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STOCK PRICES AND INFLATION:
EVIDENCE FROM JORDAN, SAUDI ARABIA,
KUWAIT, AND MOROCCO

Adel Al-Sharkas and Marwan Al-Zoubi

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Abstract

This paper attempts to investigate monthly stock price indexes and good price indexes for selected Mashreq and Maghreb countries: Jordan, Saudi Arabia, Morocco, and Kuwait for 2000-2009 using cointegration methods. Our findings support the long-run relationship between stock prices and goods prices. The long-run Fisher elasticities of stock prices with respect to goods prices are in the range of 1.01 to 1.36 across the four countries under study. With the exception of Kuwait, the Fisher effect coefficient estimates are significantly greater than one. In the case of Jordan, Saudi Arabia, and Morocco the empirical results support the Fisher hypothesis, with estimated coefficients near unity.

ملخص

تحاول هذه الورقة بحث المؤشرات الشهرية لأسعار الأسهم ومؤشرات الأسعار الجيدة لدول المشرق والمغرب المختارة : الأردن و المملكة العربية السعودية والمغرب والكويت للاعوام من 2000-2009 و ذلك باستخدام طرق التكامل المشترك. و ندعم نتائجنا العلاقة طويلة المدى بين أسعار الأسهم وأسعار السلع. ومع ذلك ، تتأثر الأسهم من التضخم على المدى القصير. اما على المدى الطويل ، تتراوح مرونة فيشر لأسعار الأسهم بالنسبة لأسعار السلع بين 1.01-1.36 في البلدان الأربعة قيد الدراسة . و باستثناء الكويت ، ترتفع تقديرات معامل فيشر أكثر بكثير من واحد. اما بالنسبة لكلا من الأردن ، و المملكة العربية السعودية والمغرب، ندعم النتائج التجريبية فرضية فيشر ، مع معاملات مقدرة تقترب من الوحدة.

1. Introduction

The relationships between stock returns and inflation rates or stock prices and goods prices have been the subject of numerous research papers during the last five decades. Mixed results were obtained ranging from no relationship, to negative and positive relationships. However, the short-term negative relationship and the long-term positive relationship seem to be well established in the literature.

Theoretically, stocks are assumed to be inflation neutral for unexpected inflation which should have a negative effect on stock prices. The standard discounted cash flow model calculates stock prices as the present value of future expected cash flows. For stocks to be inflation neutral and represent a good long-term hedge against inflation, firms should pass on any increase in inflation rates on future cash flows. Investors on the other hand should discount the adjusted cash flows by rate of return or discount the real cash flows by the same real discount rate.

This argument is now known as the Fisher hypothesis (1930) or Fisher effect. Fisher argues that stocks are claims against real assets therefore they are neutral and uncorrelated with inflation rates. Stock returns are equal to the real rate of return plus an inflation premium. Since this is an equilibrium relationship, it is supposed to hold in the long-run. Hoguet (2009) reports that when inflation rates were historically high or accelerating in the US, price/earnings ratios were declining, a phenomenon which puzzled researchers and practitioners as well.

The literature offers three theories as possible explanations for this relationship. Fisher hypothesis is more applicable in the long-run as an equilibrium relationship. The behavioral hypothesis proposed by Modigliani and Cohn (1979) has been used to explain the short-run relationship, or what is known today as the *inflation illusion phenomena*. This hypothesis says that investors mistakenly price future real cash flows by discounting by the nominal rate of return. The proxy hypothesis proposed by Fama (1981) argues that because stock returns are positively related to future real economic growth, as inflation increases, real economic growth declines and become more volatile which pushes investors to require higher risk premiums to cover the additional risk. Stock prices therefore start declining accordingly. This hypothesis is also supported by Sharpe (1999).

The debate is still ongoing although standard investment and finance theory implies that stocks are good hedging instruments at least in the long run. It is however necessary to differentiate between the effect of anticipated inflation and unanticipated inflation. While stocks should be inflation neutral, it is reasonable to assume that their prices react negatively to inflation shocks. New investors require risk premium for any increase in unanticipated inflation, and this implies a decline in stock prices. However, one should admit that investors sometimes do not follow standard investment theory because of behavioral reasons and lack of sophisticated knowledge about finance fundamentals. In markets with more and more institutional investors, the behavioral hypothesis becomes less realistic on the grounds that institutional investors are expected to be sophisticated investors.

