An Empirical Investigation of Stock Market Behavior in the Middle East and North Africa

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Abstract

This paper analyzes 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.

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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 331st in the Fortune Global 500 list, with an estimated revenue of USD 20.9 billion and equity worth USD 16.6 billion, ranking 6th among international chemical producers.\(^1\) Koç Holdings of Turkey, with over USD 18 billion in revenue in 2006, is ranked 358th in the Fortune Global 500. The Israeli hi-tech industry, nicknamed ‘Silicon Wadi,’ is considered second in global influence after California’s Silicon Valley.\(^2\) Dubai in the United Arab Emirates (UAE) is trying to position itself as an important financial center between Hong Kong and London; it was top-ranked in the September 2008 Global Financial Centres Index survey question on where financial firms would like to establish new offices. 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

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\(^1\)Sources: the SABIC 2006 annual report and the 2006 Fortune Global 500.

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 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 Global Index (DJG).³

Our check of CAPM efficiency and financial market integration in this setup is based upon examination of estimates of, respectively, alpha and beta.⁴ 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 (ICAPM) 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.

³The DJG is a broad measure of the global stock market which targets 95% coverage of markets open to foreign investment. It tracks 46 countries, of which 24 are developed markets and 22 are emerging markets. More details can be found at http://www.djindexes.com/globalindexes/.

⁴If 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.
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.

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. Table 1 lists some descriptive statistics for the economies of these countries. Bahrain, Kuwait, Oman, and Saudi Arabia, represent oil exporting and rich Persian Gulf basin markets; each is classified by the World Bank as a high-income country. Israel and Turkey, the two predominantly non-Arab countries we study, are, respectively, an OECD ‘accession candidate’ country and an OECD member country; the first is a high-income and the second is an upper-middle-income country. Jordan and Egypt, both non oil-exiting lower-middle-income countries, have strong trade and financial ties to the Persian Gulf region oil exporters; for example, remittances from Egyptian workers in Gulf Cooperation Council (GCC) countries account for 50% of total remittances from Egyptians working abroad and such remittances from Jordanian workers comprise roughly 14% of Jordanian GNP.5 Morocco is a representative lower-middle-income North African country, but in many ways its market is more integrated with Europe than with the rest of the MENA countries; for example, two-thirds of Morocco’s balance of payments is linked to transactions with the European Union.6

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

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are reported by Standard and Poor's/IFCG. The same data for Israel and our proxy for the world financial market index, the DJG 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 2. 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 DJG. This, along with the MENA estimated unconditional second moments being of the same order of magnitude as for the DJG, 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 unit root tests, the augmented Dickey-Fuller and Elliott et al. (1996) tests, on the data. The associated test statistics are also reported in Table 2. The results show 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 excess 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 such as ours, we regress the excess returns of the national market against the excess returns of an index composite of international markets, the DJG.
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. We postulate that oil prices have an impact on market performance in the most important oil-producing region in the world; we use both the growth rate and the squared growth rate of oil prices. In addition, we also test for the possibility of a relationship between expected local excess returns and the squared values of world excess returns. Inclusion of this variable may capture some nonlinear departure from the traditional international CAPM. Further, we 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 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.”

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

\[
 r_i^t = \alpha + \beta r_W^t + \epsilon_i^t \tag{1}
\]

where \( r_i^t \) is the market excess return in country \( i \), \( r_W^t \) is the world market excess return, and \( \epsilon_i^t \) is assumed to be a white noise innovation process. As mentioned earlier, we use the DJG 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 are presented in the first two rows of Table 3. 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 signif-

\[7\] In many MENA markets, there is no concept of a domestic corporate debt market.

\[8\] We outline this procedure and explain why we favor it in the Appendix, where we also define the Hinich factor dummy variables we use.
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: Egypt (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 $\hat{\beta}$ for Egypt is statistically significant, we feel it is not economically significant since its value is rather small (0.067). The value of $\hat{\beta}$ for Israel (0.594) is economically significant and implies less volatility relative to the world capital market. In the case of Turkey, the large value of $\hat{\beta}$ (1.014) implies strong movement with world equity markets.

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 factor model is given by:

$$r^i_t = \alpha + \beta r^M_t + \sum_{j=1}^{3} \delta_j F_{t,j} + \sum_{k=1}^{M} \gamma_k d^k_{t,k} + \epsilon^i_t,$$

(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 $d^k_{t,k}$ variables represent the Hinich factors.$^9$

The results are reported in the third through last rows in Table 3. 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. Three exceptions concern Egypt, Israel, and Saudi Arabia: use of the factor variables leads to the conclusion that these markets, in contrast to the outcome under a simpler specification, are CAPM inefficient. The Turkish market continues to be CAPM-efficient, and the $\hat{\beta}$’s for both Israel and Turkey 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 as well as CAPM inefficient.

Few of the factor variables are significant at conventional significance levels. We have two observations on this. First, given the important role that roughly half of the MENA countries play in the world oil market, we find it surprising that, based upon our estimated models, the growth

$^9$We used West Texas Intermediate spot oil prices from the US Department of Energy’s database as our oil price measure.
rate of the price of oil is apparently not conditionally correlated with aggregate equity returns in
these markets; Egypt, which is not a large exporter of oil, is the only country for which the p-value
for the null hypothesis that the coefficient on the log difference in the spot oil price equals zero
is less than 0.10.\textsuperscript{10} Second, the variability of the price of oil is a statistically significant factor for
three of the four GCC countries, Bahrain, Kuwiat, and Oman, as well as for Egypt. Third, the
variability of world market excess returns appears to be strongly correlated with market excess
returns in Egypt, Morocco, and Turkey. Fourth, the coefficients on only 7 out of the 48 Hinich
factors used are statistically significant. The framework we use in Section 5 provides an alternative
approach for modeling possible nonlinearity in the excess returns in these markets.

