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Consideration the Relationship between the Inflation and Stock Price with Panel Cointegration Approach

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Abstract
The analytical literature mentions the possibility of a negative and a positive relationship both. The purpose of this study is to examine the relationship between the inflation and stock prices in Asian selected countries for 2000 – 2012 by using quarterly data. In Economics literature many studies tried to examine whether stocks are perfect hedge against inflation. The answer is not conclusive. In this paper, using data from Asian selected countries stock market, the relationship between inflation and stock returns. The long-run relationship is estimated using a full-modified OLS. Pedroni's heterogeneous panel cointegration test reveals a long-run equilibrium relationship between inflation and stock price. The empirical results have shown that Fisher Hypothesis, which asserts that stocks are perfect hedge against inflation, has been rejected and also it is revealed that stocks are a weak hedge against inflation in Asian selected stock market. However in the very long run as we observe from the co-integrating equation, inflation influences stock prices and that too in a positive direction. Also Results obtained indicate the presence of the long run and the short run relationship between inflation and stock price. In the short run and in the long run, inflation is found to be Granger cause of economic stock price, and vice versa.
Keywords: inflation, stocks, money volume, panel cointegration.

1. Introduction

Access to high and steady growth of stock market return associated with low inflation rate is the basic subject of macroeconomic policies. The relationship between inflation and stock price growth, subject that has been discussed in recent decades. As we discussed here, despite the fact that price increases should shield corporations from the effects of inflation, they actually end up eating up much of the profits in asset requirements. But this cash requirement of the business is hidden from investors, who continue to see rising earnings. Therefore, what happens to stock prices during inflationary times?

They key to understanding the argument below is to recognize that as inflation increases, central banks increase interest rates to reduce the money supply and slow inflation down: When interest rates are high, people find it expensive to borrow, and therefore there is less money floating around. With interest rates are high, people require higher returns on stocks. Well, it's not so easy to just increase earnings for a stock, so its price has to adjust downward. However, as we bring in the stock market prices the relationship between price and quantity may turn out to be more complex than a simplistic one, as though usually. The stock market prices may be related to the domestic inflation and even if domestic inflation may not affect quantity produced directly there can be substantial impact of stock market prices on quantity produced. Hence, two important questions that we are bothered about from empirical standpoint are whether domestic inflation and stock market prices are in any manner connected – and if so what is the nature of relationship - and secondly whether stock market prices affect the real variables in a significant way and again if so, what is the nature of relationship? In the developed world the stock market controls the real sector hugely whereas in the Indian context the stock market used to be quite superfluous in this respect. That is because the stock market was controlled by only a few players (Chakravarty and Mitra, 2010).

The early survey on the behaviour of stock return was done by Fama (1970). The Fama theory of efficient market hypothesis suggests that stock markets are efficient because they reflect the fundamental macroeconomic behavior. The term efficiency implies that a financial market incorporates all relevant information (including macroeconomic fundamentals) in the market and thus the observed outcome is the best possible one under the circumstances. Chakravarty (2006) explore the relationship between stock price and some key macro variables and gold price in India for the period 1991-2005. The study used Granger non-causality test procedure developed by Toda and Yamamoto (1995).Bhattacharya and Mukherjee (2002) showed a two-way causation between stock price and the rate of inflation, while index of industrial production lead the stock price. Studies suggesting a negative relationship between stock prices and inflation (Fama, 1981) envisage that high inflation predicts an economic downturn and keeping in view this the firms start selling off their stock. An increase in the supply of stock then reduces the stock prices. Since stocks reflect firms' future earning potential an expected economic downturn prompts firms to sell off the financial stocks and thus high inflation and low stock prices tend to go together. On the other hand, a positive relationship is also possible between inflation and stock prices as unexpected inflation raises the firms' equity value if they are net debtor (Chakravarty and Mitra, 2013).

The purpose of this paper is to empirically examine the long-run and the short run relationship and the causal relationship between inflation and stock price in Asian selected countries. We combine cross-sectional and time series data to examine the relationship between inflation and stock price, using updated data for selected Asian countries such as: Iran, Japan, china, Hong – Kong, Australia, Korea, Qatar, United Arab Emirate, Oman and Saudi Arabia for the years 2000–2012.