Stocks should be good hedging assets against inflation but they are also sensitive to future changes in inflation rates. Investors may have incentives to adjust their valuation of stock piecemeal. That is, every time they are faced with a new inflation rate, they make the adjustment instantaneously as if they are dealing with a short-term security with short-term investment horizon.

In this paper, we contribute further evidence on the long-run Fisher effect for stocks by using stock prices and goods prices. Consistent with Engle and Granger (1987), this approach allows us to fully use long-run information contained in the levels of the variables, as

opposed to focusing on partial long-run information contained in selected holding periods. Additionally, by using levels rather than first differences, we avoid using data over long historical periods, in which the accuracy and relevance of the data series can be compromised.

We examine monthly time series of stock price indexes and goods price indexes (consumer price indexes) for four Arab countries (Jordan, Saudi Arabia, Kuwait, and Morocco) from January 2000 to December 2009. While the use of cointegration analysis abounds in the empirical literature that previously examined the hypothesis, such as Mishkin (1992), Anari and Kolari (2001) and Shrestha, Chen and Lee (2002), we have found no studies that employed cointegration analysis in testing the hypothesis for these countries. In addition, we are applying a relatively new and powerful methodology, the generalized forecast error variance decomposition components and the generalized impulse response functions computed from estimated unrestricted vector autoregressive (UVAR) models.

Unlike the traditional forecast error variance decomposition and impulse response functions, these approaches do not require orthogonalization of shocks and is invariant to the ordering of the variables in the UVAR model, while the widely used Choleski factorization is known to be sensitive to the ordering of the variables. Finally, the UVAR is employed to avoid using arbitrary choice of the restrictions needed to settle the identification.

Our study provides evidence on the long-term Fisher effect on stocks in four Arab countries markets – Jordan, Saudi Arabia, Kuwait, and Morocco. While we perform the regression analysis with stock return and inflation rates, we use the levels of stock prices and corresponding changes in inflation in our cointegration tests. Evidence has shown that it is impossible to measure fully both contemporaneous and inter-temporal correlations between stock returns in real terms and inflation when variables are evaluated by their first differences (Gallagher, 1986).

As Hendry (1986) and Juselius (1991) observe, when a time series is differenced, long-run information contained in the levels of variables is lost. Consistent with the findings of Granger (1986), Engle and Granger (1987) and Anari and Kolari (2001), the use of levels allows us to fully evaluate long-term information that may be contained in continuous time series variables, as opposed to focusing on the segmented information variables whose values in their first differences are susceptible to different lengths of observation.

The findings of this study support the long-term Fisher effect between stock prices and corresponding changes in the price of goods as measured by the consumer price index; namely, stock prices appear to reflect a time-varying memory associated with inflation shocks that make stock portfolios a reasonably good hedge against inflation over the long run in four Arabic equity markets. This means that the stock markets are relatively efficient in impounding forthcoming inflation from the concurrent changes in stock prices, and investors can adjust their portfolios accordingly.

This paper is organized as follows. Section 2 presents the literature review. Section 3 describes the data. Section 4 introduces the VAR model. Section 5 reports evidence from the VAR model. Section 6 concludes.

2. Literature Review

Traditionally, it was believed that stocks provide a good hedge against inflation given that stocks represent claims on real assets so that stock returns are positively correlated with actual inflation. In recent years, empirical research showed that inflation affects stock returns negatively (Sharpe 1999). In the United States, high expected inflation and accelerating inflation have been associated with decreasing price/earnings ratios.

Stocks seem to be better hedges against inflation in the medium and long- term compared to the short term.