As a robustness check, we also consider specifications in which we augment equation (2) by
including the following additional factors: (1) variants of Hamilton’s (1996) ‘net oil price increase’
(NOPI) measures; (2) log differences of bilateral exchange rates against the USD; and (3) following
Hamilton (1996), what we call ‘net exchange rate increase’ (NERI) measures.\textsuperscript{11} The idea behind
inclusion of these measures is that they may serve as proxies for oil price and exchange rate surprises;
our use of exchange rate log first differences as exchange rate surprise measures is motivated by
the voluminous literature, following Meese and Rogoff (1983), which has documented the difficulty
of beating random walk out-of-sample forecasts of exchange rate movements. Regarding the key
question of segmentation from versus integration with international capital markets, in none of the
many additional regressions run are the results different from those described above. That is, the
\( \hat{\beta} \)’s for the Israeli and Turkish markets are both statistically and economically significant, while
none of the \( \hat{\beta} \)’s for the other MENA markets meet these two criteria. With respect to the issue
of CAPM-efficiency, in most cases the \( \hat{\alpha} \)’s are both statistically and economically significant.\textsuperscript{12}
Further, in most cases the oil price and exchange rate surprise proxy variables we use are not

\textsuperscript{10}This set of results appears to be consistent with work, such as Blanchard and Galí (2008), showing that the
macroeconomic effects of oil price movements in the first decade of the 21\textsuperscript{st} century differ considerably from those
observed in the 1970s. We note, however, that Hamilton (2009), for example, questions the conclusions reached by

\textsuperscript{11}Since we use daily data, whereas Hamilton (1996) used quarterly data, we use shorter time periods over which
to compute the relevant maxima in defining our ‘net increase’ measures. More specifically, we compute the k-day
‘net increase’ (NI\textsubscript{k}) for variable \( x_t \) as: \( NI_k = \max(0, \{\ln(x_t) - \ln[\max(x_{t-k}, x_{t-k-1}, ..., x_{t-1})]\}) \), where \( k = 5, 20, \) and
65, for \( x_t \) = the price of oil and the bilateral exchange rate between the currency of the particular MENA country
and the USD.

\textsuperscript{12}The exceptions are Israel and Turkey. In the case of Israel, when the 5-day NOPI and NERI variables are
added, the results imply CAPM-efficiency; but in all other cases CAPM-inefficiency is implied. In the case of Turkey,
if only the log difference of the exchange rate is added, the results imply CAPM-efficiency; in all other cases, the
results imply CAPM-inefficiency.
statistically significant factors.\textsuperscript{13}

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 ICAPM, the expected conditional excess return for market \( i \) should vary positively with its conditional variance:

\[
E_{t-1}[r^i_t] = \mu + \lambda \text{Var}_{t-1}[r^i_t],
\]

where the parameter \( \lambda \) is the coefficient of relative risk aversion of the representative agent.\textsuperscript{14} \( \lambda \) is also referred to as the risk premium associated with market risk. If the ICAPM holds, then \( \mu = 0 \).

To investigate whether there is a risk-return trade-off of the ICAPM 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:

\[
r^i_t = \mu + \lambda h^i_{t-1} + \epsilon_t,
\]

where \( \epsilon_t = \sqrt{h_t} \epsilon_t, \epsilon_t \sim N(0,1) \), and \( h^i_t \) is the conditional variance of \( r^i_t \). The first GARCH conditional volatility structure we use is the Exponential GARCH (EGARCH) model of Nelson (1990):

\[
\ln(h^i_t) = \omega + \alpha g(z_{t-1}) + \beta \ln(h^i_{t-1})
\]

\[
g(z_t) = \theta z_t + \delta [z_t - \mathbb{E}[z_t]],
\]

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

\textsuperscript{13}Full details are available upon request, but here we list the variables for which the \( p \)-value on the estimated coefficient is less than 0.10 for the null hypothesis that the coefficient equals zero: for Bahrain, the 65-day \textit{NOPI}, as well as the 5-day and 20-day \textit{NERI}; for Egypt, both the 20-day \textit{NOPI} and \textit{NERI}; for Israel, the 5-day and 65-day \textit{NERI}; for Saudi Arabia, the 20-day \textit{NERI}; for Turkey, the 5-day and 65-day \textit{NERI}.

\textsuperscript{14}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).
Our second GARCH specification follows Glosten et al. (1993) (GJR):

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

(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 4.

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 \( \lambda \)'s are also significantly positive for Bahrain, Egypt, and Jordan, but not for Saudi Arabia. Both the EGARCH-M and GJR GARCH-M \( \lambda \)'s are economically reasonable for Egypt (6.881 and 4.787) and Jordan (6.808 and 5.292), while those for Bahrain appear to be too high to be economically meaningful (40.317 and 30.314).\(^{15}\) The EGARCH-M \( \lambda \) 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 \( \lambda \)'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 \( \mu \) is insignificant for Jordan, but the EGARCH-M estimated intercept for Jordan is significant. For all other MENA markets, \( \mu \) 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; we explore this below. This may also be due to compensation for jump risk; see, for example, Pan (2002).

The strongest evidence in favor of the ICAPM 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 \( \mu \). Holding constant the statistically significant intercepts, our positive risk-return trade-off results also support the ICAPM 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 ICAPM 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

\(^{15}\)The arguments of Kandel and Stambaugh (1990), however, imply that the \( \lambda \)'s for Bahrain may not be too high to make economic sense.
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., $\theta$ in equation (6) and $\gamma$ 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.

As a check of possible misspecification of the conditional mean in equation (4), we also estimate a large set of additional EGARCH-M and GJR GARCH-M models in which we use several of the factors considered in Section 3: log first differences of oil prices and bilateral exchange rates against the USD, as well as the various $k$-day NOPI and NERI measures. In these additional models, we do not observe any differences in the statistical significance/insignificance of the intercept terms relative to what we obtain with our more parsimonious specification discussed above. Further, in most cases there is no change in the evidence on the risk-return trade-off question.\footnote{The exceptions are as follows. First, when we augment the EGARCH-M model for the Moroccan market by including the 20-day NOPI and NERI variables, the $\hat{\lambda}$ of 6.1 we obtain is statistically significant at the 1% significance level. This point estimate for $\lambda$ is quite close to its analogue reported in Table 4, but the $p$–value for the zero null hypothesis in that case is greater than 0.10. Second, each time we add one of the three pairs of $k$–day NOPI and NERI measures to the GJR GARCH-M model for Egypt, the $\hat{\lambda}$ is not significant at even the 10% level of significance; for their analogue in Table 4 the $p$–value for the zero null hypothesis is less than 0.05. Full details on this set of additional estimated EGARCH-M and GJR GARCH-M models are available upon request.}

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 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 call this the Markov-switching intertemporal CAPM (MS-ICAPM). 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 $S^i_t$ be a latent state variable for market $i$ which can take on two values, with $S^i_t = 1$ denoting that market $i$ is integrated with international equity markets in observation $t$ and $S^i_t = 2$ denoting...
it is segmented. Define:

\[ \phi_{i,t-1} = \text{Prob}(S^i_t = 1|\mathcal{F}_{t-1}), \]

