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Trade Openness and the Behaviour of Stock Prices in Iran

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Abstract
This paper investigates the relationship between trade openness, banking sector developments, real GDP and Tehran Stock Exchange (TSE) price index covering the period of 1993Q1– 2011Q4. Previous studies have investigated the direct effects of macroeconomic variables on the stock prices in Iran. This study for the first time in previous literature in Iran examines the direct effects of trade openness on stock prices in Iran. We have tested an econometric model of stock prices in compliance with Basu and Morey (2005). The ARDL bounds testing approach is applied to examine the cointegration. The error correction model (ECM) is also applied to estimate the model in short-run. Our empirical findings confirm the existence of cointegration between the series. We find that trade openness has a negative effect on TSE price index in longrun and short-run. Moreover, banking sector developments and real GDP have a positive impact on TSE price index in long-run and short-run.
Keywords: Stock Prices; Trade Openness, Banking Sector Development

1. Introduction

The effect of macroeconomic variables on the stock prices is a well-established theory in the financial economics literature. Numerous studies (Chen, 1991; Chen *et al.*, 1986; Fama, 1991; Huang and Kracaw, 1984; Pearce and Roley, 1988; Wei and Wong, 1992; Lanne, 2002; Lewellen, 2003; Campbell and Yogo, 2003; Janson and Moreira, 2004, Donaldson and Maddaloni, 2002, Goyal and Yamada , 2004; and Ang and Maddaloni, 2005) have been modeled the relation between asset prices and real economic activities in terms *GDP* growth, industrial production rate, short-term interest rate, inflation rate, interest rate spread, exchange rate, current account balance, unemployment rate, fiscal balance, etc. More studies are focused on the developed countries such as the US, UK and Japan. Surprisingly, there has been very few works done on the relationships between trade openness and stock prices (i.e., Li *et al*., 2004; Basu and Morey, 2005).

This study contributes to this literature by examining the dynamic links between trade openness, banking sector developments, real *GDP* and stock prices in Iran. Our study is different from the previous studies in Iran in two ways. The number of empirical studies has been conducted investigating the direct effects of macroeconomic variables on the stock prices in Iran (e.g. Foster and Kharazi, 2008; Safdari *et al*., 2011). We, for the first time in previous literature in Iran, examine the direct effects of trade openness on stock prices in Iran. A second way our study is different is as follows. We investigate the impact of financial opening and trade openness together on stock prices in long-run and short-run. Basu and Morey (2005) indicate financial opening alone without trade opening will not lead to efficiency in the stock prices. We will test this hypothesis on stock prices in Iran.

The rest of the paper is as follows: Section 2 reviews existing literature; Section 3 describes econometric modeling and estimation techniques; Section 4 deals with the data and empirical analysis and Section 5 concludes the study.

2. Literature Review

According to the literatures, several theories explain the effects of macroeconomic variables on stock prices. Among these theories are the efficient market hypothesis (*EMH*) and the arbitrage price theory (*APT*). The *EMH* advocates that stock market prices fully and rationally incorporate all relevant information. The basic idea underlying the *EMH* developed by Fama (1965, 1970) is that asset prices promptly reflect all available information such that abnormal profits cannot be produced regardless of the investment strategies utilized. The theory of asset pricing, in general, demonstrates how assets are priced given the associated risks. The *APT* suggested by Ross (1976) has been an influential form of asset price theory. *APT* is a general form of Sharpe’s (1964) capital asset price model (*CAPM*). While the *CAPM* suggests that a single common factor drives asset prices or expected returns, the *APT* advocates that they are driven by multiple macroeconomic factors.

In the last three decades, numerous studies have examined the dynamic relationships between stock market behaviour and economic activity, particularly for developed stock markets such as the U.S (Hashemzadeh and Taylor, 1988; Malliaris and Urrutia, 1991; Abdullah and Hayworth, 1993; Dhakal, Kandil, and Sharma, 1993; Sadorsky, 1999; Ratanapakorn and

Sharma, 2007), United Kingdom (UK) (Thornton, 1993; Abdullah, 1998), Germany (Thornton, 1998), and Japan (Kim and Morenom, 1994; Mukherjee and Naka,1995). Next, several studies have investigated this relationship for developing countries such as Malaysia (Ibrahim, 1999; Ibrahim, 2006), Pakistan (Zafar *et al*, 2008; Hasan and Tarij, 2009). Finally, other studies have emphasized comparisons of developing countries and developed countries, or of developing against developed countries. For example, Singapore and U.S. (Keung *et al*., 2006), U.S. and Japan (Humpe and Macmillan, 2009), U.S., Germany, UK, and Canada (Najand and Rahman, 1991), Hong Kong, Singapore, South Korea, Taiwan, and Thailand (Wenshwo, 2002) and Saudi Arabia, Kuwait, and Bahrain (Malik and Hammoudeh, 2007). However, studies in this area are different in terms of their hypotheses and the methods used.