However, Fisher (1930) contends that return on assets move one – for – one with anticipated inflation. That is, stocks should be inflation neutral but stock prices react negatively to high unexpected inflation. Fisher argues that real stock returns are related to real factors and that stocks should maintain their purchasing power in the long -term. Fama and Schwert (1977) argue that anticipated inflation negatively affected stock returns during the period 1953-1971 and concluded that stocks are not good hedges against inflation. Cohn and Modigliani (1979) argue that U.S investors undervalued stocks because they discounted (mistakenly) future real cash flows by using nominal rates of return. They use quarterly data over 1953-1977 on price/earnings ratios and inflation rates in their analysis.

Lansing (2004) confirms the behavioral anomaly of investors discounting future real cash flows by nominal rates of return.

It is well established in the literature that rising inflation and future real economic growth are negatively correlated. During the 1970s, the U.S. experienced a decline in economic activity when inflation was rising (Hoguet 2009). This view is confirmed by Fama (1981). Fama argues that stock returns are positively related to expected real economic growth. Future real economic growth is, on the other hand, associated with low inflation rates. When future economic growth (in real terms) is expected to decline due to high inflation, investors required higher risk premiums on their stocks.

The negative short-term relationship between stock returns and anticipated and unanticipated inflation is reported by Geske and Roll (1983) and Jaffe and Mandelker (1976) and Wei (2009). Wei found that stock returns' negative reaction to unanticipated inflation is higher during economic contractions than expansions. On the other hand, The long – run positive relationship (Fisher effect) is reported by many authors. Jaffe and Mandelker (1976) reported positive relationship over a long period (1875 -1970). Boudoukh and Richardson (1993) confirmed the same result applying one-year and five-year holding period returns during 1802-1990 in both the United Kingdom and the United States.

Anari and Kolari (2001) used stock prices and goods prices instead of the first difference in order to overcome the problem that the first difference eliminates the long – run information. They use monthly stock price indices and goods price indices for Canada, France, Germany, the United kingdom and the United states during 1953-1998. They employ the cointegration technique developed by Johansen (1988) for those goods prices and stock prices are cointegrated and non-stationary and confirm the long memory Fisher effect which says that stocks are good inflation hedges over a long holding period. However, they also report the negative initial effect in all six countries.

Luintel and Paudyal (2006) support previous results and report the existence of the long – run hedging relationship in the UK stock market.

Although the short-run negative effect (the inflation illusion as named by Modigliani and Chon (1979)) and the long – run hedging Fisher effect are well established in empirical research, Ely and Robinson (1997) found no long – run relationship. They apply the Johansen's (1988) method on sixteen countries during 1957-1992 .

Aga and Kocaman (2006) tested the impact of price/earnings ratios, industrial price indices (IPI) and the consumer price indices (CPI) on returns of stocks traded in Istanbul Stock Exchange. They claim that macroeconomic variables such as inflation rates should have two possible effects.

The direct effect hypothesis implies that stock markets normally react negatively to bad news and positively to good news. The policy signaling hypothesis implies that it is possible for the market to react positively to adverse movements in macroeconomic variables due to anticipated government remedial actions.

Their findings indicate that only the price/earnings ratio appears to be significant in explaining the movements in stock returns, while industrial price indices and consumer price indices are not. Exponential GARCH model was applied to test the impact of CPI and IPI on stock return and volatility. They also found that these variables are not statistically significant in explaining stock returns and volatility.

3. An Overview of the Data

This study covers four Arab equity markets: Jordan, Saudi Arabia, Kuwait, and Morocco. Monthly consumer price index (CPI) and monthly stock prices are used. Stock price indices are taken from their respective homepages. Consumer price indices for all countries are from Monthly Financial Statistics (International Monetary Fund). The sample period begins in January 2000 and ends in December 2009. All variables are transformed into natural logarithms. Using data from different countries enables comparative analysis to check the robustness of the results. In this regard, there is some concern about the power of tests to detect cointegration for different sample sizes. Based on the literature on this subject (e.g., see Hakkio and Rush (1991) and Junttila (2001)), we infer that our sample sizes are sufficiently large to provide reliable cointegration tests. Hakkio and Rush (1991) say that "cointegration is a long-run concept and hence requires long span of data to give tests for cointegration much power rather than merely large numbers of observations." Also, testing several models for multiple countries allows for the consistency of the estimated test statistics to be observed.

