1965) variation. This model states that the expected excess returns of an asset are linearly dependent on excess market returns. Empirically, the systematic risk of the asset is estimated by regressing its excess returns on some measure of excess returns of a broad equity market measure. To apply the model to an international setting, the excess returns of a national market are regressed against the excess returns of an index composite of international markets.

There are well-documented criticisms to the CAPM and two remedies are often considered. The most common approach is to use the Fama and French (1996) methodology. We can not use this scheme since Fama-French factors are not available for the majority of the markets we study. As an alternative, we use a variant of multifactor models. A classic example is Chen et al. (1986), who link stock market performance to a set of well-known macroeconomic factors. Since this model is developed mainly for developed markets and, moreover, some of the variables used in Chen et al. (1986), such as the default and term premia, are not recorded for many MENA markets, we opt for an alternative formulation.[3] We postulate that oil prices have an impact on market performance in the most important oil-producing region in the world. In addition, we also test for the possibility of a relationship between expected excess returns and the squared values of both lagged local excess returns and world excess returns. Inclusion of these two variables may capture some nonlinear departure from the traditional international CAPM. We also allow for the possibility that there are time-specific events which may have an impact on the behavior of expected returns. The typical approach is to conduct event study analysis. We chose an alternative, based on the research of Hinich (1996), Hinich and Patterson (2005), and Hinich and Serletis (2007), which in our opinion is at least as effective, if not superior, for markets with limited coverage of information. Through

[3] In many MENA markets, there is no concept of domestic corporate debt market.

our use of this methodology, we produce dummy variables we call "Hinich factors" which indicate if a given observation falls within an "episode" of nonlinearity.[4]

The Sharpe (1964) and Lintner (1965) formulation of the CAPM is given by:

$$r_t^i = \alpha_i + \beta r_t^W + \epsilon_t^i \tag{1}$$

where r_t^i is the market excess return in country i, r_t^W is the world market excess return, and ϵ_t^i is assumed to be a white noise innovation process. As mentioned earlier, we use the DJW index returns as a proxy for world market returns and the 3-month US T-bill rate as a proxy for the global risk-free rate. The above variant of the international CAPM assumes there is no exchange-rate risk. Under certain conditions, exchange-rate risk is not priced independently from market risk; see, for example, Adler and Dumas (1983).

A necessary condition for the i^{th} market to be CAPM efficient is $\alpha = 0$. If $\beta = 0$, then the i^{th} market is segmented from the international capital market. International CAPM results for the full sample are presented in the first two rows of Table 2. The models were estimated by OLS and Newey-West HAC standard errors were computed; see Newey and West (1987).

The empirical results show that for the plain vanilla international CAPM, the $\hat{\alpha}$'s are significantly different from zero at conventional significance levels for five countries: Bahrain, Kuwait, Oman, Jordan, and Morocco. This implies that these MENA markets are CAPM inefficient. Though all of these point estimates are very small, on the order of 10^{-4}, they are of the same order of magnitude for the sample means of the daily excess returns in these markets, implying that they are economically significant. In contrast, the $\hat{\alpha}$'s are insignificant for Saudi Arabia, Egypt, Israel, and Turkey, suggesting that these markets are CAPM efficient. Second, the $\hat{\beta}$'s are significantly different from zero for three markets: Bahrain (10% level), Israel (1% level), and Turkey (1% level). All other international CAPM $\hat{\beta}$'s are insignificant at conventional levels, implying that risk premia in these markets are priced locally. While the $\hat{\beta}$ for Bahrain is statistically significant, we feel it is not economically significant since its value is rather small (0.028). In the case of Turkey, $\hat{\beta}$ is large (0.522), implying strong integration with world equity markets. For Israel, the value of $\hat{\beta}$ is smaller than Turkey's (0.164), but it is still economically significant.

In the next step, we test whether augmenting the model with the factors discussed above affects these results on asset pricing efficiency and capital market integration obtained with the simple international CAPM; exclusion of these factors can be a source of omitted variables bias. The

[4]We outline this procedure and explain why we favor it in the Appendix, where we also define the Hinich factor dummy variables we use.

factor model is given by:

$$r_t^i = \alpha + \beta r_t^M + \sum_{j=1}^{3} \delta_j F_{t,j} + \sum_{k=1}^{M} \gamma_k d_{t,k}^i + \epsilon_t^i, \tag{2}$$

where the factors $F_{t,j}$ are the log differences in the daily spot oil price ($j = 1$), the squared log differences in the spot oil price ($j = 2$), and the squared world market excess returns ($j = 3$), and the $d_{t,k}^i$ variables represent the Hinich factors.[5]

The results are reported in the third through last rows in Table (2) as in the previous case, the models were estimated by OLS and Newey-West HAC standard errors were computed. On the whole, inclusion of the additional factor components does not change the conclusions under the static international CAPM specification. The Israeli and Turkish markets both continue to be efficient in the sense that their $\hat{\alpha}$'s have high $p-$values, and their $\hat{\beta}$'s are both statistically and economically significant. Under the factor model specification, the evidence still suggests that the other MENA markets are segmented from international capital markets. Interestingly, use of the factor variables leads to the conclusion that the Saudi Arabian market, in contrast to the outcome under a simpler specification, is CAPM-inefficient; five other MENA markets remain CAPM-inefficient by estimation of equation (2).

Very few of the factor variables are significant at conventional significance levels. We have two observations on this. First, given the important role most of the MENA countries play in the world oil market, we find it surprising that, based upon our estimated models, the price of oil is apparently not conditionally correlated with aggregate equity returns in these markets. Second, if there is an important nonlinear aspect to behavior of excess returns in the MENA equity markets, it apparently is not captured by the additional factors we consider. The framework we use in Section 5 provides an alternative approach for modeling possible nonlinearity in the excess returns in these markets.

4 Constant-Parameter Intertemporal CAPM

Merton (1973) extended the static CAPM of Sharpe (1964) and Lintner (1965) to an intertemporal framework which allows for a changing set of investment opportunities. In his intertemporal CAPM, the expected conditional excess return for market i should vary positively with its conditional variance:

$$\mathrm{E}_{t-1}[r_t^i] = \mu + \lambda \mathrm{Var}_{t-1}[r_t^i], \tag{3}$$

[5]We used West Texas Intermediate spot oil prices from the US Department of Energy's database.

where the parameter λ is the coefficient of relative risk aversion of the representative agent.[6] λ is also referred to as the risk premium associated with market risk. If the intertemporal CAPM holds, then $\mu = 0$.

To investigate whether there is a risk-return trade-off of the intertemporal CAPM sort in MENA financial markets, we fit GARCH-M models to the excess returns series. Bekaert and Harvey (1997) emphasize that equity returns in emerging markets exhibit substantial asymmetry in volatility, possibly due to a leverage effect in which firms' leverage increases with negative returns. Accordingly, we use two GARCH-M specifications developed to allow for such asymmetry. In both cases the conditional mean for the excess returns in market i is given by:

$$r_t^i = \mu + \lambda h_{t-1}^i + \varepsilon_t, \tag{4}$$

where $\varepsilon_t = \sqrt{h_t} e_t$, $e_t \sim N(0,1)$, and h_t^i is the conditional variance of r_t^i. The first GARCH conditional volatility structure we use is the Exponential GARCH (EGARCH) model of Nelson (1990):

$$\begin{aligned} \ln(h_t^i) &= \omega + \alpha g(z_{t-1}) + \beta \ln(h_{t-1}^i) & (5) \\ g(z_t) &= \theta z_t + \delta[|z_t| - \mathbb{E}|z_t|], & (6) \end{aligned}$$

where $z_t = \varepsilon_t / \sqrt{h_t}$ and $\delta = 1$. We refer to equations (4), (5), and (6) jointly as an EGARCH-M model.