The paper is organized as follows: In Section 2 reviews the existing literature. Section 3 we provide a brief discussion of the panel unit root test and the panel cointegration procedure. Empirical results are provided in Section 4. Final section contains the conclusions and policy implications.

2. Literature review

Xin Liu (2014) examined the relationship between inflation and stock price. He found that stock price has a positive impact on inflation rate, though in a subtle, complex way. Chakravarty and Mitra (2013) examined the relationship between inflation and stock price. This study examines the nature of relationship between inflation and stock price movement. The analytical literature mentions the possibility of a negative and a positive relationship both. Using the VAR framework based on monthly data for wholesale price index, index of industrial production, exchange rate, stock prices and foreign institutional investment we note that stock prices have an impact on inflation whereas the causality in the reverse direction is not prominent. The results from the impulse response function tend to suggest that the nature of relationship is rather negative. When stock prices are low the firms are reluctant to tap the capital market. Unless bank finance can substitute adequately for the capital market firm's investment plans would be hit and production would decline. This may result in a price rise as the market demand may exceed the supply. An important policy implication is augmentation of production by encouraging investment through inexpensive bank finance. However in the very long run as we observe from the co-integrating equation, inflation influences stock prices and that too in a positive direction. Unexpected inflation raises the firm's equity value if they are net debtor. Similarly tightening of monetary policy can reduce inflation and stock prices both as individuals will be left with less money to buy goods or buy stocks.Geetha and et al (2011) in their paper entitled the relationship between inflation and stock market: evidence from Malaysia, USA and China considered the relationship between inflation and stock market price. The study aims to find the relationship between inflation and stock returns. Inflation was distinguished as expected and unexpected inflation. The study revealed that there is long run relationship between expected and unexpected inflation with stock returns but there is no short run relationship between these variables for Malaysia and US but it exists for China. Based on the data for the Greek economy Ioannidis et al. (2005), used ARDL cointegration technique in conjunction with Granger causality tests to detect possible long-run and short-run effects between inflation and stock market prices and also the direction of these effects. The results provide evidence in favour of a negative long-run causal relationship between the series after 1992. In the context of Turkish economy the coefficients of IPI and CPI do not turn out to statistically significant in the equation for stock prices implying that they do not explain the stock prices (Aga and Kocaman, 2006). The stock traders are made up of professional traders who buy and sell shares all day long, hoping to profit from changes in share prices. They are not really interested in the long-term profitability or the value of assets of the company. When traders believe that others will buy shares (in the expectation that prices will rise), then they will buy as well, hoping to sell when the price actually rises. If others believe the same thing, then the wave of buying pressure will, in fact, cause the price to rise (Aga and Kocaman, 2006). Thus the stock demand and so also the stock prices rise when the economy is about to enter an upswing and on the other hand they all fall when the economy is about to experience a downswing. Thus just before the upswing occurs an increased stock price and a modest inflation can coincide and similarly just before the downswing starts a depressed stock price accompanied by a high inflation may co-exist. In the Indian context the growth boom since 2003-04 has been accompanied by a rise in savings and investment rate of the corporate sector, stock price increase, foreign investment and so on. The financial and monetary market policies must try to keep in view the private investment that is required to maintain the growth tempo (see Desai, 2011). For the structural development of the capital market and for growth to take place it is important that the RBI's monetary policy must look into the issue of inflation management (Desai, 2011). Price stability should be the main goal of the monetary policy because it is only slow and stable inflation which is conducive to growth (Chakravarty and Mitra, 2013).

3. Data and Methodology

Panel data provide a large number of point data, increasing the degrees of freedom and reducing the collinearity between regressors. Therefore, it allows for more powerful statistical tests and normal distribution of test statistics. It can also take heterogeneity of each cross-sectional unit into account, and give "more variability, less collinearity among variables, more degrees of freedom, and more efficiency" (Baltagi, 2001). In this paper, regressions are based on data concerning a group of Asian countries over the period 2000 - 2012. Inflation and stock price are a common statistic for representing the level of a particular country within a certain time range. Our analysis is based on quarterly data from 2000 to 2012 drawn from the following sources: latest RBI's weekly statistical bulletin, website and press reports, WDI.