The *EMH* and *APT* are silent about which precise events or economic factors likely influence asset prices. This silence opens the door to investigating a wide range of relevant events both at the microeconomic and macroeconomic levels of a stock market. Also, according to the 'intuitive financial theory', various macroeconomic variables affect stock market behaviour (Maysami and Koh, 2000; Gjerde and Sættem, 1999). Several theoretical and empirical frameworks that try to explain the fluctuations of stock prices and macroeconomic variables are interested in finding a high-frequency, statistical relationship between these variables.

Many studies examined the relationship between monetary policy and stock prices (e.g., Ratanapakorn and Sharma 2007; Rahman and Mustafa 2008; Humpe and Macmillan 2009). Some studies investigate the short-run and long-run relationship between stock prices and exchange rates (e.g., Aggarwal 1981; Soenen and Hennigar 1988; Ajayi and Mougoue 1996; Salifu *et al*., 2007). Several empirical frameworks have been tested in an effort to explain the relationships between inflation and stock prices (e.g., Fama 1981; Geske and Roll 1983; Lee 1992). The stock market-output nexus has also been extensively studied (e.g., Fama 1990, 1991; Geske and Roll 1983; Kwon and Shin, 1999; Laopodis and Sawhney, 2002).

There is a huge body of literature that analyses the effects of trade openness, especially its impacts on the economic performance of developing countries (for a survey, see Santos-Paulino, 2005). Although the scope of research in recent years has moved beyond the goods markets to the financial sector (see, for example, Braun and Raddatz, 2008; Baltagi *et al*., 2009), only two studies (i.e., Li *et al*., 2004; Basu and Morey, 2005) explicitly examine the association between trade liberalization and stock market informational efficiency in developing countries.

3. Methodology and data

The aim of this study is to evaluate the relationship of trade openness, real *GDP* and banking sector developments with *TSE* price index in case of Iran. In doing so, many studies are examined the effect of real *GDP* on the stock market (see, Chen *et al*. 1986; Rousseau and Wachtel, 2000). Basu and Morey (2005) develop a model that explores the effect of trade openness on stock price behavior. The model predicts that stock returns show a non-zero serial correlation in a closed economy. However, once the country opens on the trade front, the stock returns show zero serial correlation. The model also establishes that financial opening alone without trade opening will not lead to a gain in efficiency in the stock prices.

Following previous studies, we have developed an econometric model of the stock price in compliance with Basu and Morey (2005). We use real *GDP*, trade openness, banking sector developments and *TSE* price index within a univariate framework in case of Iran. The general form of stock price model is constructed as follows:

$$SI = f\,(GDP, BD, OT) \tag{1}$$

We have transformed all the variables into natural-log form to make Eq. (2) estimable. All variables are employed with their natural logarithms form to reduce heteroskedasticity. The estimable form of equation is modeled as follows

$$\ln SI_t = \beta_0 + \beta_1 LnGDP_t + \beta_2 lnBD_t + \beta_3 \ln OT_t + U_t \tag{2}$$

Where, *SI* is *TSE* price index, *GDP* is real gross domestic product, *BD* is measure of banking sector development (banking sector credit available to the private sector as a percentage of *GDP*) and *OT* measure of trade openness (exports plus imports divided by *GDP*). The all of time series data are taken from the Central Bank of Iran (*CBI*) online database for the 1993Q1–2011Q4 period.