4. Methodology

4.1 Unit root tests

The first step in our statistical analysis is to analyze the stationary properties of the macro time series considered in this study by applying the unit root. Applying the unit root test will do this. Unit root tests are important in examining the stationary of a time series, which is a matter of concern in three important areas. First, a crucial question in the ARIMA modeling of a single time series is the number of times the series needs to be first differenced before an ARMA model is fit. Each unit root requires a first differencing operation. Second, stationary of regressors is assumed in the derivation of standard inference procedures for regression models. Nonstationary regressors invalidate many standard results and require special treatment. Third, in cointegration analysis, an important question is whether the disturbance term in the cointegrating vector has a unit root.

The Augmented Dickey-Fuller Test (ADF) is applied in this paper. The ADF test consists of running a regression of the first difference of the series against the series lagged once, lagged difference terms, and optionally, a constant and a time trend. With two lagged difference terms.

There are three choices in running the ADF test regression: to include a constant term in the regression, to include a linear time trend, or to determine how many lagged differences are to be included in the regression.

In each case the test for a unit root is a test on the coefficient of the regression. If the coefficient is significantly different from zero then the hypothesis that y contains a unit root is rejected and the hypothesis is accepted that y is stationary rather than integrated.

The output of the ADF test consists of the t-statistic on the coefficient of the lagged test variable and critical values for the test of a zero coefficient. A large negative t-statistic rejects the hypothesis of a unit root and suggests that the series is stationary. Under the null hypothesis of a unit root, the reported t-statistic does not have the standard t-distribution. We must refer to the critical values presented in the test output. The reported critical values are chosen on the basis of the number of observations and the estimation option.

After running the ADF test, If the Dickey-Fuller t-statistic is smaller (in absolute value) than the reported critical values, we cannot reject the hypothesis of nonstationarity and the existence of a unit root. We would conclude that our series might not be stationary. We may then wish to test whether the series is I(1) (integrated of order one) or integrated of a higher order. A series is I(1) if its first difference does not contain a unit root.

4.2 The VAR model

This study adopts an unrestricted vector autoregression (UVAR) framework to analyze the dynamic relationship between the variables. The UVAR does not impose arbitrary restrictions of the effects of the endogenous variables. It was common in earlier VAR-type analyses to rely on a Choleski factorization. Unfortunately, the Choleski factorization is known to be sensitive to the ordering of variables when the residual covariance matrix is non-diagonal. This paper employs generalized forecast error variance decomposition developed in Koop, Pesaran and Lee (1996) and Pesaran and Shin (1998) to deal with this problem. Unlike the orthogonalized forecast error variance decomposition, the generalized approach is invariant to the ordering of the variables in the UVAR model. The generalized forecast error variance decomposition from the UVAR model is computed in order to investigate interrelationships within the system. The empirical work undertaken in this study is based on estimating the UVAR on eight definitions of money. The UVAR approach, introduced by Sims (1980), suggests a standard tool to analyze time series relationships among macroeconomic variables. A VAR is a system in which every equation has the same right hand variables, and those variables include lagged values of all of the endogenous variables. VARs are well suited to forecasting variables where each variable helps forecast other variables.

The mathematical form of a UVAR is

$$y_t = A_1 y_{t-1} + \dots + A_N y_{t-N} + Bx_t + \varepsilon_t \quad (1)$$

Here y_t is a vector of endogenous variables; x_t is a vector of constant, N is the vector autoregressive order, A_i are matrices of lag coefficients of y_t up to some lag length N , and ε_t is a vector of innovations y_{t-1} x_t . The components of ε_t vector are each white noise process with zero mean, constant variance, and are individually serially uncorrelated. However, the components of ε_t vector could be contemporaneously correlated. UVARs have proven successful for forecasting systems of interrelated time series variables.

Vector autoregression is also frequently used, although with considerable controversy, for analyzing the dynamic impact of different types of random disturbances on systems of variables. However, the estimated coefficients of UVARs themselves are difficult to interpret. We will look at the generalized forecast error variance decomposition and the generalized impulse response functions of the system to draw conclusions about a UVAR.