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

\[ r^i_t = \phi^i_{t-1}\lambda_{t-1}\text{Cov}_{i-1}[r^i_t, r^W_t] + (1 - \phi^i_{t-1})\lambda^i_{t-1}\text{Var}_{i-1}[r^i_t] + \varepsilon^i_t, \tag{9} \]

where \( \lambda_{t-1} \) and \( \lambda^i_{t-1} \) are the time-varying risk premia associated with world market systematic risk and country-specific idiosyncratic risk.\(^\text{17}\) While the MS-ICAPM allows the probability of integration, \( \phi^i_{t-1} \), to vary across time, we assume that the transition probabilities \( p^i_{1,1} = \text{Prob}(S^i_t = 1|S^i_{t-1} = 1) \) and \( p^i_{2,2} = \text{Prob}(S^i_t = 2|S^i_{t-1} = 2) \) are constant. Time variation in the risk premia is allowed as follows:

\[ \lambda_{t-1} = \exp(\psi Z_{t-1}) \tag{10} \]
\[ \lambda^i_{t-1} = \exp(\psi^i Z^i_{t-1}) \tag{11} \]

where \( \psi \) and \( \psi^i \) are parameter vectors, and \( Z_t \) and \( Z^i_t \) 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 \( \lambda \) and \( \lambda^i \) 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, by maximum likelihood assuming, as in Bekaert and Harvey (1995), that the stochastic error terms in equation (9) are normally distributed. Estimation is carried out in two stages. First, we compute \( \text{Var}_{t-1}[r^i_t] \) and \( \text{Cov}_{t-1}[r^i_t, r^W_t] \) using a rolling window estimation scheme.\(^\text{18}\) Second, we form the likelihood function according to the model in equation (9) and maximize it. The sampling error of the first-stage estimates is not accounted for in the standard errors of the second stage, such that

\(^\text{17}\)Following Bekaert and Harvey (1995), we do not include an intercept term in equation (9).

\(^\text{18}\)We fix a sub-sample period of \( m \) days for calculating the variance of \( r^i_t \) and the covariance between \( r^i_t \) and \( r^W_t \), 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.
these standard errors may be understated. To avoid local optima, we perturb our starting values and re-estimate the model 50 times for each market.

Bekaert and Harvey (1995) impose the restriction that the global price of risk is the same across countries, such that there is a unique price for the same source of risk. However, since the sample sizes we have vary across the markets we study, we do not do this, which allows us to use all of our data for estimation of the model on each market. Further, this facilitates comparison with our static international CAPM and constant-parameter ICAPM results.

Following Bekaert and Harvey (1995), we use a set of global and local instrumental variables as components of, respectively, $Z_t$ and $Z^i_t$, 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 DJG 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. 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.

We find that including $Z_t$ and $Z^i_t$, and hence allowing for time-varying premia, does not improve the MS-ICAPM estimation results significantly. In fact, in several cases there are problems with the size of the estimated parameters. 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., $\phi^i_{t-1}$ 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).

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19 In both Andreou and Ghysels (2002) and Ghysels et al. (2005), for example, the sampling error from first-stage rolling window computed sample moments is also ignored in second-stage estimation.

20 For the two countries for which the MS-ICAPM evidence most strongly suggests integration with world capital markets, Israel and Turkey, the estimates of the global price of risk are quite close, less than one standard error apart. Further details available upon request.

21 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.

22 More specifically, many of the elements of the parameter vectors $\hat{\psi}$ and $\hat{\psi}^i$ are unreasonably large in magnitude, as are the associated standard errors, “blowing up” in both the positive and negative directions. We speculate that this is due to the state variables in $Z_t$ and $Z^i_t$ being noisy and thus uninformative. In light of these results, we decided not to pursue estimation of an extension of the model in which we allow the transition probabilities to be time-varying, since, if we were to follow Bekaert and Harvey (1995), a subset of presumed noisy $Z^i_t$ would be used to capture time variation in the transition probabilities.
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 $\hat{\phi}_{it-1}$ 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 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 $\hat{\phi}_{it-1}$ values presented in the top panel Table 5 support these conclusions. The median of $\hat{\phi}_{it-1}$ 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 $\hat{\phi}_{it-1}$ in each case is at the low polar value of 0.0, implying full weight on local pricing of risk at the median. For Kuwait and Morocco, the median of $\hat{\phi}_{it-1}$ is, respectively, 0.200 and 0.253, indicating a more so intermediate case.

To help interpret these results, it is useful to consider data on foreign direct investment (FDI) and cross-border merger and acquisition (M&A) activity for the MENA economies. In 2007, the sum of the stock of inward and outward FDI comprised approximately 105% and 64% of GDP for, respectively, Bahrain and Israel; in absolute terms the volume of both FDI and M&A activity is higher for Turkey than it is for Israel. In contrast, for Kuwait and Saudi Arabia in 2007 the sum of the stock of inward and outward FDI equaled, respectively, 14% and 26% of GDP. While all of the countries of the MENA region attract inward FDI, most of those for which our MS-ICAPM results imply median local pricing of risk carry out, as a fraction of GDP, rather low levels of outward FDI and cross-border M&A investment. Accordingly, the median values we obtain for $\hat{\phi}_{it-1}$ across the MENA countries appear to be consistent with the following: the higher is the volume FDI (especially outward FDI) and cross-border M&A investment flows for a country, the greater is the tendency towards global pricing of risk in its capital market.

Information about the persistence of the unobserved integrated and segmented states is provided by the estimates of the MS-ICAPM transition probabilities $p_{1,1}^i$ and $p_{2,2}^i$ in the middle panel of Table 5. 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 Jordan.

The bottom panel of Table 5 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. 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 substantially 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 MS-ICAPM 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, in 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 13th, 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 25th, 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, possibly 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 14th. 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

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24The 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.
institutional factors, the program faced severe problems in November and December of 2000; for more details, see Alper (2001). On December 1st, 2001, the overnight interbank interest rate reached 1,700%. By December 5th, the financial system was about to collapse. As a result, the IMF extended a rescue package worth USD 10 billion on December 6th, 2000. As is seen in Figure 2, Turkey’s financial markets generally seem to have been integrated moving towards the end of December. The sharp drop in $\hat{\phi}_{t-1}$ 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 19th and February 2nd 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. 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 5 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 11th, 2001, up through the initial phase of the following US-led invasion of Afghanistan. Recall that our earlier analysis implies Kuwaiti financial markets, 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 11th attacks, pricing of risk in Kuwait was generally local; on almost every day, $\hat{\phi}_{t-1}$ was considerably below 0.5. Then, immediately after the September 11th 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 7th, 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

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25Exacerbating the sense of crisis in the country, a rather public row between President Ahmet Necdet Sezer and Prime Minister Bulent Ecevit broke out on February 19th, 2001.
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 MS-ICAPM analysis. That said, pricing in these markets is done locally on roughly twenty percent of the trading days in our sample. Our constant-parameter ICAPM results suggests there is no risk-return trade-off in the Israeli and Turkish markets. This is supported to some extent by our MS-ICAPM evidence 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, as per our static international CAPM and MS-ICAPM results, 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 MS-ICAPM 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 ICAPM analysis or our MS-ICAPM exercise; our results on this question generated by these two different approaches, however, are not consistent for these MENA markets in most cases.