3.1. Unit root test

In order to investigate the possibility of panel cointegration, first, it is necessary to determine the existence of unit roots in the data series. For this study we have chosen the Im, Pesaran and Shin (IPS, hereafter), which is based on the well-known Dickey-Fuller procedure.

Im, Pesaran and Shin denoted IPS proposed a test for the presence of unit roots in panels that combines information from the time series dimension with that from the cross section dimension, such that fewer time observations are required for the test to have power. Since researchers have found the IPS test to have superior test power for analyzing long-run relationships in panel data, we will also employ this procedure in this study. IPS begins by specifying a separate ADF regression for each cross-section with individual effects and no time trend:

$$\Delta y_{it} = \alpha_i + \rho_i y_{i,t\ 1} + \sum_{j=1}^{pi} \beta_{ij} \Delta y_{i,t\ j} + \varepsilon_{it} \quad (1)$$

Where $i = 1, \ldots, N$ and $t = 1, \ldots, T$

IPS use separate unit root tests for the N cross-section units. Their test is based on the Augmented Dickey-fuller (ADF) statistics averaged across groups. After estimating the separate ADF regressions, the average of the t-statistics for p_1 from the individual ADF regressions, $t_{iT_i}(p_i)$:

$$\bar{t}_{NT} = \frac{1}{N} \sum_{i=1}^{N} t_{iT}\ (p_i \beta_i) \quad (2)$$

The t-bar is then standardized and it is shown that the standardized *t-bar* statistic converges to the standard normal distribution as N and T $\longrightarrow \infty$. IPS (1997) showed that *t-bar* test has better performance when N and T are small. They proposed a cross-sectional demeaned version of both test to be used in the case where the errors in different regressions contain a common time-specific component (Nor'Aznin and et al, 2010).

3.2. Panel Cointegration Tests

The next step is to test for the existence of a long run relationship among inflation and stock price. A common practice to test for cointegration is Johansen's procedure. However, the power of the Johansen test in multivariate systems with small sample sizes can be severely distorted. To this end, we need to combine information from time series as well as cross-section data once again. In this context three panel cointegration tests are conducted.

First, we use a test due to Levin and Lin (1993) in the context of panel unit roots, to estimate residuals from (supposedly) long run relations. Levin and Lin (1993) consider the model

$$y_{it} = \rho_i y_{i,t-1} + z_{it}' \gamma + u_{it} \quad (3)$$

Where z_{it} are deterministic variables, u_{it} is iid$(0,\sigma^2)$ and $\rho_i = \rho$. The test statistic is at t-statistic on ρ given by

$$t_\rho = \frac{(\hat{\rho}-1)\sqrt{\sum_{i=1}^{N}\sum_{t=1}^{T} \tilde{y}_{i,t-1}^2}}{s_e} \quad (4)$$

Where

$$\tilde{y}_{it} = y_{it} - \sum_{s=1}^{T} h(t,s) y_{is}, \quad \tilde{u}_{it} = u_{it} - \sum_{s=1}^{T} h(t,s) u_{is} \quad h(t,s) = z_t' \left(\sum_{t=1}^{T} z_t z_t' \right) z_s,$$

$$s_e^2 = (NT)^{-1} \sum_{i=1}^{N} \sum_{t=1}^{T} \tilde{u}_{it}^2,$$

And $\hat{\rho}$ is the OLS estimate of ρ. It must be noted that Levin and Lin (1993) tests may have substantial size distortion if there is cross-sectional dependence (O'Connell, 1998). Also, Harris and Tzavalis (1999) find that small T yields Levin and Lin tests which are substantially undersized and have low power. A drawback of the Levin and Lin or Harris and Tzavalis tests is that they do not allow for heterogeneity in the autoregressive coefficient, ρ.