4.2.1 The Generalized Forecast Error Variance Decomposition

Innovation accounting analysis refers to two tools used to trace the impact of shocks (innovations) in the VAR system. These tools were introduced by Sims (1980) to measure the dynamic interaction among the variables. The first, the forecast error variance decomposition (FEVD), analyzes the errors the model would tend to make if it is used to

forecast its variables. The FEVD shows how much of the average squared forecast error, which the model tends to make, is caused by innovations associated with each of the variables in the model. The FEVD of a variable, thus, can suggest that forces associated with one variable are major influences on the evolution of another variable.

The GFEVD shows how much of the average squared forecast error, which the model tends to make, is caused by innovations associated with each of the variables in the model. The GFEVD of a variable thus can suggest that forces associated with one variable are major influences on the evolution of another variable. In other words, the GFEVD of a VAR provides information about the relative importance of the random innovations. It was common in earlier VAR-type analyses to rely on a Choleski factorization. Unfortunately, the innovation accounting results, based on the Choleski factorization, are sensitive to the ordering of variables in the VAR model. In this paper, we apply generalized forecast error variance decomposition developed by Koop, Pesaran and Lee (1996) and Pesaran and Shin (1998) to deal with this problem. Unlike the orthogonalized method, the generalized approach is invariant to the ordering of the variables and does not impose the constraint that the underlying shocks to the VAR are orthogonalized before decompositions are computed. The generalized approach explicitly takes into account the contemporaneous correlation of the variables in the VAR model. The approach provides meaningful results at all the horizons including initial impact.

We calculate a separate variance decomposition for each endogenous variable. The first column is the forecast error of the variable for different forecast horizons. The source of this forecast error is variation in the current and future values of the innovations. The remaining columns give the percentage of the variance due to specific innovations. One period ahead, all of the variation in a variable comes from its own innovation, so the first number is always 100 percent.

4.2.2 Generalized Impulse Response Function

The other tool, the impulse response function, shows how one variable responds over time to a single innovation in itself or in another variable. Specifically, it traces the effect on current and future values of the endogenous variable of a one standard deviation shock to one of the innovations. Innovations or surprise movements are jointly summarized by the error terms of the UVAR model.

4.2.3 VAR Specification Issues

The following issues are related to specifying VAR models. Alternative specifications differ with respect to ordering of variables, method of "trend" removal, lag length on the VAR equations, and level of temporal aggregation. These issues must necessarily be addressed beyond the choice of variables to be included.

4.2.4 Lag length

The empirical evidence from a VAR model is very sensitive to the choice of lag length in the equations of the model. Alternative choices will give different innovations series and, thus, will likely make a difference in the variance decomposition results.

The appropriate lag length could be tested using the likelihood ratio test, the Akaike Information Criterion, or the Schwarz Criterion. In this study, the lag length will be specified based on these criteria and the results obtained in each case will be compared. Changing the lag length will also test the robustness of the empirical results.

5. Empirical Results

We first test whether the 8 time series are nonstationary. To determine the stationary properties of the series, we use the ADF unit root test and Phillips-Perron (Phillips and

Perron, 1990, PP) tests. Table 1 shows that the stock prices (SP) and consumer goods prices (CP) are generally nonstationary in the level. Therefore, the cointegration test will be applied to examine the long-run relationship between stock prices and goods prices. Table 2 represents the result for the first difference of the variables. The results show that the first differences of the series are stationary. In addition, these findings indicate that all variables employed in regressions are stationary and therefore would not cause spurious regression outcomes.

Our next task is to check whether the series are cointegrated. Specifically, having established the presence of a unit root in the first-difference of each variable, we need to test whether the series in each country has different unit roots (non-cointegrated), or shares the same unit root (cointegrated). Cointegrated variables, if disturbed, will not drift apart from each other and thus possess a long-run equilibrium relationship. The existence of cointegration would imply that the two series would not drift too far apart. A non-stationary variable, by definition, tends to wander extensively over time, but a pair of non-stationary variables may have the property that a particular linear combination would keep them together, that is, they do not drift too far apart. Under this scenario, the two variables are said to be cointegrated, or possess a long-run stable relationship.