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

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26 While neither of these specifications is likely to be the “true” model, we note that each implies the other model is misspecified.

27 This assertion is based on the following results. We find evidence in favor of a significantly positive value of the EGARCH-M parameter $\lambda$, but MS-ICAPM local risk premia that are not significantly different from zero, for Bahrain, Egypt, and Jordan. For Kuwait, Oman, and Morocco, we find the opposite. There is evidence in favor of a positive risk-return trade-off through both the ICAPM and MS-ICAPM approaches only for Saudi Arabia.

28 What we have in mind is the following. For most of the sample period we study, oil prices were increasing. This fact, combined with our finding that most of the Arabic MENA markets were segmented from international capital markets at the time, is the basis of our idea that there is an inverse relationship between financial market integration for these economies and oil price increases. We believe this is further supported by the simultaneous crash across world markets which followed the outbreak of the September 2008 financial crisis and the subsequent collapse of the price of oil.
Appendix: Episodic Nonlinear Event Detection

To produce additional explanatory variables for our multifactor CAPM regressions, 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 \( \ell \). 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 \( n \)th frame by \( \{ z^n_t \} \). The goal is to detect evidence in favor of third-order correlation in the \( n \)th frame using the Hinich (1996) portmanteau bicorrelation test statistic:

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

\[
B_n(r, u) = \sum_{t=1}^{\ell-r} z^n_t z^n_{t+r} z^n_{t+u}
\]

Under the null hypothesis that the observed process is pure white noise (i.i.d.), if \( \ell \) 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. 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.

Call each of these \( M \) frames a “significant frame” in which there is, following Hinich and Serletis (2007), a “nonlinear event.”

29 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.
For each country \( i \) we create “Hinich factor” dummy variables, \( d_{i,t,k}^i \), \( k = 1, \ldots, M \), corresponding to each of the \( M \) significant frames. The values of these binary variables are determined as follows: if observation \( t \) for country \( i \)’s excess returns series falls in significant frame \( k \) for country \( i \), then \( d_{i,t,k}^i = 1 \); \( d_{i,t,k}^i = 0 \) otherwise.\(^{30}\)

There are several chronological tables available for important or potentially influential financial, economic, and political events in emerging financial markets. Useful references include Henry (1999) and the online tables maintained by Bekaert and Harvey at Duke University; the latter was last updated in September of 2004.\(^{31}\) We find that most of the pre-September 2004 detected significant frames coincide with important events reported in the chronology of Bekaert and Harvey; see Table 6.

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\(^{30}\)The code is available at: http://www.gov.utexas.edu/hinich/files/T23/

\(^{31}\)The URL for these tables is: http://www.duke.edu/~charvey/Country_risk/couindex.htm
Figure 1: Histograms of Estimated Daily Integration Probabilities

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

$$r_i^t = \phi_{it-1} \lambda \text{Cov}_{t-1}[r_i^t, r_w^t] + (1 - \phi_{it-1}) \lambda^i \text{Var}_{t-1}[r_i^t] + \epsilon_i^t,$$

where $r_i^t$ is the market excess return in country $i$, $r_w^t$ is the world market excess return, $\phi_{it-1} = \text{Prob}(S_i^t = 1|\mathcal{F}_{t-1})$, $S_i^t = 1$ 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 $\lambda$ and $\lambda^i$, 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 $\hat{\phi}_{t-1}^i$ for Israel, Turkey, and Kuwait

Notes: See notes to Figure 1 for explanation of $\hat{\phi}_{t-1}^i$. 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 11th and during the subsequent start of the US invasion of Afghanistan.
### Table 1: Summary Statistics for MENA Economies

<table>
<thead>
<tr>
<th></th>
<th>Bahrain</th>
<th>Kuwait</th>
<th>Oman</th>
<th>Saudi Arabia</th>
<th>Egypt</th>
<th>Israel</th>
<th>Jordan</th>
<th>Morocco</th>
<th>Turkey</th>
</tr>
</thead>
<tbody>
<tr>
<td>2008 Per Capita GDP</td>
<td>$37,300</td>
<td>$57,400</td>
<td>$20,200</td>
<td>$20,500</td>
<td>$5,400</td>
<td>$28,300</td>
<td>$5,100</td>
<td>$4,000</td>
<td>$11,900</td>
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<td>Composition of GDP:</td>
<td></td>
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<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>-Agriculture</td>
<td>0.3%</td>
<td>0.3%</td>
<td>2.1%</td>
<td>3.1%</td>
<td>13.4%</td>
<td>2.7%</td>
<td>3.6%</td>
<td>14.7%</td>
<td>8.5%</td>
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<tr>
<td>-Industry</td>
<td>43.6%</td>
<td>52.2%</td>
<td>37.2%</td>
<td>61.6%</td>
<td>37.6%</td>
<td>31.7%</td>
<td>10.1%</td>
<td>38.9%</td>
<td>28.6%</td>
</tr>
<tr>
<td>-Services</td>
<td>56.0%</td>
<td>47.5%</td>
<td>60.7%</td>
<td>35.4%</td>
<td>48.9%</td>
<td>65.6%</td>
<td>86.3%</td>
<td>46.5%</td>
<td>62.9%</td>
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<td>Oil Exports:</td>
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<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
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<tr>
<td>-Global Rank</td>
<td>45</td>
<td>11</td>
<td>23</td>
<td>1</td>
<td>58</td>
<td>70</td>
<td>86</td>
<td>64</td>
<td></td>
</tr>
<tr>
<td>-BBL/Day</td>
<td>0.24MM</td>
<td>2.20MM</td>
<td>0.73MM</td>
<td>8.55MM</td>
<td>0.15MM</td>
<td>0.08MM</td>
<td>0MM</td>
<td>0.02MM</td>
<td>0.11MM</td>
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<tr>
<td>Population</td>
<td>0.7MM</td>
<td>3.5MM</td>
<td>3.2MM</td>
<td>27.6MM</td>
<td>81.7MM</td>
<td>7.5MM</td>
<td>6.2MM</td>
<td>34.3MM</td>
<td>71.9MM</td>
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</table>