Finally, to overcome the problem of heterogeneity that arises in both tests we use Fisher's test to aggregate the p-values of individual Johansen maximum likelihood cointegration test statistics, see Maddala and Kim (1998). If p_i denotes the p-value of the Johansen statistic for the ith unit, then we have the result $-2\sum_{i=1}^{N} \log p_i \sim \chi_{2N}^2$. The test is easy to compute and, more importantly, it does not assume homogeneity of coefficients in different countries (Christopoulos and Tsionas, 2004).

The next step is to test for the existence of a long-run cointegration inflation and stock price using panel cointegration tests suggested by Pedroni (1999 and 2004). We will make use of seven panel cointegration by Pedroni (1999), since he determines the appropriateness of the tests to be applied to estimated residuals from a cointegration regression after normalizing the panel statistics with correction terms (Nor'Aznin and et al, 2010). The procedures proposed by Pedroni make use of estimated residual from the hypothesized long-run regression of the following form:

$$y_{i,t} = \alpha_i + \delta_i t + \beta_{1i}x_{1i,t} + \beta_{2i}x_{2i,t} + \cdots + \beta_{Mi}x_{Mi,t} + \varepsilon_{i,t} \quad (5)$$

For t = 1,…..,T; I = 1,….,N; m = 1, …., M,

Where T is the number of observations over time, N number of cross-sectional units in the panel, and M number of regressors. In this set up, α_i is the member specific intercept or fixed effects parameter which varies across individual cross-sectional units. The same is true of the slope coefficients and member specific time effects, $\delta_i t$. Pedroni (1999 and 2004) proposes the heterogeneous panel and heterogeneous group mean panel test statistics to test for panel cointegration. He defines two sets of statistics. The first set of three statistics $Z_{\hat{v},N,T}$, $Z_{\hat{\rho}N,T\ 1}$ and $Z_{tN,T}$ are based on pooling the residuals along the within dimension of the panel. The statistics are as follows

$$Z_{\hat{v},N,T} = T^2 N^{3/2} \sum_{i=1}^{N}\sum_{t=1}^{T} \hat{L}_{11i}^2\, \hat{e}_{i,t\ 1}^2 \quad (6)$$

$$Z_{\hat{\rho},N,T\ 1} = T\sqrt{N} \sum_{i=1}^{N}\sum_{t=1}^{T} \hat{L}_{11i}^2\, \hat{e}_{i,t\ 1}^2 \quad \sum_{i=1}^{N}\sum_{t=1}^{T} \hat{L}_{11i}^2\, (\hat{e}_{i,t\ 1}\, \Delta\hat{e}_{i,t}\ \ \hat{\lambda}_i) \quad (7)$$

$$Z_{tN,T} = \tilde{\sigma}_{N,T}^2 \sum_{i=1}^{N}\sum_{t=1}^{T} \hat{L}_{11i}^2\, \hat{e}_{i,t\ 1}^2{}^{1/2} \sum_{i=1}^{N}\sum_{t=1}^{T} \hat{L}_{11i}^2\, (\hat{e}_{i,t\ 1}\, \Delta\hat{e}_{i,t}\ \ \hat{\lambda}_i) \quad (8)$$

Where $\hat{e}_{i,t\ 1}$ is the residual vector of the OLS estimation of Equation (5) and where the other terms are properly defined in Pedroni. The second set of statistics is based on pooling the residuals along the between dimension of the panel. It allows for a heterogeneous autocorrelation parameter across members. The statistics are as follows:

$$\tilde{Z}_{\hat{\rho}N,T\ 1} = \sum_{i=1}^{N}\sum_{t=1}^{T} \hat{e}_{i,t\ 1}^2{}^{1} \sum_{t=1}^{T}(\hat{e}_{i,t\ 1}\, \Delta\hat{e}_{i,t}\ \ \hat{\lambda}_i) \quad (9)$$

$$\tilde{Z}_{tN,T\ 1} = \sum_{i=1}^{N}\sum_{t=1}^{T} \hat{e}_{i,t\ 1}^2{}^{1/2} \sum_{t=1}^{T}(\hat{e}_{i,t\ 1}\, \Delta\hat{e}_{i,t}\ \ \hat{\lambda}_i) \quad (10)$$