Our second GARCH specification follows Glosten et al. (1993) (GJR):

$$h_t^i = \omega + \alpha \varepsilon_{t-1}^2 + \gamma I_{\{\varepsilon_{t-1}<0\}} \varepsilon_{t-1}^2 + \beta h_{t-1}^i, \tag{7}$$

where $I_{\{\varepsilon_{t-1}<0\}}$ is an indicator function which takes on the value of 1 when $\varepsilon_{t-1} < 0$ and 0 otherwise. We refer to equations (4) and (7) jointly as a GJR GARCH-M model.

We obtain parameter estimates by joint maximum likelihood estimation of the conditional mean and variance equations for both the EGARCH-M and GJR GARCH-M models. In all cases, convergence in estimation is achieved in 50 or less iterations. The results are reported in Table 3.

Using the EGARCH-M specification, there is a statistically significant positive risk-return trade-off in four of the MENA markets: Bahrain, Saudi Arabia, Egypt, and Jordan. The GJR GARCH-M estimated $\hat{\lambda}$'s are also significantly positive for Bahrain, Egypt, and Jordan, but not for Saudi Arabia. Both the EGARCH-M and GJR GARCH-M $\hat{\lambda}$'s are economically reasonable for Egypt (7.881 and 4.787) and Jordan (6.808 and 5.292), while those for Bahrain appear to be too high to

[6] This conditional single-factor formulation follows under the assumption that the variance of the change in wealth is much larger than the variance of the change in the state variable with which wealth varies; see Merton (1980).

be economically significant (40.317 and 30.314).[7] The EGARCH-M $\hat{\lambda}$ for Saudi Arabia (3.056) is also economically sensible. In no other MENA market is there a statistically significant risk-return trade-off. For two MENA markets, Israel and Turkey, all $\hat{\lambda}$'s are negative but not statistically significant.

Under both the EGARCH-M and GJR GARCH-M specifications, the estimated intercepts are insignificant for Oman and Morocco. The GJR GARCH-M $\hat{\mu}$ is insignificant for Jordan, but the EGARCH-M estimated intercept for Jordan is significant. For all other MENA markets, $\hat{\mu}$ is significant using both the EGARCH-M and GJR GARCH-M models. This may reflect the absence, in our conditional mean equations, of other state variables which covary with the excess return in these MENA markets. This may also be due to compensation for jump risk; see, for example, Pan (2002).

The strongest evidence in favor of the intertemporal CAPM is offered by the GJR GARCH-M conditional mean intercept and slope estimates for Jordan. In this case, there is an economically and statistically significant risk-return trade-off coupled with a statistically insignificant $\hat{\mu}$. Holding constant the statistically significant intercepts, our positive risk-return trade-off results also support the intertemporal CAPM for Bahrain and Egypt under both GARCH-M specifications, and for Jordan and Saudi Arabia via the EGARCH-M specification. It is interesting to note that the evidence in favor of the intertemporal CAPM is quite weak for both Israel and Turkey, the two markets for which the static international CAPM strongly suggest integration with world equity markets.

Our use of the EGARCH and GJR conditional variance models was motivated by the observation of Bekaert and Harvey (1997) on volatility asymmetry in emerging markets. Accordingly, we think it is helpful to examine the extent to which our asymmetric GARCH-M models are consistent with such asymmetry. With the exception of Morocco, the estimated values of the asymmetry parameters, i.e., θ in equation (6) and γ in equation (7), are statistically significant at conventional significance levels for all markets. However, the signs of these parameters are consistent with the leverage effect, i.e., $\hat{\theta} < 0$ and $\hat{\gamma} > 0$, for only three MENA markets: Kuwait, Israel, and Turkey.

5 Markov-Switching Intertemporal CAPM

International finance theory includes an active line of research studying market integration versus segmentation. Some examples related to our study include Harvey (1991), Errunza et al. (1992), Harvey (1995), Bekaert and Harvey (1997), and more recently Bekaert et al. (2008). The thrust of this line of research is the study of country-specific versus global pricing of risk premia. As noted

[7] The arguments of Kandel and Stambaugh (1990), however, imply that the $\hat{\lambda}$'s for Bahrain may not be too high to make economic sense.

by Bekaert and Harvey (1995), empirical evidence suggests that expected returns of assets with the same level of exposure to risk factors are influenced by their "nationality." Such results are consistent with incomplete equity market integration.

Bekaert and Harvey (1995) propose a conditional regime-switching model which generalizes the Sharpe (1964), Lintner (1965), and Merton (1973) asset pricing models to allow for time-varying weights on local and global pricing of an asset. We use this framework to study the extent to which the MENA financial markets' degree of integration with world capital markets changes across time.

Let S_t^i be a latent state variable for market i which can take on two values, with $S_t^i = 1$ denoting that market i is integrated with international equity markets in observation t and $S_t^i = 2$ denoting it is segmented. Define:

$$\phi_{t-1}^i = \text{Prob}(S_t^i = 1 | \mathcal{F}_{t-1}), \tag{8}$$

where $\mathcal{F}_{t-1}$ is the observation $t-1$ information set. As before, let r_t^i and r_t^W be, respectively, the excess return for market i and the world market. Bekaert and Harvey (1995) model r_t^i as:

$$r_t^i = \phi_{t-1}^i \lambda_{t-1} \text{Cov}_{t-1}[r_t^i, r_t^W] + (1 - \phi_{t-1}^i) \lambda_{t-1}^i \text{Var}_{t-1}[r_t^i] + \varepsilon_t^i, \tag{9}$$

where λ_{t-1} and λ_{t-1}^i are the time-varying risk premia associated with world market systematic risk and country-specific idiosyncratic risk.[8] While the above framework allows the probability of integration, ϕ_{t-1}^i, to vary across time, we assume that the transition probabilities $p_{1,1}^i = \text{Prob}(S_t^i = 1 | S_{t-1}^i = 1)$ and $p_{2,2}^i = \text{Prob}(S_t^i = 2 | S_{t-1}^i = 2)$ are constant. Time variation in the risk premia is allowed as follows:

$$\lambda_{t-1} = \exp(\psi' Z_{t-1}) \tag{10}$$

$$\lambda_{t-1}^i = \exp({\psi_i}' Z_{t-1}^i), \tag{11}$$

where ψ and ψ_i are parameter vectors, and Z_t and Z_t^i are vectors of state variables that capture world market information and country i specific information at time t. We also consider the case in which the risk premia λ and λ^i are constant:

$$\lambda = \exp(c_1) \tag{12}$$

$$\lambda^i = \exp(c_2). \tag{13}$$

Through use of the exponential function in (10)-(11) and (12)-(13), we constrain each risk premium to be positive.