Notes: These data obtained from the 2008 *CIA World Factbook*. Per capita GDP is stated in 2008 USD and was computed on a purchasing power parity basis. Due to rounding errors, the sectoral shares for the composition of GDP do not necessarily add up to 100% for each country.
Table 2: Sample Statistics for Daily Excess Returns and Unit Root Tests

<table>
<thead>
<tr>
<th></th>
<th>Dow Jones Global</th>
<th>Bahrain</th>
<th>Kuwait</th>
<th>Oman</th>
<th>Saudi Arabia</th>
<th>Egypt</th>
<th>Israel</th>
<th>Jordan</th>
<th>Morocco</th>
<th>Turkey</th>
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<tbody>
<tr>
<td>Mean</td>
<td>3.96e-05</td>
<td>2.91e-4</td>
<td>7.66e-4</td>
<td>6.64e-4</td>
<td>5.23e-4</td>
<td>2.65e-4</td>
<td>3.41e-4</td>
<td>4.78e-4</td>
<td>3.81e-4</td>
<td>2.86e-4</td>
</tr>
<tr>
<td>Standard Deviation</td>
<td>0.009</td>
<td>0.005</td>
<td>0.008</td>
<td>0.008</td>
<td>0.015</td>
<td>0.015</td>
<td>0.014</td>
<td>0.012</td>
<td>0.009</td>
<td>0.034</td>
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<tr>
<td>Skewness</td>
<td>-0.214</td>
<td>0.227</td>
<td>-0.357</td>
<td>0.043</td>
<td>-1.107</td>
<td>-0.218</td>
<td>-0.456</td>
<td>0.048</td>
<td>0.004</td>
<td>-0.132</td>
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</thead>
<tbody>
<tr>
<td>ADF Log-Level Data</td>
<td>-1.533</td>
<td>-2.581</td>
<td>-2.345</td>
<td>-2.248</td>
<td>-0.858</td>
<td>-0.737</td>
<td>-1.386</td>
<td>-1.223</td>
<td>0.389</td>
<td>-1.882</td>
</tr>
<tr>
<td>ERS Log-Level Data</td>
<td>18.355</td>
<td>132.703</td>
<td>17.039</td>
<td>64.135</td>
<td>41.304</td>
<td>240.388</td>
<td>27.182</td>
<td>75.318</td>
<td>92.739</td>
<td>14.280</td>
</tr>
<tr>
<td>ADF Excess Returns</td>
<td>-43.626*</td>
<td>-28.966*</td>
<td>-43.092*</td>
<td>-23.338*</td>
<td>-44.533*</td>
<td>-34.013*</td>
<td>-49.242*</td>
<td>-50.805*</td>
<td>-42.033*</td>
<td>-51.085*</td>
</tr>
<tr>
<td>ERS Excess Returns</td>
<td>0.042*</td>
<td>0.075*</td>
<td>0.025*</td>
<td>0.032*</td>
<td>0.023*</td>
<td>0.021*</td>
<td>0.033*</td>
<td>0.059*</td>
<td>0.058*</td>
<td>0.020*</td>
</tr>
</tbody>
</table>

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. ‘ADF’ and ‘ERS’ stand, respectively, for the augmented Dickey-Fuller and Elliott et al. (1996) unit root test statistics. *, †, and ‡ denote rejection of the unit root null hypothesis at the 1%, 5%, and 10% significance levels, respectively. An intercept term and linear time trend were included in the regressions run for the unit root tests on the log-level stock index data, and the linear time trend was excluded from the excess returns unit root tests.
Table 3: Static International CAPM and Factor Model Results

<table>
<thead>
<tr>
<th>International CAPM</th>
<th>Bahrain</th>
<th>Kuwait</th>
<th>Oman</th>
<th>Saudi Arabia</th>
<th>Egypt</th>
<th>Israel</th>
<th>Jordan</th>
<th>Morocco</th>
<th>Turkey</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\alpha$</td>
<td>-0.003</td>
<td>-0.003</td>
<td>-0.003</td>
<td>0.005</td>
<td>0.067</td>
<td>0.594</td>
<td>0.032</td>
<td>0.014</td>
<td>1.014</td>
</tr>
<tr>
<td>$\beta$</td>
<td>0.002</td>
<td>-0.003</td>
<td>0.012</td>
<td>0.020</td>
<td>0.019</td>
<td>0.011</td>
<td>-0.006</td>
<td>0.003</td>
<td>0.019</td>
</tr>
<tr>
<td>$\delta_1$</td>
<td>0.003</td>
<td>-0.003</td>
<td>0.012</td>
<td>0.020</td>
<td>0.019</td>
<td>0.011</td>
<td>-0.006</td>
<td>0.003</td>
<td>0.019</td>
</tr>
<tr>
<td>$\delta_2$</td>
<td>0.003</td>
<td>-0.003</td>
<td>0.012</td>
<td>0.020</td>
<td>0.019</td>
<td>0.011</td>
<td>-0.006</td>
<td>0.003</td>
<td>0.019</td>
</tr>
<tr>
<td>$\delta_3$</td>
<td>0.003</td>
<td>-0.003</td>
<td>0.012</td>
<td>0.020</td>
<td>0.019</td>
<td>0.011</td>
<td>-0.006</td>
<td>0.003</td>
<td>0.019</td>
</tr>
<tr>
<td>$\gamma_1$</td>
<td>-0.003</td>
<td>-0.003</td>
<td>0.012</td>
<td>0.020</td>
<td>0.019</td>
<td>0.011</td>
<td>-0.006</td>
<td>0.003</td>
<td>0.019</td>
</tr>
<tr>
<td>$\gamma_2$</td>
<td>-0.003</td>
<td>-0.003</td>
<td>0.012</td>
<td>0.020</td>
<td>0.019</td>
<td>0.011</td>
<td>-0.006</td>
<td>0.003</td>
<td>0.019</td>
</tr>
<tr>
<td>$\gamma_3$</td>
<td>-0.003</td>
<td>-0.003</td>
<td>0.012</td>
<td>0.020</td>
<td>0.019</td>
<td>0.011</td>
<td>-0.006</td>
<td>0.003</td>
<td>0.019</td>
</tr>
<tr>
<td>$\gamma_4$</td>
<td>-0.003</td>
<td>-0.003</td>
<td>0.012</td>
<td>0.020</td>
<td>0.019</td>
<td>0.011</td>
<td>-0.006</td>
<td>0.003</td>
<td>0.019</td>
</tr>
<tr>
<td>$\gamma_5$</td>
<td>-0.003</td>
<td>-0.003</td>
<td>0.012</td>
<td>0.020</td>
<td>0.019</td>
<td>0.011</td>
<td>-0.006</td>
<td>0.003</td>
<td>0.019</td>
</tr>
<tr>
<td>$\gamma_6$</td>
<td>-0.003</td>
<td>-0.003</td>
<td>0.012</td>
<td>0.020</td>
<td>0.019</td>
<td>0.011</td>
<td>-0.006</td>
<td>0.003</td>
<td>0.019</td>
</tr>
<tr>
<td>$\gamma_7$</td>
<td>-0.003</td>
<td>-0.003</td>
<td>0.012</td>
<td>0.020</td>
<td>0.019</td>
<td>0.011</td>
<td>-0.006</td>
<td>0.003</td>
<td>0.019</td>
</tr>
<tr>
<td>$\gamma_8$</td>
<td>-0.003</td>
<td>-0.003</td>
<td>0.012</td>
<td>0.020</td>
<td>0.019</td>
<td>0.011</td>
<td>-0.006</td>
<td>0.003</td>
<td>0.019</td>
</tr>
</tbody>
</table>