We employ the autoregressive distributed lag (*ARDL*) bounds testing approach to cointegration developed by Pesaran and Shin (1999) and Pesaran *et al*. (2001) to explore the existence of long-run relationship between trade openness, banking sector development, real *GDP* and *TSE* prices. This approach has numerous advantages in comparison with other cointegration methods such as Engle and Granger (1987), and Johansen and Juselius (1990) procedures. The bounds testing approach is applicable irrespective of whether variables are I(0) or I(1). Moreover, it is efficient estimator even if samples are small and some of the regressors are endogenous. Finally, a dynamic unrestricted error correction model (*UECM*) can be derived from the *ARDL* bounds testing through a simple linear transformation. The *UECM* integrates the short-run dynamics with the long-run equilibrium without losing any long-run information. The *ARDL* model for Eq. (2) may follow as:

$$\Delta Ln\, SI_t = \alpha + \sum_{i=1}^{n} \beta_{i0}\, \Delta\, LnSI_{t-i} + \sum_{i=1}^{n} \beta_{i1} \Delta LnGDP_{t-i} + \sum_{i=1}^{n} \beta_{i2} \Delta LnBD_{t-i} + \sum_{i=1}^{n} \beta_{i3} \Delta Ln\, OT_{t-i} + \gamma_1 Ln\, SI_{t-1} + \gamma_2 Ln\, GDP_{t-1} + \gamma_3 Ln\, BD_{t-1} + \gamma_4 Ln\, OT_{t-1} + \varepsilon_t \tag{3}$$

Where Δ and ε_t are the first difference operator and the white noise term, respectively. The optimal lag structure of the first-differenced regression is selected Akaike information criterion (*AIC*) and Schwarz Bayesian Criterion (*SBC*). The bound testing procedure is based on the joint *F*-statistic is that tested the null of no cointegration. For example, H0: $\gamma_r \neq 0$, against the alternative of H1: $\gamma_r \neq 0$, r = 1, 2, …, 4. Accordingly, in bounds test two set of critical values (lower and upper critical bounds) compute for a given significance level. Lower critical bound is applied if the regressors are I(0) and the upper critical bound is used for I (1). If the calculated *F*statistics exceeds the upper critical value, we conclude in favour of a long-run relationship. If the calculated *F*-statistics is below the upper critical value, we cannot reject the null hypothesis of no cointegration. However, if it lies between the bounds, a conclusive inference cannot be made without knowing the order of integration of the underlying regressors. To check the robustness of the *ARDL* model, we apply diagnostic tests. The diagnostics tests are checking for normality of error term, serial correlation, autoregressive conditional heteroskedasticity, White heteroskedasticity and the functional form of the empirical model.

4. Empirical results

The first step is to find integrating properties of the variables before proceeding to the *ARDL* bounds testing approach to cointegration for a long run relationship. It is necessary to test the stationarity properties of the variables to ensure that none of the variables is stationary at I(2) or beyond that order of integration. In doing so, Phillips and Perron (1988) and augmented Dickey Fuller (1979) unit root tests are applied to examine the unit root properties of the variables. The results of *PP* and *ADF* unit root tests are presented in Table 1. The unit root analysis indicates that all the series are non-stationary at their level form with intercept and trend. At 1st differenced level, Tehran Stock Exchange (*TSE*) price index, a measure of banking sector development, a measure of trade openness and real *GDP* are integrated. This implies that all the variables are integrated at I(1).

Table 1: Results of Unit Root Test

Series	Order	PP [1]	ADF [2]
LnSI	Level 1st difference	-1.9570 -3.2197	-2.2283 -4.2880
LnBD	Level 1st difference	-2.6452 -3.3519	-2.6131 -2.9227
LnOT	Level 1st difference	-3.0494 -3.2829	-2.4386 -4.2260
LnGDP	Level 1st difference	-1.6542 -3.2037	-2.2834 -3.7712

1 Phillips and Perron (1988) and 2 Augmented Dickey-Fullers (1979)

Note: *, ** and *** represent significance at 1, 5 and 10% level respectively

The selection of lag to *ARDL* procedure is a very important step. Thus, before proceeding to the *ARDL* bounds testing, appropriate lag length of the variables should be selected by using *AIC* and *SBC* criterions. It is pointed out by Lütkepohl (2006) that *AIC* lag length criteria provide efficient and consistent results to capture dynamic relation. In this condition, maximum lags will be determined by a researcher with respect to sample size. Given the Quarterly data available for estimation, we set the maximum lag order of the various variables in the model equal to four. So, using *AIC* and *SBC* criteria, optimal lag length of the variables is 4 which are reported in Table 2 as follows:

Table 2: Selection of Lag Length Criteria

Order	LL	AIC	SBC	LR test	Adjusted LR test
4	860.747	792.7470	715.8159	……….	………
3	801.8860	749.8860	691.0563	CHSQ(16) = 72.3557[.000]	52.0626[.000]
2	713.3342	677.3342	636.6059	CHSQ(32) = 218.052 [.000]	156.2712[.000]
1	570.8462	550.8462	528.2194	CHSQ(48) = 408.614 [.000]	292.8402[.000]
0	41.2085	37.2085	32.6832	CHSQ(64) = 1338. 50[.000]	959.2544[.000]