We estimate the model, in both the constant risk premia and time-varying risk premia cases,

[8]Following Bekaert and Harvey (1995), we do not include an intercept term in equation 9.

by maximum likelihood. Estimation is carried out in two stages. First, we compute $\mathrm{Var}_{t-1}[r_t^i]$ and $\mathrm{Cov}_{t-1}[r_t^i, r_t^W]$ using a rolling window estimation scheme.[9] Second, we form the likelihood function according to the model in equation (9) and maximize it. To avoid local optima, we perturb our starting values and re-estimated the model 50 times for each market.

Following Bekaert and Harvey (1995), we use a set of global and local instrumental variables as components of, respectively, Z_t and Z_t^i to study the behavior of the time-varying risk premia in the MENA markets. The global instrumental variables we use are the log differences on the DJW market capitalization, the default spread captured by changes in the difference between Moody's Aaa and Baa bond yields, changes in the yields on US commercial paper, and the term structure spread captured by the difference between the US 10-year bond and 3-month T-bill yields.[10] These variable are designed to capture fluctuations in expectations of the world business cycle. The local instrumental variables we use include the returns on the market index, changes in market dividend payments, and changes in market valuation in each country i.

We find that including Z_t and Z_t^i, and hence allowing for time-varying premia, does not improve the estimation results significantly. In fact, in several cases there are problems with the size of the estimated parameters.[11] As a result, we only discuss the results obtained through estimation of the constant risk premia model.

We are interested in the behavior over time of the estimated probabilities of integration, i.e., ϕ_{t-1}^i for each market i. High values of these probabilities show that pricing of assets in market i is done primarily with respect to the covariance of the assets with the world market excess return (integration), and low probabilities imply mostly local pricing of risk (segmentation).

Figure 1 shows the histograms of these probabilities for all countries in this study. It is worth noting that these histograms are generally bimodal, with probability masses concentrated in the "high" end of plot (integration) and in the "low" end (segmentation). Inspection of the histograms suggests that Bahrain, Israel, and Turkey are considerably more integrated than the other MENA countries, with 40% or more of asset pricing days having very high $\hat{\phi}_{t-1}^i$ values, while Egypt, Jordan, Oman, and Saudi Arabia are overwhelmingly segmented, with 60% or more of the estimated probabilities of integration being quite low. Though Kuwait and Morocco show a more mixed

[9]We fix a sub-sample period of m days for calculating the variance of r_t^i and the covariance between r_t^i and r_t^W, and roll the sample one day forward to compute for the next pair of statistics. In order to find a sensible value for m, we look at the estimated partial autocorrelation function of the squared excess returns and include all the lags that have a significant impact on the current level.

[10]The default spreads, US commercial paper yields, and term structure data are all from the FRED II data bank maintained by the St. Louis FED. The maturity of the default spreads data is 30 years and the maturity of the commercial paper yields is 3 months.

[11]More specifically, many of the elements of the parameter vectors $\hat{\psi}$ and $\hat{\psi}_i$ are unreasonably large in magnitude, "blowing up" in both the positive and negative directions.

picture than the other Arabic countries, with a higher relative tendency towards integration, the degree of integration in these two countries is generally quite low.

The summary statistics on the $\hat{\phi}^i_{t-1}$ values presented in the top panel Table 4 support these conclusions. The median of $\hat{\phi}^i_{t-1}$ for Bahrain, Israel, and Turkey is, respectively, 1.0, 0.783, and 0.754, suggesting a median tendency towards predominantly global pricing of risk in these markets. On the other hand, for Egypt, Jordan, Oman, and Saudi Arabia the median of $\hat{\phi}^i_{t-1}$ in each case is at the low polar value of 0.0, implying a median high weight on local pricing of risk. For Kuwait and Morocco, the median of $\hat{\phi}^i_{t-1}$ is, respectively, 0.200 and 0.253, indicating a more so intermediate case.

Information about the persistence of the unobserved integrated and segmented states is provided by the estimates of the transition probabilities $p^i_{1,1}$ and $p^i_{2,2}$ in the middle panel of Table 4. For four countries, Kuwait, Israel, Morocco, and Turkey, both of these staying probabilities are greater than 0.80, indicating a strong degree of persistence of both states. For Bahrain, the estimated probability of staying in the segmented state is quite low, at roughly 0.20, while the degree of persistence of the integrated state is considerably higher. The opposite holds for Oman, Saudi Arabia, Egypt, and Morocco.

The bottom panel of Table 4 presents the estimated global and local risk premia, $\hat{\lambda}$ and $\hat{\lambda}^i$, for each market. In brackets under each estimated risk premium appears the $p-$value for a likelihood ratio test of the null hypothesis that the coefficient equals zero against the one-sided alternative that it is positive.[12] On the whole, these results also support our conclusions obtained from inspection of the histograms in Figure 1. For both Israel and Turkey, the global risk premium is significantly positive while the local risk premium is not. For Kuwait, Oman, and Saudi Arabia, the opposite holds; for Saudi Arabia the $p-$value for the local risk premium is considerably higher than it is for either Kuwait or Oman. Once again, Morocco offers an intermediate case in that both the global and local risk premium are significantly positive at conventional significance levels. For Bahrain, Egypt, and Jordan, neither risk premium is significant at the 10% significance level.

Figure 2 presents time series plots of these estimated probabilities for three MENA countries during three particularly volatile sub-samples. Our objective is to show how our results suggest that an increase in "instability" appears to lead to a shift away from the general trend in the pricing of risk. That is, if there is usually global (local) pricing of risk in the country's financial markets, then during a period of increased instability, due to political, economic, or other factors, there is a shift towards local (global) pricing of risk.

First, consider the behavior of the estimated probabilities of integration for the Israeli market during the buildup to and through the summer 2006 war in Lebanon. As is seen in Figure 2, during

[12]The risk premium parameter in question was set equal to zero in the "constrained" model. All risk premia estimated in both the "constrained" and "unconstrained" models were constrained to be positive.

the month of June the Israeli market swung between local and international pricing in the wake of increasing violence between the Israeli Defense Force (IDF) and militants in the Gaza Strip. On July 13$^{\text{th}}$, 2006, two Israeli soldiers were kidnapped by the Lebanese paramilitary organization Hezbollah. On the same day Hezbollah launched missile attacks into northern Israel. Following so soon after the abduction of a soldier in the Gaza Strip on July 25$^{\text{th}}$, the IDF responded with a ferocious wave of air raids and artillery assaults on Lebanon. As the Second Lebanon War began, our results suggest that there was a dramatic shift to local pricing of risk, arguably with the ongoing war as the main risk factor. As the likelihood of a ceasefire grew during the early part of August, the market increasingly priced assets in line with integration; a ceasefire went into effect on August 14$^{\text{th}}$. By mid-August 2006, the estimated integration probabilities were close to one, implying a high degree of integration, which our earlier discussion suggests is reflective of the median behavior for the Israeli market.