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, $r_i = \alpha + \beta r^W + \varepsilon_i$ and $r_i = \alpha + \beta r^M + \sum_{j=1}^{M} \delta F_{i,j} + \sum_{k=1}^{K} \gamma_k d^k_{i} + \varepsilon_i$, equations (1) and (2), where $r_i$ is the market excess return in country $i$, $r^W$ is the world market excess return, the factors $F_{i,j}$ 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 $d^k_{i}$ variables represent the Hulich factors, and $\varepsilon_i$ is assumed to be a white noise innovation process.
Table 4: Intertemporal CAPM GARCH-M Results

<table>
<thead>
<tr>
<th>EGARCH-M</th>
<th>Bahrain</th>
<th>Kuwait</th>
<th>Oman</th>
<th>Saudi Arabia</th>
<th>Egypt</th>
<th>Israel</th>
<th>Jordan</th>
<th>Morocco</th>
<th>Turkey</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\hat{\mu}$</td>
<td>$-8.89e-4^*$</td>
<td>$1.20e-3^*$</td>
<td>$-2.00e-4$</td>
<td>$5.71e-4^*$</td>
<td>$-1.53e-3^*$</td>
<td>$6.17e-4^*$</td>
<td>$-5.43e-4^*$</td>
<td>$-1.97e-4$</td>
<td>$2.28e-3^*$</td>
</tr>
<tr>
<td>$\lambda$</td>
<td>$40.31^*$</td>
<td>$-2.489$</td>
<td>$11.895$</td>
<td>$3.056^*$</td>
<td>$6.881^*$</td>
<td>$-0.221$</td>
<td>$6.808^*$</td>
<td>$5.920$</td>
<td>$-1.321$</td>
</tr>
<tr>
<td>$\omega$</td>
<td>$-1.005^*$</td>
<td>$-0.730^*$</td>
<td>$-0.589^*$</td>
<td>$-0.039^*$</td>
<td>$-0.091^*$</td>
<td>$-0.500^*$</td>
<td>$-0.011^*$</td>
<td>$-0.603^*$</td>
<td>$-0.131^*$</td>
</tr>
<tr>
<td>$\alpha$</td>
<td>$0.188^*$</td>
<td>$0.201^*$</td>
<td>$0.142^*$</td>
<td>$0.174^*$</td>
<td>$0.122^*$</td>
<td>$0.196^*$</td>
<td>$0.080^*$</td>
<td>$0.310^*$</td>
<td>$0.189^*$</td>
</tr>
<tr>
<td>$\beta$</td>
<td>$0.900^*$</td>
<td>$0.923^*$</td>
<td>$0.936^*$</td>
<td>$0.992^*$</td>
<td>$0.987^*$</td>
<td>$0.941^*$</td>
<td>$0.992^*$</td>
<td>$0.936^*$</td>
<td>$0.980^*$</td>
</tr>
<tr>
<td>$\hat{\sigma}$</td>
<td>$0.283^*$</td>
<td>$-0.247^*$</td>
<td>$0.117^*$</td>
<td>$0.134^*$</td>
<td>$0.191^*$</td>
<td>$-0.455^*$</td>
<td>$0.543^*$</td>
<td>$0.055$</td>
<td>$-0.187^*$</td>
</tr>
</tbody>
</table>

<table>
<thead>
<tr>
<th>GJR GARCH-M</th>
<th>Bahrain</th>
<th>Kuwait</th>
<th>Oman</th>
<th>Saudi Arabia</th>
<th>Egypt</th>
<th>Israel</th>
<th>Jordan</th>
<th>Morocco</th>
<th>Turkey</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\hat{\mu}$</td>
<td>$-6.85e-3^*$</td>
<td>$5.89e-4^*$</td>
<td>$1.32e-5$</td>
<td>$5.86e-4^*$</td>
<td>$-8.23e-4^*$</td>
<td>$7.90e-4^*$</td>
<td>$-3.62e-4$</td>
<td>$-1.44e-4$</td>
<td>$1.59e-3^*$</td>
</tr>
<tr>
<td>$\lambda$</td>
<td>$30.31^*$</td>
<td>$3.907$</td>
<td>$8.721$</td>
<td>$0.901$</td>
<td>$4.787^*$</td>
<td>$-1.555$</td>
<td>$5.292^*$</td>
<td>$5.339$</td>
<td>$-0.936$</td>
</tr>
<tr>
<td>$\omega$</td>
<td>$3.79e-6^*$</td>
<td>$4.10e-6$</td>
<td>$3.91e-6^*$</td>
<td>$1.20e-6^*$</td>
<td>$1.80e-6^*$</td>
<td>$9.47e-6^*$</td>
<td>$5.27e-7^*$</td>
<td>$6.09e-6^*$</td>
<td>$2.13e-5^*$</td>
</tr>
<tr>
<td>$\alpha$</td>
<td>$0.142^*$</td>
<td>$0.086^*$</td>
<td>$0.065^*$</td>
<td>$0.098^*$</td>
<td>$0.061^*$</td>
<td>$0.047^*$</td>
<td>$0.051^*$</td>
<td>$0.202^*$</td>
<td>$0.073^*$</td>
</tr>
<tr>
<td>$\beta$</td>
<td>$0.786^*$</td>
<td>$0.843^*$</td>
<td>$0.888^*$</td>
<td>$0.926^*$</td>
<td>$0.944^*$</td>
<td>$0.850^*$</td>
<td>$0.963^*$</td>
<td>$0.742^*$</td>
<td>$0.889^*$</td>
</tr>
<tr>
<td>$\gamma$</td>
<td>$-0.087^*$</td>
<td>$0.031^*$</td>
<td>$-0.020^*$</td>
<td>$-0.041^*$</td>
<td>$-0.022^*$</td>
<td>$0.105^*$</td>
<td>$-0.036^*$</td>
<td>$-0.025$</td>
<td>$0.047^*$</td>
</tr>
</tbody>
</table>