AIC=Akaike Information Criterion
SBC=Schwarz Bayesian Criterion

Narayan (2005) pointed out that the critical bounds developed by Pesaran *et al*. (2001) are not suitable for a small sample. Our sample consists of T =76; we use critical bounds developed by Narayan (2005). The results of the *ARDL* bounds testing approach to cointegration are reported in Table 3. Our computed F-statistic exceeds upper critical bound at 5% significance level once *TSE* price index is used as predicted variable. This confirms the presence of cointegration between the variables over the period of 1993Q1–2011Q4. This entails that real *GDP*, measure of banking sector development; trade openness and *TSE* price index are cointegrated for a long-run relationship in the case of Iran.

Table 3*:* ARDL Bound Test to Long-run Cointegration

Test Statistics	CalculatedValue	Lag - order	Significance level	Bound Calculated Value	
				I(0)	*I(1)*
F-Statistics	6.0655	4	1%	4.385	5.615
			5%	3.219	4.378
			10%	2.711	3.823

The results of the impacts of real *GDP*, a measure of banking sector development and measure of trade openness on *TSE* price index are presented in Table 4. The findings indicate that real *GDP* has a positive impact on *TSE* price index. This implies that 4.2796% increase in *TSE* price index is linked with 1% rise in real *GDP*. The effect of measure of banking sector development on *TSE* price index is positive and significant at 10% level. A 1 percent rise in a measure of banking sector development increases *TSE* price index by 0.5416% keeping other things constant. The effect of measure of trade openness on *TSE* price index is negative and significant at 1% level. Keeping other things constant, a 1% increase on measure of trade openness lowers *TSE* price index by 2.7654%.

Table 4: Estimated long-run Coefficients the *ARDL* Approach; Dependent variable is *LnSI*

Regressors	Coefficient	Standard Error	T-Ratio [Prob]
LnGDP	4.2796	0.068906	2.3000[.025]
LnOT	-2.7654	0.031574	-3.2217[.002]
LnFD	0.54163	0.051893	1.8805[.065]
Intercept	-50.3915	0.13137	-2.0211[.048]
R-Squared	0.9984	R-Bar-Squared	0.99964
DW-statistic	1.8148	F-stat. F(13,46)	12661.5[.000]

After finding long run effect of real *GDP*, trade openness and banking sector development, on *TSE* price index, next step is to investigate their short run dynamics. For this purpose, we have applied error correction model (*ECM*). The results are reported in Table 5.

Table 5: Error Correction Representation for *ARDL* Model, Dependent variable is Dependent variable is □ *LnSI* - Preferred Specification

Regressors	Coefficient	Standard Error	T-Ratio [Prob]
□ LnGDP	0.38881	0.02495	2.2861[.026]
□ LnFD	0.049208	0.03202	1.7110[.092]
□ LnOT	-0.25124	0.01143	-3.1268 [.003]
Intercept	-4.5781	0.20696	-2.0305 [.047]
ECT(-1)	-0.090851	0.01729	-5.5774[.000]
R-Squared	0.8431	R-Bar-Squared	0.8257
DW-statistic	1.8148	F-stat. F(7,63)	48.385[.000]

The impact of real *GDP* and Banking sector development on *TSE* price index is positive and it is statistically significant. A 1% rise in real *GDP* causes 0.3888% rise in *TSE* price index. In addition, a 1% rise in banking sector development causes 0.0492% rise in *TSE* price index. The trade openness declines *TSE* price index and statistically it is significant. The estimate of *ECT* (-1) is negative and significant at 1% level corroborating our proven long run association between real *GDP*, Banking sector development, trade openness and *TSE* price index. The coefficient of the error correction term (*ECT*) is equal to -0.0908 and it is statistically significant. According to this estimation, the speed of adjustment is very slow. It is an indication of very slow and significant adjustment process for Iranian economy in any shock to *TSE* price index model. In addition, the *ECM* can explain 84% of fluctuation of *TSE* price index.