Our second case focuses on the Turkish market during the financial turmoil of December 2000 to February 2001. In 2000, the Turkish central bank and government implemented a currency peg-based stabilization program aimed at ending decades of high inflation. For various reasons, including reliance of the program on inflows of "hot money," a weak banking system, and other institutional factors, the program faced severe problems in November and December of 2000; for more details, see Alper (2001). On December 1$^{\text{st}}$, 2001, the overnight interbank interest rate reached 1,700%. By December 5$^{\text{th}}$, the financial system was about to collapse. As a result, the IMF extended a rescue package worth USD 10 billion on December 6$^{\text{th}}$, 2000. As is seen in Figure 2, Turkey's financial markets generally seem to have been integrated moving towards the end of that December. The sharp drop in $\hat{\phi}^i_{t-1}$ at the start of January 2001 may reflect the large bets that hedge funds and other investors were making against the Turkish lira. The peaking of the integration probabilities between January 19$^{\text{th}}$ and February 2$^{\text{nd}}$ coincided with propagation of news regarding the IMF's package and attempts by the government to calm the markets. By this point in time the nearly USD 6 billion in capital that had exited the country as the financial crisis broke out in late 2000 had flowed back. But during the month of February 2001, the peg-based stabilization program was abandoned.[13] Our results suggest that as this major policy reversal occurred, the market was paying exclusive attention to local risks. It is not until a couple weeks into March 2001 that markets returned to integrated pricing, which our results in Figure 1 and Table 4 suggest is the norm for Turkey.

In the third case, we look at the behavior of the Kuwait stock exchange around the terrorist attacks in the US on September 11$^{\text{th}}$, 2001, up through the initial phase of the following US-led invasion of Afghanistan. Recall that our earlier analysis suggests that Kuwaiti financial markets,

[13]Exacerbating the sense of crisis in the country, a rather public row between President Ahmet Necdet Sezer and Prime Minister Bulent Ecevit which broke out on February 19$^{\text{th}}$, 2001.

along with those of most other countries in the Persian Gulf basin, are generally segmented. Figure 2 suggests that, for the month prior to the September 11$^{\text{th}}$ attacks, pricing of risk in Kuwait was generally local; on almost every day, $\hat{\phi}^i_{t-1}$ was considerably below 0.5. Then, immediately after the September 11$^{\text{th}}$ attacks, there was a marked shift to global pricing of risk. This continued through the start of US and British bombing on Taliban communication and military facilities in Afghanistan on October 7$^{\text{th}}$, 2001, and throughout the month of October. By the start of November 2001, there was a return to the segmented state for Kuwait.

6 Conclusions

In this paper, we provide a detailed study of the behavior of equity markets in the MENA region through use of several variants of the CAPM. Our study is, we believe, the most comprehensive empirical analysis of the risk and return dynamics in the MENA markets to date. Given the strong growth and importance of these markets, we believe our results will be of interest to the finance literature as well as financial practitioners and policy makers.

A major concern of the paper is the extent to which these markets are integrated with world capital markets, and we found that for all of the MENA markets there is substantial time variation in the degree of such integration. The Israeli and Turkish markets are strongly integrated with world equity markets. This conclusion is supported by both our static international CAPM and Markov-switching intertemporal CAPM analysis.[14] That said, pricing in these markets is done locally on roughly twenty percent of the trading days in our sample. Our constant-parameter intertemporal CAPM results suggests there is no risk-return trade-off in the Israeli and Turkish markets. This is arguably supported by our Markov-switching intertemporal CAPM models estimated for these countries, since the estimated "local" risk premia are not significantly greater than zero.

While the other MENA markets are generally strongly segmented from international capital markets, pricing in them is done globally on at least ten percent of the trading days in our sample. Bahrain appears to be an exception, in that the vast bulk of the estimated integration probabilities are greater than 0.90; but since the estimated risk premia for Bahrain are not significantly greater than zero, we have doubts about the reliability of our dominant global pricing of risk finding in this case. For each of these countries, evidence in favor of a positive risk-return trade-off is provided by either our constant-parameter intertemporal CAPM analysis or our Markov-switching intertemporal CAPM exercise; our results on this question generated by these two different approaches, however, are not consistent for these MENA markets.

Our study suggests that investment in most of the Arabic MENA markets, at least for the

[14] While neither of these specifications is likely to be the "true" model, we note that each implies the other model is misspecified.

sample period we study, provides returns uncorrelated with global markets, and thus would serve as financial instruments with which portfolio diversification could be improved. However, in the midst of the global financial crisis which erupted in September 2008, returns in these markets also plummeted. We speculate that there is an economically important link between oil price movements and the extent to which these markets are integrated with global capital markets. More specifically, we suggest that, all else equal, financial market integration decreases with oil price increases and vice versa. We plan on addressing this question in future work.

Appendix: Episodic Nonlinear Event Detection

To produce additional explanatory variables for both our multifactor and conditional volatility models, we are interested in identifying periods containing significant events for the behavior of financial market returns in a particular country of interest. We chose an approach which uses the data to isolate events which are significant. More specifically, to achieve this objective we apply the "episodic nonlinear event detection" method of Hinich and Serletis (2007) explained below. This procedure is based on Hinich (1996), who introduces a test for third-order correlation which can be viewed as the time-domain analogue of the bispectrum test of Hinich (1982).

We prefer this line of attack over postulating when an event could have occurred and then testing for significant changes based on this guess; see Binder (1998) for an overview of the event study literature and its application in finance. Typical event study analysis depends on transparent and readily available financial reporting. These criteria may be lacking for at least some MENA markets. While very well-known events are trivially detectable, there are events that may not be as obvious unless the data are studied carefully. Alternatively, it is possible that an event that appears significant at first glance may not be as influential empirically.

To carry out the exercise, we break the series into 50-day frames, approximately equivalent to 10 trading weeks. Let the length of each frame be ℓ. We standardize the data in each frame by subtracting the mean and dividing by the frame's standard deviation. Denote the standardized data in the nth frame by $\{z_t^n\}$. The goal is to detect evidence in favor of third-order correlation in the nth frame using the Hinich (1996) portmanteau bicorrelation test statistic, which follows:

$$H_n = \sum_{r=2}^{L}\sum_{u=1}^{r-1}(\ell-u)^{-1}[B_n(r,u)]^2 \tag{14}$$

$$B_n(r,u) = \sum_{t=1}^{\ell-r} z_t^n z_{t+r}^n z_{t+u}^n \tag{15}$$

Under the null hypothesis that the observed process is pure white noise (iid), if ℓ is sufficiently large and $L=\ell^c$ where $0<c<0.5$, then $H_n \sim \chi^2_{L(L-1)/2}$. Under this null hypothesis, $U=F(H_n)$

has a uniform (0,1) distribution, where F is the cumulative distribution function of $\chi^2_{L(L-1)/2}$. Using FORTRAN code provided by Hinich, and setting $c = 0.4$ as suggested in Hinich and Serletis (2007), we apply the test to the excess returns data for each country to extract the M frames for which the null hypothesis is rejected at the 5% significance level.[15] Call each of these M frames a "significant frame" in which there is, following Hinich and Serletis (2007), a "nonlinear event." For each country i we create "Hinich factor" dummy variables, $d^i_{t,k}, k = 1, ..., M$, one corresponding to each of the M significant frames. The values of these binary variables are determined as follows: if a given observation from the excess returns series for country i falls in the kth significant frame, then $d^i_{t,k} = 1$; $d^i_{t,k} = 0$ otherwise.[16]