We test the cointegration hypothesis with the methodology suggested by Johansen (1988) and Johansen and Juselius (1990). Because Johansen tests are performed within a VAR framework, and the results from VARs are sensitive to the lag length (Hafer and Sheehan (1991)), attention should be paid to lag length. Because the data are monthly, and based on lag-selection tests using the Sims (1980) criterion, we introduced twelve monthly lags in the Johansen system. Applying the MLE approach, we show in table 3 results from Johansen's trace test to determine whether a long-term relation exists between each pair of stock prices and goods prices (CPI).

A brief description of the test is in order. Let

$$\Delta x_t = \sum_{i=1}^{p-1} \Gamma_i \Delta x_{t-i} + \pi x_{t-1} + \varepsilon_t \quad (2)$$

Where x_t and ε_t are (n by 1) vectors and π is an (n by n) matrix of parameters. The Johansen (1988) methodology requires estimating the system of equations in (2) and examining the rank of matrix π . If $\text{rank}(\pi) = 0$, then there is no stationary linear combination of the $\{x_{it}\}$ process, the variables are not co-integrated. Since the rank of a matrix is the number of non-zero eigenvalues (λ), the number of $\lambda > 0$ represents the number of co-integrating vectors among the variables. The test for the non-zero eigenvalues is normally conducted using the following two test statistics:

$$\lambda_{trace}(r) = -T \sum_{i=r+1}^n \ln(1 - \hat{\lambda}_i) \quad (3)$$

$$\lambda_{max}(r, r+1) = -T \ln(1 - \hat{\lambda}_{r+1}) \quad (4)$$

Where $\hat{\lambda}_i$ is the estimated eigenvalues, and T is the number of valid observations. Note that

λ_{trace} statistic is simply the sum of λ_{max} statistic. In equation (3), λ_{trace} tests the null hypothesis that the number of distinct co-integrating vectors is less than or equal to r against a general alternative. λ_{max} statistic tests the null hypothesis of r co-integrating vectors against

r+1 co-integrating vectors. Johansen and Juselius (1990) and Osterwald-Lenum (1992) derive the critical values of λ_{trace} and λ_{max} by simulation method.

Our cointegration tests parallel those made by Anari and Kolari (2001) for their study on European, U.S., and Japanese markets. In testing the long-run relation between each pair of stock prices and goods prices, the null hypothesis states that there is no co-integration relation.

The results of the Johansen trace test are provided in table 3. These results suggest that one co-integrating vector (or a long-run relation) between goods prices measured by the consumer price index (CPI) and the stock price index exists in each country under study. We conclude from the similarity in results across several countries in this study that it takes stock prices a long time to return to their long run relation when there is unexpected movement in goods prices, and that the co-integrating tests are robust. This is similar to the case with European, U.S., and Japanese markets as reported by Anari and Kolari (2001) and Khil and Lee (2000).

5.1 Vector Error Correction Model (VEC)

Cointegration exists when a group of nonstationary variables has a linear combination of them that is stationary. Cointegration means that although many developments can cause permanent changes in the individual elements of the group, there is some long-run equilibrium relation tying the individual components together. If the group is cointegrated, then it is not correct to fit a VAR to the differenced data [Hamilton (1994)].

As argued by Engle and Granger (1987), the VAR estimated with cointegrated data (without including the cointegration term) will be misspecified. However, another representation of VAR, the Vector Error Correction model (VEC), can be used. It is a VAR model for data in difference from augmented by the error correction term. In a VEC model the short-run dynamics of the variables in the group are influenced by the deviation from an equilibrium relationship.

As the VEC specification only applies to a cointegrated series, we should run the Johansen cointegration test prior to VEC specification. This test is needed to confirm that the variables are cointegrated and to determine the number of cointegrating equations. Estimation of a VEC model proceeds by first determining one or more cointegrating equations using the Johansen procedure. The first difference of each endogenous variable is then regressed on a one period lag of the cointegrating equation(s) and lagged first differences of all the endogenous variables in the system.