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 $r_i^t = \mu + \lambda h_i^{1/2} + \epsilon_t$, where $r_i^t$ is the market excess return in country $i$, $\epsilon_t = \sqrt{h_t} \epsilon_t$, $\epsilon_t \sim N(0, 1)$, and $h_t^1$ 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 $\ln(h_t^1) = \omega + \alpha (z_{t-1}^1 + \beta \ln(h_{t-1}^0))$, where $g(z_t) = \theta z_t + \delta |z_t| - E|z_t|$, $z_t = \epsilon_t / \sqrt{h_t}$, and $\delta = 1$. In the GJR GARCH-M model, the conditional variance is given by $h_t^1 = \omega + \alpha z_{t-1}^2 + \gamma I_{s_t < 0} \epsilon_t^2 + \beta h_{t-1}^0$, where $I_{s_t < 0}$ is an indicator function which takes on the value of 1 when $s_t < 0$ and 0 otherwise.
Table 5: Markov-Switching Intertemporal CAPM Results

<table>
<thead>
<tr>
<th>Sample Statistics of $\hat{\phi}_i^t$</th>
<th>Bahrain</th>
<th>Kuwait</th>
<th>Oman</th>
<th>Saudi Arabia</th>
<th>Egypt</th>
<th>Israel</th>
<th>Jordan</th>
<th>Morocco</th>
<th>Turkey</th>
</tr>
</thead>
<tbody>
<tr>
<td>Mean</td>
<td>0.660</td>
<td>0.365</td>
<td>0.297</td>
<td>0.326</td>
<td>0.220</td>
<td>0.598</td>
<td>0.342</td>
<td>0.394</td>
<td>0.592</td>
</tr>
<tr>
<td>Median</td>
<td>1.0</td>
<td>0.200</td>
<td>0.0</td>
<td>0.0</td>
<td>0.0</td>
<td>0.783</td>
<td>0.0</td>
<td>0.253</td>
<td>0.754</td>
</tr>
<tr>
<td>Standard Deviation</td>
<td>0.434</td>
<td>0.337</td>
<td>0.439</td>
<td>0.458</td>
<td>0.404</td>
<td>0.381</td>
<td>0.443</td>
<td>0.327</td>
<td>0.377</td>
</tr>
</tbody>
</table>

<table>
<thead>
<tr>
<th>Transition Probabilities</th>
<th></th>
<th></th>
<th></th>
<th></th>
<th></th>
<th></th>
<th></th>
<th></th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td>$\hat{p}_{1,1}$</td>
<td>0.65</td>
<td>0.85</td>
<td>0.21</td>
<td>0.31</td>
<td>0.11</td>
<td>0.97</td>
<td>0.29</td>
<td>0.83</td>
<td>0.95</td>
</tr>
<tr>
<td></td>
<td>(0.16)</td>
<td>(0.03)</td>
<td>(0.03)</td>
<td>(0.02)</td>
<td>(0.02)</td>
<td>(0.01)</td>
<td>(0.01)</td>
<td>(0.03)</td>
<td>(0.01)</td>
</tr>
<tr>
<td>$\hat{p}_{2,2}$</td>
<td>0.22</td>
<td>0.95</td>
<td>0.70</td>
<td>0.68</td>
<td>0.77</td>
<td>0.94</td>
<td>0.68</td>
<td>0.93</td>
<td>0.98</td>
</tr>
<tr>
<td></td>
<td>(0.32)</td>
<td>(0.01)</td>
<td>(0.02)</td>
<td>(0.02)</td>
<td>(0.01)</td>
<td>(0.01)</td>
<td>(0.02)</td>
<td>(0.01)</td>
<td>(0.09)</td>
</tr>
</tbody>
</table>

<table>
<thead>
<tr>
<th>Risk Premia</th>
<th></th>
<th></th>
<th></th>
<th></th>
<th></th>
<th></th>
<th></th>
<th></th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td>$\hat{\lambda}$</td>
<td>2.458</td>
<td>1.383</td>
<td>2.69e−6</td>
<td>0.346</td>
<td>0.058</td>
<td>7.731</td>
<td>0.186</td>
<td>5.011</td>
<td>6.337</td>
</tr>
<tr>
<td></td>
<td>[0.607]</td>
<td>[0.317]</td>
<td>[0.206]</td>
<td>[0.055]</td>
<td>[0.206]</td>
<td>[4.59e−6]</td>
<td>[0.044]</td>
<td>[0.020]</td>
<td>[2.0e−5]</td>
</tr>
<tr>
<td>$\hat{\lambda}^i$</td>
<td>2.379</td>
<td>8.765</td>
<td>7.766</td>
<td>0.789</td>
<td>0.951</td>
<td>0.548</td>
<td>1.848</td>
<td>3.886</td>
<td>0.033</td>
</tr>
<tr>
<td></td>
<td>[0.138]</td>
<td>[1.0e−5]</td>
<td>[0.001]</td>
<td>[0.088]</td>
<td>[0.237]</td>
<td>[0.371]</td>
<td>[0.527]</td>
<td>[0.023]</td>
<td>[0.527]</td>
</tr>
</tbody>
</table>