There are several chronological tables available for important or potentially influential financial, economic, and political events in emerging financial markets. The reader may want to refer to Henry (1999) or the online tables maintained by Bekaert and Harvey at Duke University.[17] We find that most of detected significant frames coincide with the important events reported in the chronology of Bekaert and Harvey. This qualitative comparison supports our view in applying Hinich analysis in detecting significant frames for each return series.[18]

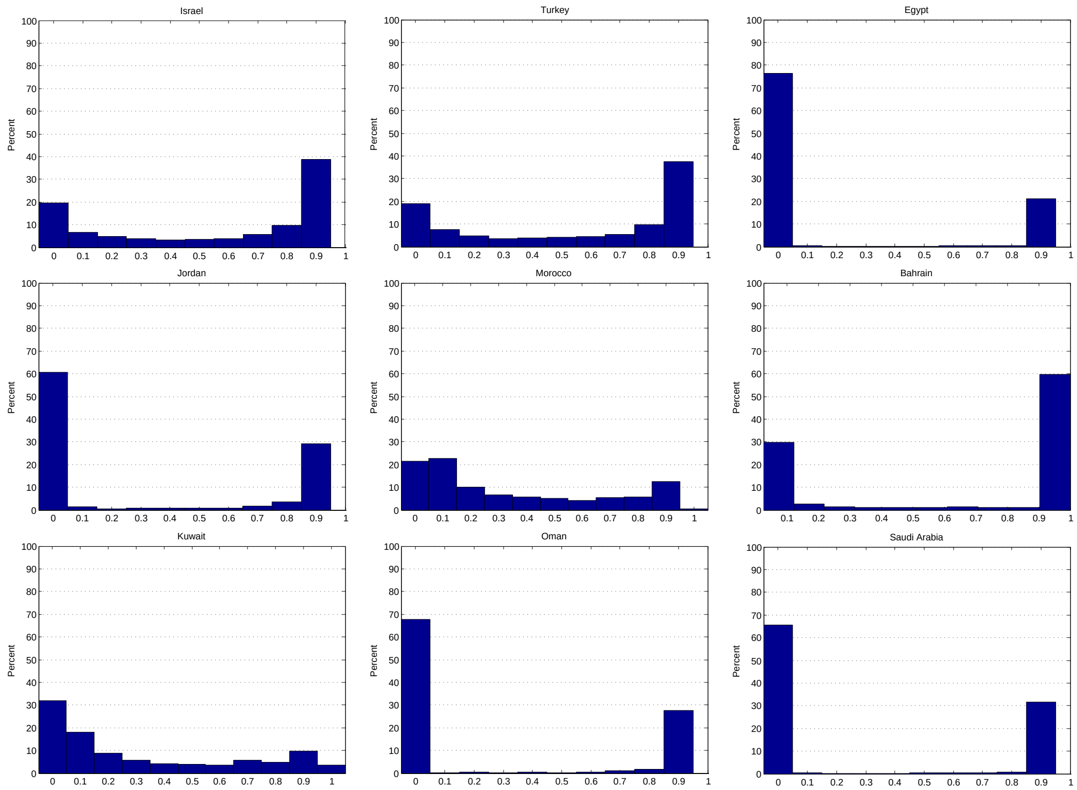
[15]More specifically, we run the test on the residuals obtained by fitting a low-order autoregressive process to the data for each frame; there is no evidence of second-order correlation at conventional significance levels for each residual series.

[16]The code is available at: `http://www.gov.utexas.edu/hinich/files/T23/`

[17]The URL for these tables is: `http://www.duke.edu/~charvey/Country_risk/couindex.htm`

[18]Further details are available upon request.

Figure 1: Histograms of Estimated Daily Integration Probabilities

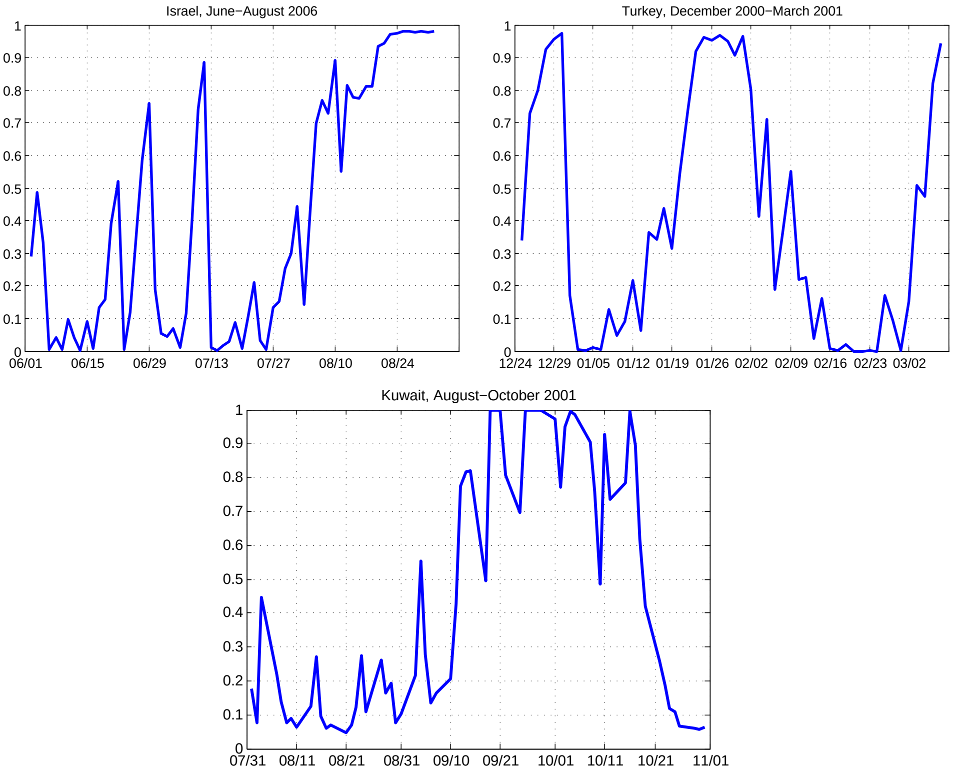


Notes: Each plot is a histogram of $\hat{\phi}^i_{t-1}$ values obtained from maximum likelihood estimation for each market of:

$$r^i_t = \phi^i_{t-1}\lambda \text{Cov}_{t-1}[r^i_t, r^W_t] + (1 - \phi^i_{t-1})\lambda^i \text{Var}_{t-1}[r^i_t] + \varepsilon^i_t,$$

where r^i_t is the market excess return in country i, r^W_t is the world market excess return, $\phi^i_{t-1} = \text{Prob}(S^i_t = 1|\mathcal{F}_{t-1})$, $S^i_t = 1$ denotes that market i is integrated with international equity markets in observation t, $S^i_t = 2$ denotes it is segmented, $\mathcal{F}_{t-1}$ is the observation $t-1$ information set, and both λ and λ^i, the risk premia associated with, respectively, world market systematic risk and country-specific idiosyncratic risk, were restricted to be positive.