These statistics compute the group mean of the individual conventional time series statistics. The asymptotic distribution of each of those five statistics can be expressed in the following form:

$$\frac{X_{N,T}\ \mu\sqrt{N}}{\sqrt{V}} \rightarrow N(0,1) \quad (11)$$

Where $X_{N,T}$ is the corresponding form of the test statistics, while μ and v are the mean and variance of each test respectively. They are given in Table 2 in Pedroni (1999). Under the alternative hypothesis, Panel v statistics diverges to positive infinity. Therefore, it is a one sided test were large positive values reject the null of no cointegration. The remaining statistics diverge to negative infinity, which means that large negative values reject the null (Al-Awad and Harb, 2005).

3.3. Fully Modified Ordinary Least Squares (FMOLS) Estimation

In this section we adopt FMOLS procedure from Christopoulos and Tsionas (2004). In order to obtain asymptotically efficient and consistent estimates in panel series, non-exogeneity and serial correlation problems are tackled by employing fully modified OLS (FMOLS) introduced by Pedroni (1996). Since the explanatory variables are cointegrated with a time trend, and thus a long-run equilibrium relationship exists among these variables through the panel unit root test and panel cointegration test, we proceed to estimate the Equation (2) by the method or fully modified OLS (FMOLS) for heterogeneous cointegrated panels. This methodology allows consistent and efficient estimation of cointegration vector and also addresses the problem of non-stationary regressors, as well as the problem of simultaneity biases. It is well known that OLS estimation yields biased results because the regressors are endogenously determined in the I(1) case. Following cointegrated system for panel data

$$y_{it} = \alpha_i + x_{it}'\beta + u_{it} \quad (12)$$

$$x_{it} = x_{i,t-1} + e_{it} \quad (12)$$

Where $\xi_{it} = [u_{it}, e_{it}']$ is stationary with covariance matrix Ω_i. Following Phillips and Hansen (1990) a semi-parametric correction can be made to the OLS estimator that eliminates the second order bias caused by the fact that the regressors are endogenous. Pedroni (2000) follows the same principle in the panel data context, and allows for the heterogeneity in the short run dynamics and the fixed effects. Pedroni's estimator is (Christopoulos and Tsions, 2004)

$$\hat{\beta}_{FM} - \beta = \left(\sum_{t=1}^{N} \hat{\Omega}_{22i}^{-2} \sum_{t=1}^{T}(x_{it} - \bar{x}_t)^2\right)^{-1} . \sum_{i=1}^{N} \hat{\Omega}_{11i}^{-1}\, \hat{\Omega}_{22i}^{-1} \left(\sum_{t=1}^{T}(x_{it} - \bar{x}_t)u_{it}^* - T\hat{\gamma}_i\right) \quad (13)$$

$$\hat{u}_{it}^* = u_{it} - \hat{\Omega}_{22i}^{-1}\hat{\Omega}_{21i},\ \ \hat{\gamma}_i = \hat{\Gamma}_{21i} + \hat{\Omega}_{21i}^0 - \hat{\Omega}_{22i}^{-1}\hat{\Omega}_{21i}(\hat{\Gamma}_{22i} + \hat{\Omega}_{22i}^0) \quad (14)$$

Where the covariance matrix can be decomposed as $\Omega_i = \Omega_i^0 + \Gamma_i + \Gamma_i$ where Ω_i^0 is the contemporaneous covariance matrix, and Γ_i is a weighted sum of autocovariances. Also, $\hat{\Omega}_i^0$ denotes an appropriate estimator of Ω_i^0. In this study, we employed panel group FMOLS test from Pedroni (1996, 2000). An important advantage of the panel group estimators is that the form in which the data is pooled allows for greater flexibility in the presence of heterogeneity of the cointegrating vectors. Test statistics constructed from the panel group estimators are designed to test the null hypothesis $H_0: \beta_i = \beta_0$ for all I against the alternative hypothesis $H_A: \beta_i \neq \beta_0$, so that the values for β_i are not constrained to be the same under the alternative hypothesis. Clearly, this is an important advantage for applications such as the present one, because there is no reason to believe that if the cointegrating slopes are not equal to one, which they necessarily take on some other arbitrary common value. Another advantage of the panel group estimators is that the point estimates have a more useful interpretation in the event that the true cointegrating vectors are heterogeneous. Specifically, point estimates for the panel group estimator can be interpreted as the mean value for the cointegrating vectors (Nor'Aznin and et al, 2010).