The Johansen efficient maximum likelihood test is used to examine the existence of a long-term relationship between stock prices (SP) and consumer goods prices (CP). It is applied using alternative lag lengths in the VAR. Consider a VAR model of order k:

$$X_t = C + A_1 X_{t-1} + \dots + A_k X_{t-k} + v_t \quad (5)$$

Where C is a 2 x 1 vector of constants, A_k are 2 x 2 matrices of coefficients to be estimated, and vector v_t represents the unexpected movements in SP and CP.

It should be noted the VAR model provides information about the short-run relation between stock prices and inflation. For the estimation of the longrun relation between the two variables, we follow Anari and Kolari (2001) and compare our findings with those reported by Khil and Lee (2000) on Pacific-Basin markets in the short run. Anari and Kolari use the VECtor error-correction (VEC) model by Johansen (1991). Equation (5) can be written as:

$$\Delta x_t = \delta + \Gamma_1 \Delta x_{t-1} + \Gamma_2 \Delta x_{t-2} + \dots + \Gamma_{t-k} \Delta x_{t-k+1} + \Pi x_{t-k} + v_t \quad (6)$$

Where Γ_i and Π are 2x2 matrixes, and k is the lag order. The rank of matrix Π gives the number of cointegrating vectors, which are longrun relations between SP and CP. Anari and Kolari show that the term Πx_{t-k} represents the long-term relation between SP and CP, and a long run relation between SP and CP can be evaluated by:

$$\Delta S_t = \sum_{k=1}^{n-1} a_k \Delta S_{t-k} + \sum_{k=1}^{n-1} b_k \Delta C_{t-k} + e(S_{t-1} - c - dP_{t-1}) \quad (7)$$

Where the summation term represents the short run relation between stock prices and goods prices, and the error correction term e represents the speed of adjustment of stock prices to unexpected changes in inflation. The term, which is the vector of deviations from the longrun relation between stock prices and goods prices, can be normalized and its longrun equation can be expressed as:

$$S_t = c + dP_t \quad (8)$$

If the variables are in log terms, the coefficient (d) in this equation is the elasticity of stock prices with respect to goods prices, otherwise known as the Fisher coefficient (Anari and Kolari, 2001).

Using equation (8), the MLE estimates on longrun relations between stock prices and the CPI for the sample period are shown in table 4. It should be noted that the estimated Fisher coefficients (d) are in the range of 1.01 to 1.36. These coefficients are distributed as follows: Jordan = 1.01, Saudi Arabia = 1.20, Morocco = 1.04, and Kuwait = 1.36. In all countries, the long-run Fisher effect is supported because of the positive signs for the estimated (d) coefficient of CPI in equation (8).

Since all variables are expressed as logarithms, the CPI's coefficient (d) in each equation shows the elasticity of the changes in stock prices with respect to corresponding changes in inflation. For instance, when the estimated coefficient (d) is 1.01 in Jordan, this means that for every increase of 1% in CPI, the stock index is expected to increase by 1.01% over the sample period.

We apply t-tests to examine whether the estimated Fisher coefficient is less than, equal to or greater than unity. The results in table 4 show that estimates for Jordan, Morocco, and Saudi Arabia are greater than unity, whereas those for Kuwait are less than or equal to unity at a 5% level significance.

In table 4, the estimates of the speed of adjustment coefficients (e) lie between 0.01 and 0.03, which means that it takes a long time for stock prices to return to their long-run relation following an unexpected movement in goods prices.

6. Conclusions

This paper attempts to examine monthly stock price indexes and goods price indexes for four Arab countries: Jordan, Saudi Arabia, Morocco, and Kuwait for 2000-2009 using cointegration methods. The results of the cointegration test support the long-run relationship between stock prices and goods prices. The long-run Fisher elasticities of stock prices with respect to goods prices are in the range of 1.01 to 1.36 across the four countries under study.

With the exception of Kuwait the Fisher effect coefficient estimates are significantly greater than one. In the case of Jordan, Saudi Arabia, and Morocco the empirical results support the Fisher hypothesis, with estimated coefficients near unity. The results also reveal that stock prices in the those four Arab markets have a long memory with respect to inflation shocks that make stocks a reasonably good inflation hedge over a long holding period.