Figure 2: Sub-Sample Periods of $\hat{\phi}^i_{t-1}$ for Israel, Turkey, and Kuwait



Notes: See notes to Figure 1 for explanation of $\hat{\phi}^i_{t-1}$. These time series plots are presented to demonstrate how the estimated probability of market i being integrated with international equity markets varied: in Israel around the time of the Lebanon/Hezbollah War of 2006; in Turkey during its exchange-rate crisis of 2000-01; and in Kuwait prior to and following the terrorist attacks of September 11^{th} and during the subsequent start of the US invasion of Afghanistan.

Table 1: Sample Statistics for Daily Excess Returns

	Dow Jones World	Bahrain	Kuwait	Oman	Saudi Arabia	Egypt	Israel	Jordan	Morocco	Turkey
Mean	3.96e-05	2.91e-4	7.66e-4	6.64e-4	5.23e-4	2.65e-4	3.41e-4	4.78e-4	3.81e-4	2.86e-4
Standard Deviation	0.009	0.005	0.008	0.008	0.015	0.015	0.014	0.012	0.009	0.034
Skewness	-0.214	0.227	-0.357	0.043	-1.107	-0.218	-0.456	0.048	0.004	-0.132
Kurtosis	5.029	11.363	10.329	20.226	22.559	9.344	8.592	13.100	7.061	9.282

Notes: The excess returns series were computed by subtracting, for each observation, the daily 3-month secondary market US T-bill rate from the log difference of the market total return index in each country. The last observation for each series is February 15, 2008. For Egypt, Jordan, Israel, Morocco, and Turkey, the initial observation is July 7, 1997. The data start on March 1, 2000 for Bahrain, Kuwait, and Saudi Arabia, and the first observation for Oman is July 17, 2000.

Table 2: Static International CAPM and Factor Model Results

		Bahrain	Kuwait	Oman	Saudi Arabia	Egypt	Israel	Jordan	Morocco	Turkey
International CAPM	$\hat{\alpha}$	2.92e−4‡	7.66e−4*	6.63e−4*	5.22e−4	2.65e−4	3.40e−4	4.78e−4‡	3.81e−4†	2.82e−4
		(1.43e−4)	(1.93e−4)	(2.27e−4)	(3.58e−4)	(3.24e−4)	(2.72e−4)	(2.39e−4)	(2.07e−4)	(6.32e−4)
	$\hat{\beta}$	0.028†	−0.009	−0.003	−0.020	0.040	0.164*	−0.026	0.028	0.522*
		(0.014)	(0.018)	(0.026)	(0.029)	(0.040)	(0.043)	(0.025)	(0.023)	(0.103)
Factor Model	$\hat{\alpha}$	2.81e−4†	8.04e−4*	0.001*	8.11e−4†	5.92e−4	4.01e−4	6.95e−4‡	7.04e−4*	7.18e−4
		(1.68e−4)	(2.44e−4)	(2.78e−04)	(4.44e−4)	(3.83e−4)	(2.98e−4)	(2.93e−4)	(2.48e−4)	(7.29e−4)
	$\hat{\beta}$	0.027 †	−0.011	−5.22e−4	−0.018	0.032	0.159*	−0.029	0.027	0.516*
		(0.014)	(0.019)	(0.025)	(0.030)	(0.038)	(0.043)	(0.025)	(0.023)	(0.102)
	$\hat{\delta}_1$	0.013*	−0.003	0.012	−0.005	0.023†	0.007	0.004	0.003	−0.010
		(0.005)	(0.007)	(0.008)	(0.016)	(0.014)	(1.22e−2)	(0.009)	(0.008)	(0.034)
	$\hat{\delta}_2$	0.008	0.105	−0.368†	−0.120	0.003	0.345‡	−0.062	−0.056	0.280
		(0.066)	(0.091)	(0.207)	(0.287)	(0.192)	(0.141)	(0.123)	(0.092)	(0.618)
	$\hat{\delta}_3$	−1.428†	−0.446	−0.723	−0.894	−2.466	−0.670	−1.544	−1.340	−2.814
		(0.758)	(1.045)	(1.564)	(1.509)	(2.075)	(2.169)	(1.112)	(1.233)	(4.234)
	$\hat{\gamma}_1$	−8.59e−4	2.93e−4	−0.003 *	−8.50e−4	−0.002	−0.004†	6.52e−5	−0.003*	−0.005
		(0.001)	(4.71e−4)	(7.39e−4)	(0.001)	(0.001)	(0.002)	(6.37e−4)	(8.61e−4)	(0.005)
	$\hat{\gamma}_2$	−1.16e−4	−1.35e−4	−0.001	0.001	−0.002†	0.002	1.38e−4	−0.001	0.003
		(0.001)	(6.70e−4)	(0.002)	(9.27e−4)	(8.83e−4)	(0.001)	(0.001)	(0.002)	(0.004)
	$\hat{\gamma}_3$	0.002‡	−4.77e−4	0.001	−0.001		−0.004	3.57e−4	−0.002	−0.005
		(0.001)	(9.36e−4)	(8.62e−4)	(0.002)		(0.003)	(0.001)	(0.001)	(0.008)
	$\hat{\gamma}_4$	0.001†	−9.28e−4	−6.47e−4	3.04e−4		−0.002	−0.001	−0.001	−0.007
		(6.36e−4)	(0.002)	(4.97e−4)	(0.003)		(0.002)	(0.001)	(9.13e−4)	(0.007)
	$\hat{\gamma}_5$	−2.29e−4	7.50e−5		−0.006†		−4.64e−4	−0.002*	−6.40e−4	−2.33e−4
		(4.21e−4)	(7.53e−4)		(0.003)		(0.001)	(5.63e−4)	(6.40e−4)	(0.004)
	$\hat{\gamma}_6$	9.22e−4	−0.001						−8.14e−4	0.002
		(0.001)	(0.002)						(6.66e−4)	(0.003)
	$\hat{\gamma}_7$	0.001							−3.50e−4	4.42e−4
		(9.76e−4)							(0.003)	(0.002)
	$\hat{\gamma}_8$									−0.004
										(0.003)

Notes: Newey-West HAC consistent standard errors appear in parentheses. *, ‡, and † denote rejection of the null hypothesis that the parameter equals zero at the 1%, 5%, and 10% significance levels, respectively. The estimated parameters were obtained by applying OLS to, respectively, $r_t^i = \alpha_i + \beta r_t^W + \varepsilon_t^i$ and $r_t^i = \alpha + \beta r_t^M + \sum_{j=1}^{3} \beta_j F_{t,j} + \sum_{k=1}^{M} \gamma_k d_{t,k}^i + \varepsilon_t^i$, equations (1) and (2), where r_t^i is the market excess return in country i, r_t^W is the world market excess return, the factors $F_{t,j}$ are the log differences in the daily spot oil price ($j = 1$), the squared log differences in the spot oil price ($j = 2$), and the squared world market excess returns ($j = 3$), the $d_{t,k}^i$ variables represent the Hinich factors, and ε_t^i is assumed to be a white noise innovation process.