4. Empirical result

Table 1 presents the results of the IPS panel unit root test at level indicating that all variables are $I(1)$ in the constant of the panel unit root regression. These results clearly show that the null hypothesis of a panel unit root in the level of the series can be rejected at various lag lengths. We assume that there is no time trend. Therefore, we test for stationarity allowing for a constant plus time trend. In the absence of a constant plus time trend, again we found that the null hypothesis of having panel unit root is generally not rejected in all series at level form and various lag lengths. We can conclude that most of the variables are non-stationary in with and without time trend specifications at level by applying the IPS test which is also applied for heterogeneous panel to test the series for the presence of a unit root. The results of the panel unit root tests confirm that the variables are non-stationary at level.

Table 1: Panel Unit Root Test – Im, Pesaran and Shin (IPS)

Variable	Level		First order difference	
	Constant	Constant + Trend	Constant	Constant + Trend
Inflation	**-1.232** (0.786)	**-1.244** (0.375)	**-3.790** (0.000)	**-3.476** (0.001)
Stock Price	**-1.246** (0.674)	**-1.947** (0.865)	**-3.840** (0.000)	**-3.485** (0.002)

Note: Levels and first order differences denote the IPS t-test for a unit root in levels and first differences respectively. Number of lags was selected using the AIC criterion. Boldface values denote sampling evidence in favour of unit roots. We use the Eviews software to estimate this value.

Table 1 also presents the results of the tests at first difference for IPS test in constant and constant plus time trend. We can see that for all series the null hypothesis of unit root test is rejected at 95 percent critical value. Hence, based on IPS test, there is strong evidence that all the series are in fact integrated of orders one.

We can conclude that the results of panel unit root tests reported in Table1 support the hypothesis of a unit root in all variables across countries, as well as the hypothesis of zero order integration in first differences. At most of the 1 percent significance level, we found that all tests statistics in both with and without trends significantly confirm that all series strongly reject the unit root null. Given the results of IPS test, it is possible to apply panel cointegration method in order to test for the existence of the stable long-run relation among the variables.

Table 2: The Pedroni Panel Cointegration Test

Test	**Constant trend**	**Constant + Trend**
Panel *v*-Statistic	1.000	0.988
Panel *ρ*-Statistic	0.000	0.000
Panel *t*-Statistic: (non-parametric)	0.988	0.983
Panel *t*-Statistic (*adf*): (parametric)	0.001	0.002
Group *ρ*–Statistic	0.999	1.000
Group *t*-Statistic: (non-parametric)	0.000	0.000
Group *t*-Statistic (*adf*): (parametric)	0.000	0.000

Note: All statistics are from Pedroni's procedure (1999) where the adjusted values can be compared to the N(0,1) distribution. We use the Eviews software to estimate this value.

The next step is to test whether the variables are cointegrated using Pedroni's (1999, 2001, and 2004). This is to investigate whether long-run steady state or cointegration exist among the variables and to confirm of what Coiteux and Olivier (2000) state that the panel cointegration tests have much higher testing power than conventional cointegration test. Since the variables are found to be integrated in the same order $I(1)$, we continue with the panel cointegration tests proposed by Pedroni (1999, 2001, and 2004). Cointegration are carried out for constant and constant plus time trend and the summary of the results of cointegration analyses are presented in

Table 2.

In constant level, we found that that 4 out of 7 statistics reject the null hypothesis of no cointegration at the 1 percent and 5 percent level of significance. It is shown that independent variables do hold cointegration in the long run for these countries.