In this respect, our findings are similar to the evidence already reported by Anari and Kolari (2001) on American, European and Japanese stock markets.

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Table 1: Unit Root Tests

Country	SPa	SPb	SPc	CPa	CPb	CPc
Jordan						
ADF*	-1.584	-1.598	-0.304	0.861	-1.960	3.275
PP	-1.437	-1.367	-0.206	0.557	-1.896	2.710
Saudi						
ADF*	-1.343	-1.001	-0.357	1.547	-0.275	1.962
PP	-1.566	-1.392	-0.553	3.487	-1.636	2.736
Morocco						
ADF*	-0.285	-1.603	0.991	-0.631	-2.555	3.102
PP	-0.571	-1.846	0.556	0.0949	-3.165	5.216
Kuwait						
ADF*	-1.495	-1.445	-0.329	0.667	-1.274	1.931
PP	-1.401	-1.117	-0.252	3.832	-0.309	4.803

Note: *SP* and *CP* denote stock price and consumer price. The ADF and PP are the Augmented Dickey Fuller and Phillips-Perron unit root test with intercept (a), with trend and intercept (b), and with neither trend nor intercept (c), respectively.

Table 2: Unit Root Tests

Country	SPa	SPb	SPc	CPa	CPb	CPc
Jordan						
ADF*	-5.207	-5.254	-5.202	-8.710	-8.821	-8.180
PP	-1.437	-1.367	-0.206	0.557	-1.896	2.710
Saudi						
ADF*	-9.045	-9.062	-9.071	-2.819	-3.897	-1.246
PP	-1.566	-1.392	-0.553	3.487	-1.636	2.736
Morocco						
ADF*	-5.514	-5.499	-5.442	-9.176	-9.146	-8.376
PP	-0.571	-1.846	0.556	0.0949	-3.165	5.216
Kuwait						
ADF*	-6.179	-6.233	-6.171	-2.441	-4.701	-1.576
PP	-1.401	-1.117	-0.252	3.832	-0.309	4.803

Note: *SP* and *CP* denote stock price and consumer price. The ADF and PP are the Augmented Dickey Fuller and Phillips-Perron unit root test with intercept (a), with trend and intercept (b), and with neither trend nor intercept (c), respectively.

Table 3: Cointegration Tests Based on the Johansen's Trace Test

Hypothesized no. of cointegration vectors	Jordan	Saudi	Morocco	Kuwait	5% critical value
None	11.267	10.689	23.213	41.909	9.96
At most one	2.362	2.108	3.674	2.296	9.24

Likelihood Ratio. Note: The tests show that there is one cointegrating relation between stock price index and CPI in each country.

Table 4: Long-Run Relations between Stock Prices and Inflation Based on the Full Information Maximum Likelihood Estimator (MLE).

$$SP_t = c + dCP_t$$

Country	c	d	e
Jordan	-0.16	1.01* (6.55)	-0.01 (4.88)
Saudi	2.34	1.2* (11.23)	-0.04 (3.65)
Morocco	2.45	1.04* (6.23)	-0.03 (5.28)
Kuwait	2.16	1.36 (1.02)	-0.02 (3.4)

Note: * means fail to accept the null hypothesis (at %5) that the Fisher coefficient (d) is less than or equal to one and instead accept the alternative hypothesis that it is more than one. t -values are in parentheses. The term e is the speed of adjustment which is the rate of convergence to the long-run equilibrium.

Dependent variable: ΔSP			
Variables	Coefficient	t-values	Prob-values
Constant	-0.02937	-0.487251	0.6297
$\Delta SP (-1)$	0.885764	2.686028	0.0118
ΔCP	-0.283966	-0.631094	0.5329
$e_2(-1)$	-1.144181	-3.174132	0.0035
$R^2 = 0.324380$			Akaike info criterion = 0.023627
Adjusted $R^2 = 0.231192$			Schwarz criterion = 0.248092
Durbin-Watson stat = 2.012			F-statistic = 3.48 (0.02)