The results of a few diagnostic tests indicate that there is no error autocorrelation and conditional heteroskedasticity, and that the errors are normally distributed. This evidence indicates that the relationship between variables is verified (see Table 6). **Table 6**: Diagnostic Tests

Test Statistics	*LM Version*	*F Version*
A: Serial correlation	CHSQ(4) = 15.9709[.223]	F(4,59) = 4.2808[.234]
B: Functional form	CHSQ(1) = 0.95957[.327]	F(1,62) = 0.84941[.360]
C: Normality	CHSQ(2) = 3.2888[.193]	Not applicable
D: Heteroscedasticity	CHSQ(1) = .23524[.628]	F(1,69) = .22937[.634]

A: Lagrange multiplier test of residual serial correlation.

B: Ramsey's RESET test using the square of the fitted values. C:

Based on a test of skewness and kurtosis of residuals.

D: Based on the regression of squared residuals on squared fitted values.

Figure 1: Plot of Cumulative sum of Recursive Residuals

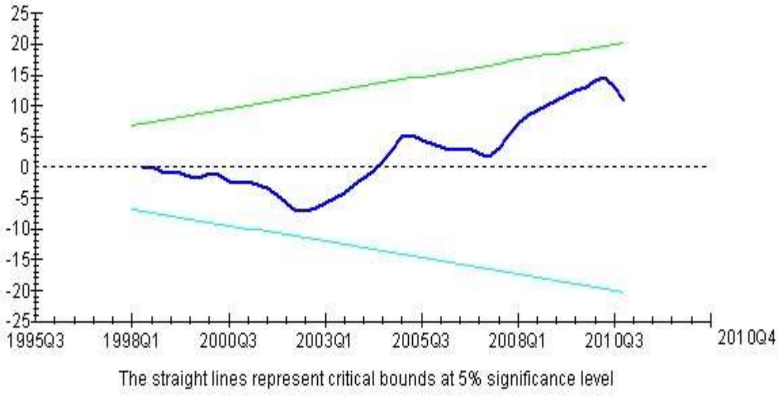
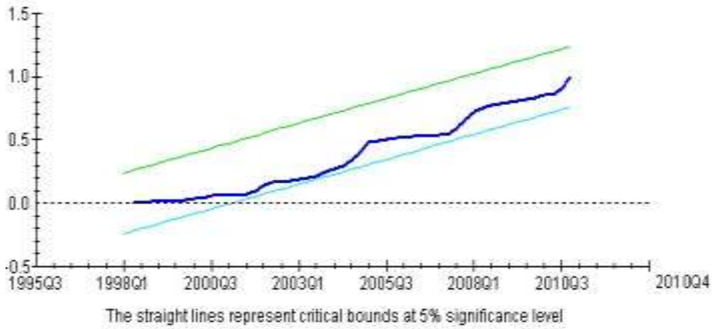


Figure 2: Plot of Cumulative sum of Squares of Recursive Residuals



5. Conclusion

This paper investigates the relationship between trade openness with *TSE* price index in Iran for 1993Q1–2011Q4 period. The bounds *F* test for cointegration test yields evidence of a long-run relationship between *TSE* price index, trade openness, real *GDP*, and banking sector developments. The results show that trade openness has a negative effect on stock prices at 1% significance level. In addition, the coefficient of banking sector developments and real *GDP* variables are positive at 5% significance level which shows that an increase in two variables results in an increase in stock prices.

This paper also explores the short-run relationship between the variables by using error correction model. The short-run results expose that trade openness has a negative impact on stock prices at 1% level of significance. The relationship between real *GDP* and stock prices is positive and statistically, it is significant at 5%. Banking sector developments is positively linked with stock prices at 10% significance level. The coefficients of estimated *ECTs* are also negative and statistically significant at 1% confidence level. These values indicate that any deviation from the long-run equilibrium between variables is corrected for each period to return the long-run equilibrium level. In addition, the coefficients of estimated *ECTs* showed that speed of adjustment is slow, and the *ECM* only can explain 84 per cent of fluctuation of *TSE* price index.

The results also don't support evidence of Basu and Morey (2005). It means that trade openness decrease stock prices in Iran. In addition, banking sector developments increase *TSE* price index. According to the Basu and Morey (2005) financial opening alone without trade, opening will not lead to efficiency in the stock prices. However, our findings don't confirm this evidence. But, it is necessary to consider that social, economic and political problems of Iran can effect on these results. First, Iran is one of developing countries and its stock market has not improved. Second, foreign trade in Iran suffered from economic instability, the trade sanctions, the freezing of Iranian assets that followed and economic isolation from the west.

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