Table 3: Intertemporal CAPM GARCH-M Results

	Bahrain	Kuwait	Oman	Saudi Arabia	Egypt	Israel	Jordan	Morocco	Turkey
EGARCH-M									
$\hat{\mu}$	$-8.89e-4^{\star}$	$1.20e-3^{\star}$	$-2.00e-4$	$5.71e-4^{\star}$	$-1.53e-3^{\star}$	$6.17e-4^{\star}$	$-5.43e-4^{\star}$	$-1.97e-4$	$2.28e-3^{\star}$
	(5.60e−5)	(1.33e−4)	(6.13e−4)	(1.70e−4)	(3.95e−4)	(2.33e−4)	(1.39e−4)	(2.72e−4)	(7.99e−4)
$\hat{\lambda}$	$40.317^{\star}$	-2.489	11.895	$3.056^{\star}$	$6.881^{\star}$	-0.221	$6.808^{\star}$	5.920	-1.321
	(3.522)	(3.935)	(9.651)	(0.241)	(1.966)	(0.140)	(2.306)	(3.957)	(0.911)
$\hat{\omega}$	$-1.005^{\star}$	$-0.730^{\star}$	$-0.589^{\star}$	$-0.039^{\star}$	$-0.091^{\star}$	$-0.500^{\star}$	$-0.041^{\star}$	$-0.603^{\star}$	$-0.131^{\star}$
	(0.198)	(0.149)	(0.125)	(0.017)	(0.021)	(0.080)	(0.0140)	(0.131)	(0.036)
$\hat{\alpha}$	$0.188^{\star}$	$0.201^{\star}$	$0.142^{\star}$	$0.174^{\star}$	$0.122^{\star}$	$0.196^{\star}$	$0.080^{\star}$	$0.310^{\star}$	$0.189^{\star}$
	(0.026)	(0.020)	(0.019)	(0.012)	(0.013)	(0.019)	(8.75e−3)	(0.030)	(0.017)
$\hat{\beta}$	$0.900^{\star}$	$0.923^{\star}$	$0.936^{\star}$	$0.992^{\star}$	$0.987^{\star}$	$0.941^{\star}$	$0.992^{\star}$	$0.936^{\star}$	$0.980^{\star}$
	(0.019)	(0.015)	(0.013)	(1.96e−3)	(2.55e−3)	(9.33e−3)	(1.58e−3)	(0.014)	(5.11e−3)
$\hat{\theta}$	$0.283^{\star}$	$-0.247^{\star}$	$0.147^{\dagger}$	$0.134^{\star}$	$0.191^{\star}$	$-0.455^{\star}$	$0.543^{\star}$	0.055	$-0.187^{\star}$
	(0.060)	(0.070)	(0.077)	(0.043)	(0.050)	(0.072)	(0.076)	(0.040)	(0.051)
GJR GARCH-M									
$\hat{\mu}$	$-6.85e-3^{\ddagger}$	$5.89e-4^{\dagger}$	$1.32e-5$	$5.86e-4^{\ddagger}$	$-8.23e-4^{\ddagger}$	$7.90e-4^{\dagger}$	$-3.62e-4$	$-1.44e-4$	$1.59e-3^{\dagger}$
	(3.27e−4)	(3.26e−4)	(4.72e−4)	(2.38e−4)	(4.11e−4)	(4.60e−4)	(2.41e−4)	(2.59e−4)	(8.37e−4)
$\hat{\lambda}$	$30.314^{\star}$	3.907	8.721	0.901	$4.787^{\ddagger}$	-1.555	$5.292^{\ddagger}$	5.339	-0.936
	(11.56)	(5.022)	(7.616)	(1.602)	(2.314)	(2.793)	(2.305)	(3.646)	(0.904)
$\hat{\omega}$	$3.79e-6^{\star}$	$4.10e-6^{\star}$	$3.91e-6^{\star}$	$1.20e-6^{\star}$	$1.80e-6^{\star}$	$9.47e-6^{\star}$	$5.27e-7^{\star}$	$6.09e-6^{\star}$	$2.13e-5^{\star}$
	(3.10e−7)	(3.44e−7)	(4.78e−7)	(1.64e−7)	(2.38e−7)	(1.30e−6)	(6.80e−8)	(5.98e−7)	(2.96e−6)
$\hat{\alpha}$	$0.142^{\star}$	$0.086^{\star}$	$0.065^{\star}$	$0.098^{\star}$	$0.061^{\star}$	$0.047^{\star}$	$0.051^{\star}$	$0.202^{\star}$	$0.073^{\star}$
	(0.011)	(0.010)	(0.009)	(6.60e−3)	(4.67e−3)	(0.010)	(2.54e−3)	(0.016)	(6.78e−3)
$\hat{\beta}$	$0.786^{\star}$	$0.843^{\star}$	$0.888^{\star}$	$0.926^{\star}$	$0.944^{\star}$	$0.850^{\star}$	$0.963^{\star}$	$0.742^{\star}$	$0.889^{\star}$
	(0.012)	(0.009)	(0.012)	(3.99e−3)	(3.81e−3)	(0.014)	(1.67e−3)	(0.015)	(6.32e−3)
$\hat{\gamma}$	$-0.087^{\star}$	$0.031^{\star}$	$-0.020^{\star}$	$-0.041^{\star}$	$-0.022^{\star}$	$0.105^{\star}$	$-0.036^{\star}$	-0.025	$0.047^{\star}$
	(0.011)	(0.011)	(0.007)	(5.93e−3)	(4.58e−3)	(0.015)	(2.89e−3)	(0.018)	(0.011)

Notes: Standard errors appear in parentheses. *, ‡, and † denote rejection of the null hypothesis that the parameter equals zero at the 1%, 5%, and 10% significance levels, respectively. The estimated parameters were obtained by maximum likelihood. In each case, the conditional mean equation is given by $r_t^i = \mu + \lambda h_t^i + \varepsilon_t$, where r_t^i is the market excess return in country i, $\varepsilon_t = \sqrt{h_t} e_t$, $e_t \sim N(0,1)$, and h_t^i is the conditional variance of the market excess return in country i. In the EGARCH-M model, the (natural logarithm of the) conditional variance is given by $\ln(h_t^i) = \omega + \alpha g(z_{t-1}) + \beta \ln(h_{t-1}^i)$, where $g(z_t) = \theta z_t + \delta[|z_t| - \mathbb{E}|z_t|]$, $z_t = \varepsilon_t/\sqrt{h_t}$, and $\delta = 1$. In the GJR GARCH-M model, the conditional variance is given by $h_t^i = \omega + \alpha\varepsilon_{t-1}^2 + \gamma I_{\{\varepsilon_{t-1}<0\}}\varepsilon_{t-1}^2 + \beta h_{t-1}^i$, where $I_{\{\varepsilon_{t-1}<0\}}$ is an indicator function which takes on the value of 1 when $\varepsilon_{t-1} < 0$ and 0 otherwise.