Table 3: FMOLS Regression (Dependent variable: Stock Price)

Country	Inflation
Iran	0.12 (0.00)
Japan	0.54 (0.01)
china	0.22 (0.02)
Hong – Kong	0.29 (0.00)
Australia	0.25 (0.00)
Korea	0.19 (0.01)
Qatar	0.65 (0.02)
United Arab Emirate	0.18 (0.03)
Oman	0.44 (0.00)
Saudi Arabia	
Panel Group	0.38 (0.00)

Note: The null hypothesis for the *t*-ratio is $H_0=\beta_i=0$; Figures in parentheses are *t*-statistics. (*) indicate significance at the 5% level. We use the Eviews software to estimate this value.

In Table 3, we found that the estimate of the coefficient for inflation is positive and statistically significant at the 5 percent level for all countries, we conclude that there is a long run relationship between inflation and stock price for these counties.

The next step is to estimate the Granger causality model with a dynamic error correction:

$$\Delta CPI_{it} = \alpha_{1j} + \sum_{i=1}^{k} \beta_{11ik} \Delta CPI_{it-k} + \sum_{i=1}^{k} \beta_{12ik} \Delta STOCKPRICE_{it-k} + \delta_{1i} ECT_{it-1} + u_{1it} \quad (15)$$

$$\Delta STOCKPRICE_{it} = \alpha_{2j} + \sum_{i=1}^{k} \beta_{21ik} \Delta CPI_{it-k} + \sum_{i=1}^{k} \beta_{22ik} \Delta STOCKPRICE_{it-k} + \delta_{2i} ECT_{it-1} + u_{2it} \quad (16)$$

Where Δ denotes first differencing and k is the lag length and is chosen optimally for each country using a step-down procedure up to a maximum of two lags.

The sources of causation can be identified by testing for the significance of the coefficients of the dependent variables in Eqs. (15) and (16). First, the short-run effect can be considered transitory. For short-run causality, we can test $H_0: \beta_{12ik} = 0$ for all i and k in Eq. (15) or $H_0: \beta_{21ik} = 0$ for all i and k in Eq. (16). Next, the long-run causality can be tested by looking at the significance of the speed of adjustment δ, which is the coefficient of the error correction term, ECT_{it-1}. The significance of k indicates the long-run relationship of the cointegrated process, and so movements along this path can be considered permanent. For long-run causality, we can test $H_0: \delta_{1i} = 0$ for all i in Eq. (15), $H_0: \delta_{2i} = 0$ for all i in Eq. (16). Finally, we can use the joint test to check for a strong causality test, where variables bear the burden of a short-run adjustment to re-establish a long-run equilibrium, following a shock to the system (Asafu-Adjaye, 2000; Chien-Chiang Lee, 2005; Abbasinejad and et al, 2012). Because all variables enter the model in stationary form, a standard F-test can be used to test the null hypothesis, which shows that all of the estimated country-specific parameters are significant. We find that a panel causality test between inflation and stock price. An examination of the sum of the lagged coefficients on the respective variables indicates that stock price (0.02) has a statistically significant impact on inflation. Moreover, the error correction term is statistically significant at the 5% level denoting a relative high speed of adjustment to long-run equilibrium. In terms of Eq. (15), it appears that inflation has a statistically significant impact on stock price in the short-run and the long run. According to F-statistic the results reveal bidirectional Granger causality between mentioned variables. However, the long run bidirectional causality between these variables is more significant than the short run causality between them.

5. Conclusion

This paper is an empirical study on the relationship between inflation and stock price in selected Asian countries' stock market. For that reason we use the panel cointegration approach. The unit root test (IPS) is used to confirm the stationarity of all variables before the cointegration test can be performed. After confirming that all variables are non-stationary at level, the panel cointegration approach is applied. Using Pedroni's, the long run cointegration test is performed to investigate the existence of the long run cointegration among the variables. Results obtained indicate the presence of the long run and the short run relationship between inflation and stock price. In the short run and in the long run, inflation is found to be Granger cause of economic stock price, and vice versa. The results of a bidirectional relationship in the short-run and in the long run show that inflation leads stock price. Our results support that current as well as past changes in inflation have significant impact on the changes in stock price in these countries. It is clear for these countries in general that in short run inflation is an important ingredient for stock market growth. The results generally show that in the long run, inflation is positive and significantly correlated with stock price of these countries.

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