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

\[ r_i^t = \phi_{i-1}^t \lambda \text{Cov} \left( r_i^t, r_i^W \right) + (1 - \phi_{i-1}^t) \lambda^i \text{Var} \left( r_i^t \right) + \epsilon_i^t, \]

where $r_i^t$ is the market excess return in country $i$, $r_i^W$ is the world market excess return, $\phi_{i-1}^t = \text{Prob}(S_i^t = 1|\mathcal{F}_{t-1})$, $S_i^t$ is a state variable which can take on two values, with $S_i^t = 1$ denoting that market $i$ is integrated with international equity markets in observation $t$ and $S_i^t = 2$ denoting it is segmented, $\mathcal{F}_{t-1}$ is the observation $t-1$ information set, $p_{1,1}^i = \text{Prob}(S_i^t = 1|\mathcal{F}_{t-1})$, $p_{2,2}^i = \text{Prob}(S_i^t = 2|\mathcal{F}_{t-1})$, and both $\lambda$ and $\lambda^i$, 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.
<table>
<thead>
<tr>
<th>Country</th>
<th>50-Day Frame</th>
<th>Event</th>
</tr>
</thead>
<tbody>
<tr>
<td>Bahrain</td>
<td>7/16/01 - 9/21/01</td>
<td>September 11 attacks in US weigh heavily on Bahraini markets</td>
</tr>
<tr>
<td>Bahrain</td>
<td>11/3/03 - 1/4/04</td>
<td>Banks in the Central Governorate started to pay full taxes</td>
</tr>
<tr>
<td>Bahrain</td>
<td>5/31/04 - 8/9/04</td>
<td>Bahrain and US conclude free-trade agreement</td>
</tr>
<tr>
<td>Egypt</td>
<td>4/8/98 - 6/16/98</td>
<td>Cairo and Alexandria stock markets to establish electronic trading system</td>
</tr>
<tr>
<td>Egypt</td>
<td>6/17/98 - 8/25/98</td>
<td>Government creates three large investment funds (LE900 million)</td>
</tr>
<tr>
<td>Egypt</td>
<td>8/26/98 - 11/3/98</td>
<td>Chair of Cairo and Alexandra stock exchanges will resign in December ’98</td>
</tr>
<tr>
<td>Egypt</td>
<td>11/4/98 - 1/12/99</td>
<td>Privatization Committee will offer stakes in additional six companies</td>
</tr>
<tr>
<td>Israel</td>
<td>6/17/98 - 8/25/98</td>
<td>Central bank cut interest rate for the eighth time in the year</td>
</tr>
<tr>
<td>Israel</td>
<td>8/26/98 - 11/3/98</td>
<td>Debate over Wye River Accord cast shadow over the markets</td>
</tr>
<tr>
<td>Israel</td>
<td>8/11/99 - 10/19/99</td>
<td>Ehud Barak struggled to piece together a coalition government</td>
</tr>
<tr>
<td>Israel</td>
<td>10/20/99 - 12/28/99</td>
<td>Israel reinstated peace talks with Palestine and Syria</td>
</tr>
<tr>
<td>Israel</td>
<td>10/4/00 - 12/12/00</td>
<td>Prime Minister Ehud Barak resigned</td>
</tr>
<tr>
<td>Jordan</td>
<td>1/13/99 - 3/24/99</td>
<td>King Hussein died</td>
</tr>
<tr>
<td>Jordan</td>
<td>11/28/01 - 2/5/02</td>
<td>Riots erupt in southern town of Maan</td>
</tr>
<tr>
<td>Morocco</td>
<td>3/24/99 - 6/1/99</td>
<td>Parliament passed amended privatization law</td>
</tr>
<tr>
<td>Morocco</td>
<td>3/8/00 - 5/16/00</td>
<td>Series of droughts hit the economy</td>
</tr>
<tr>
<td>Morocco</td>
<td>7/26/00 - 10/3/00</td>
<td>Oil discovered in country’s eastern region</td>
</tr>
<tr>
<td>Morocco</td>
<td>10/4/00 - 12/12/00</td>
<td>Tender of 35% stake in Maroc Telecom was revived</td>
</tr>
<tr>
<td>Morocco</td>
<td>11/28/01 - 2/5/02</td>
<td>Casablanca Exchange General Index sank to 5-year lows</td>
</tr>
<tr>
<td>Morocco</td>
<td>5/26/04 - 8/3/04</td>
<td>Spanish government to revise cooperation arrangement with Morocco</td>
</tr>
<tr>
<td>Oman</td>
<td>2/14/01 - 4/24/01</td>
<td>Oman became liquified natural gas exporter</td>
</tr>
<tr>
<td>Oman</td>
<td>11/21/01 - 1/29/02</td>
<td>Cost of entry visas reduced as part of efforts to promote tourism</td>
</tr>
<tr>
<td>Saudi Arabia</td>
<td>7/16/01 - 9/21/01</td>
<td>Negotiations with European Union for free-trade agreement</td>
</tr>
<tr>
<td>Saudi Arabia</td>
<td>2/11/02 - 4/19/02</td>
<td>Slumping oil income leads to large increase in government domestic borrowing</td>
</tr>
<tr>
<td>Saudi Arabia</td>
<td>8/25/03 - 10/31/03</td>
<td>British arms company BAE Systems accused of bribing Saudi officials</td>
</tr>
<tr>
<td>Saudi Arabia</td>
<td>3/22/04 - 5/28/04</td>
<td>Cut in corporate tax rate for foreign companies operating in Saudi Arabia</td>
</tr>
<tr>
<td>Turkey</td>
<td>3/24/99 - 6/1/99</td>
<td>Center-left party wins largest number of seats in parliamentary election</td>
</tr>
<tr>
<td>Turkey</td>
<td>10/4/00 - 12/12/00</td>
<td>Banking crisis due to anxiety over bank liquidity problems</td>
</tr>
<tr>
<td>Turkey</td>
<td>7/11/01 - 9/18/01</td>
<td>S&amp;P lowered Turkey’s debt rating to ‘B’</td>
</tr>
<tr>
<td>Turkey</td>
<td>9/4/02 - 11/12/02</td>
<td>Islamist-based Justice and Development Party wins landslide election victory</td>
</tr>
<tr>
<td>Turkey</td>
<td>11/13/02 - 1/31/03</td>
<td>Constitutional change allows of head of ruling party to become prime minister</td>
</tr>
<tr>
<td>Turkey</td>
<td>6/11/03 - 8/19/03</td>
<td>Parliament passes legislation reducing political role of military</td>
</tr>
</tbody>
</table>

Notes: For each country, the table lists those 50-day time frames in which the pure white noise null hypothesis is rejected at the 5% significance level via the Hinich (1996) test and for which the online Bekaert and Harvey emerging market country risk analysis database (http://www.duke.edu/~charvey/Country_risk/couindex.htm), which was last updated in September of 2004, reports an important financial, economic, or political event.
References