Table 4: Markov-Switching Intertemporal CAPM Results

	Bahrain	Kuwait	Oman	Saudi Arabia	Egypt	Israel	Jordan	Morocco	Turkey
Sample Statistics of $\hat{\phi}_t^i$									
Mean	0.660	0.365	0.297	0.326	0.220	0.598	0.342	0.394	0.592
Median	1.0	0.200	0.0	0.0	0.0	0.783	0.0	0.253	0.754
Standard Deviation	0.434	0.337	0.439	0.458	0.404	0.381	0.443	0.327	0.377
Transition Probabilities									
$\hat{p}_{1,1}^i$	0.65 (0.16)	0.85 (0.03)	0.21 (0.03)	0.31 (0.02)	0.11 (0.02)	0.97 (0.01)	0.29 (0.03)	0.83 (0.03)	0.95 (0.01)
$\hat{p}_{2,2}^i$	0.22 (0.32)	0.95 (0.01)	0.70 (0.02)	0.68 (0.02)	0.77 (0.01)	0.94 (0.01)	0.68 (0.02)	0.93 (0.01)	0.98 (0.09)
Risk Premia									
$\hat{\lambda}$	2.458 [0.607]	1.383 [0.317]	2.69e−6 [0.206]	0.346 [0.655]	0.058 [0.206]	7.731 [4.59e−6]	0.186 [0.654]	5.011 [0.020]	6.337 [2.0e−5]
$\hat{\lambda}^i$	2.379 [0.138]	8.765 [1.0e−5]	7.766 [0.001]	0.789 [0.083]	0.951 [0.237]	0.548 [0.371]	1.848 [0.527]	3.886 [0.023]	0.033 [0.527]

Notes: Standard errors appear in parentheses. The results were obtained by maximum likelihood estimation for each market of:

$$r_t^i = \phi_{t-1}^i \lambda \mathrm{Cov}_{t-1}[r_t^i, r_t^W] + (1 - \phi_{t-1}^i)\lambda^i \mathrm{Var}_{t-1}[r_t^i] + \varepsilon_t^i,$$

where r_t^i is the market excess return in country i, r_t^W is the world market excess return, $\phi_{t-1}^i = \mathrm{Prob}(S_t^i = 1|\mathcal{F}_{t-1})$, S_t^i is a state variable which can take on two values, with $S_t^i = 1$ denoting that market i is integrated with international equity markets in observation t and $S_t^i = 2$ denoting it is segmented, $\mathcal{F}_{t-1}$ is the observation $t-1$ information set, $p_{1,1}^i = \mathrm{Prob}(S_t^i = 1|S_{t-1}^i = 1)$ and $p_{2,2}^i = \mathrm{Prob}(S_t^i = 2|S_{t-1}^i = 2)$, and both λ and λ^i, the risk premia associated with, respectively, world market systematic risk and country-specific idiosyncratic risk, were restricted to be positive. In brackets under the estimated risk premia are $p-$values for likelihood ratio tests of the null hypothesis that the risk premium in question equals zero against the alternative that it is positive.

References

Adler, M., Dumas, B., 1983. International Portfolio Choice and Corporation Finance: A Synthesis. Journal of Finance 38, 925–984.

Alper, C. E., 2001. The Turkish Liquidity Crisis of 2000: What Went Wrong? Russian and East European Finance and Trade 37 (6), 51–71.

Bekaert, G., Harvey, C. R., 1995. Time-Varying World Market Integration. Journal of Finance 50 (2), 403–444.

Bekaert, G., Harvey, C. R., 1997. Emerging Equity Market Volatility. Journal of Financial Economics 43, 29–77.

Bekaert, G., Harvey, C. R., Lundblad, C. T., Siegel, S., 2008. What Segments Equity Markets? Working Paper, Columbia Business School.

Binder, J. J., 1998. The Event Study Methodology Since 1969. Review of Quantitative Finance and Accounting 11, 111–137.

Chen, N.-f., Roll, R., Ross, S., 1986. Economic Forces and the Stock Market. Journal of Business 59, 383–403.

Errunza, V., 2001. Foreign Portfolio Equity Investments, Financial Liberalization, and Economic Development. Review of International Economics 9, 703–726.

Errunza, V. R., Losq, E., Padamanabhan, P., 1992. Tests of Integration, Mild Segmentation, and Segmentation Hypothesis. Journal of Banking and Finance 16, 949–972.

Fama, E. F., French, K. R., 1996. Multifactor Explanations of Asset Pricing Anomalies. Journal of Finance 51, 55–84.

Ghysels, E., Cherkaoui, M., 2003. Emerging Markets and Trading Costs: Lessons from Casablanca. Journal of Empirical Finance 10, 169–198.

Glosten, L. R., Jaganathan, R., Runkle, D. E., 1993. On the Relation Between the Expected Value and the Volatility of the Nominal Excess Return on Stocks. Journal of Finance 48, 1779–1802.

Gulen, H., Mayhew, S., 2000. Stock Index Futures, Trading and Volatility in International Equity Markets. The Journal of Futures Markets 20 (7), 661–685.

Hamilton, J. D., 1983. Oil and Macroeconomy Since World War II. Journal of Political Economy 91 (2), 228–248.

Hamilton, J. D., 1985. Historical Causes of Postwar Oil Shocks and Recessions. Energy Journal, 97–116.

Hamilton, J. D., 1996. This Is What Happened to the Oil Price/Macroeconomy Relation. Journal of Monetary Economics 38 (2), 215–220.

Hamilton, J. D., 2003. What Is an Oil Shock? Journal of Econometrics 113, 363–398.

Harvey, C. R., 1991. The World Price of Covariance Risk. The Journal of Finance 46 (1), 111–157.

Harvey, C. R., 1995. Predictable Risk and Returns in Emerging Markets. Review of Financial Studies 8, 773–816.

Henry, P. B., 1999. Chronological Listing of Major Policy Events in Emerging Markets. Working Paper, Stanford Graduate School of Business, 1–43.

Hinich, M. J., 1982. Testing for Gaussianity and Linearity of a Stationary Time Series. Journal of Time Series Analysis 3 (3), 169–176.

Hinich, M. J., 1996. Testing for Dependence in the Input to a Linear Time Series Model. Nonparametric Statistics 6 (2 & 3), 205–221.

Hinich, M. J., Patterson, D. M., 2005. M. Belongia and J. Binner (ed.), Money Measurement and Computation. Palgrave, London, Ch. Detecting Epochs of Transient Dependence in White Noise, 61-75.

Hinich, M. J., Serletis, A., 2007. Episodic Nonlinear Event Detection in the Canadian Exchange Rate. Journal of the American Statistical Association 102 (477), 68–74.

Kandel, S., Stambaugh, R. F., 1990. Expectations and volatility of consumption and asset prices. Review of Financial Studies 3, 207–232.

Kim, E. H., Singal, V., 2000. Stock Market Opennings: Experience of Emerging Markets. Journal of Business 73, 25–66.

Lintner, J., 1965. The Valuation of Risk Assets and the Selection of Risky Investments in Stock Portfolios and Capital Budgets. Review of Economics and Statistics 47, 13–37.

Merton, R. C., 1973. An Intertemporal Capital Asset Pricing Model. Econometrica 41, 867–887.

Merton, R. C., 1980. On Estimating the Expected Return on the Market: An Explanatory Investigation. Journal of Financial Economics 8, 323–361.

Nelson, D. B., 1990. Conditional Heteroscedasticity in Asset Returns: A New Approach. Econometrica 59, 347–370.

Newey, W. K., West, K. D., 1987. A Simple, Positive Semi-definite, Heteroskedasticity and Autocorrelation Consistent Covariance Matrix. Econometrica 55, 703–708.

Pan, J., 2002. The Jump-Risk Premia Implicit in Options: Evidence From an Integrated Time-Series Study. Journal of Financial Economics 63, 3–50.

Sharpe, W. F., 1964. Capital asset prices: A Theory of Market Equilibrium under Conditions of Risk. Journal of Finance 19, 425–442.