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EDITORIAL NOTE

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The second paper, titled **Dynamic Causal Interactions of Money, Prices, Interest Rate and Output in Pakistan** by Faiz Bilquess, Tahir Mukhtar and Sidra Sohail investigates the dynamic interactions among macroeconomic variables such as money supply, prices, interest rate, exchange rate and output level, using the quarterly data for Pakistan over the period 1972Q1 to 2009Q4. For the empirical analysis the Johansen multivariate cointegration technique, Granger causality test and variance decompositions are employed. The results from the cointegration test indicate that there exists a stable long run equilibrium relationship among the macroeconomic variables of the study. The outcome of causality tests tends to support the non neutrality of money view of the Keynesians and the Monetarists at least in the short run.

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Twin Deficits and Saving-Investment Nexus in Pakistan: Evidence from Feldstein-Horioka Puzzle

Muhammad Arshad Khan[1] and Sumaira Saeed[2]

This paper tries to empirically examine the validity of Feldstein-Horioka puzzle in the presence of twin deficits for Pakistan over the period 1976 to 2010 using autoregressive distributed lag (ARDL)-bounds testing approach to cointegration. Our finding reveals a positive association between the current account balance and budget deficit in short-as well as in long-run. Toda-Yamamoto (1995) causality test also supports the existence of twin deficits. We also find a negative relationship between current account balance and investment in short-as well as in long-run. The long-run coefficient of investment is -0.75 which implies a high degree of capital mobility by rejecting the Feldstein-Horioka puzzle in case of Pakistan. However, the short-run coefficient of investment is negative but less than one (i.e. -0.46) which provides support for the moderate validity of Feldstein-Horioka puzzle but only in the short-run.

Introduction

Feldstein and Horioka (1980) find that changes in domestic investment are very sensitive to changes in domestic savings for OECD countries, which has launched a debate regarding the degree of financial integration and financial openness within the Industrialized Countries (Ho, 2003). A positive link between domestic investment relative of gross nation product (GNP) and domestic savings relative to GNP is interpreted as measure of international capital mobility. Dooley *et al*. (1987) and Feldstein and Bachetta (1989) and generally confirms the Feldstein-Horioka's findings of close associations between domestic

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savings and investment for industrial as well as developing countries. The results of Feldstein-Horioka's study showed that the saving retention coefficient, which measure the level of capital mobility in 21 member states of the Organization for Economic Cooperation and Development (OECD) was between 0.87 and 0.91, which proves relatively low capital mobility in these countries (Misztal, 2011). These results imply that 90 percent of savings in OECD member countries are used to finance domestic investment. Some evidence also supports a strong correlation between domestic savings and investment across OECD countries over time, indicating the implications of capital immobility which coins the Feldstein-Horioka puzzle (Ho, 2003). The of Feldstein-Horioka's (1980) research findings, which are contrary to economic theory, have to be referred to as "the mother of all puzzles" (Obsfeld and Rogoff, 2000, p. 9).

The other puzzle is associated with the positive associations between current account deficit and budget deficit ─ often known as twin deficits hypothesis.[3] The concept of twin deficit is not a new one but it can be traced back to the early writings of Mundell-Fleming and Keynesian theories. According to the Mundell-Fleming framework, an increase in budget deficits will cause an increase in domestic interest rate above the world rate that, in turn, will cause capital inflows and exchange rate appreciation, leading to increase in the current account deficits (Marinheiro, 2008 and Burcu, 2011). On the other hand, according to Keynesian absorption theory, an increase in budget deficits would induces domestic absorption which would results an increase in import expenditures and decrease in export revenues through the loss of international competitiveness in the international market, causing deterioration of the current account deficits (Marinheiro, 2008). Thus, there exists a positive relationship between budget deficits and current account deficits─ budget deficit causes current account deficit. Twin deficit issue is also associated to the degree of capital mobility and the Feldstein-Horioka puzzle. If domestic savings and investments are not highly correlated, then the budget deficits and the current account deficit

[3] The term "twin deficit" was used to describe the concurrent US budget and trade deficits for the period 1981-1998- both of which having reached a record levels.

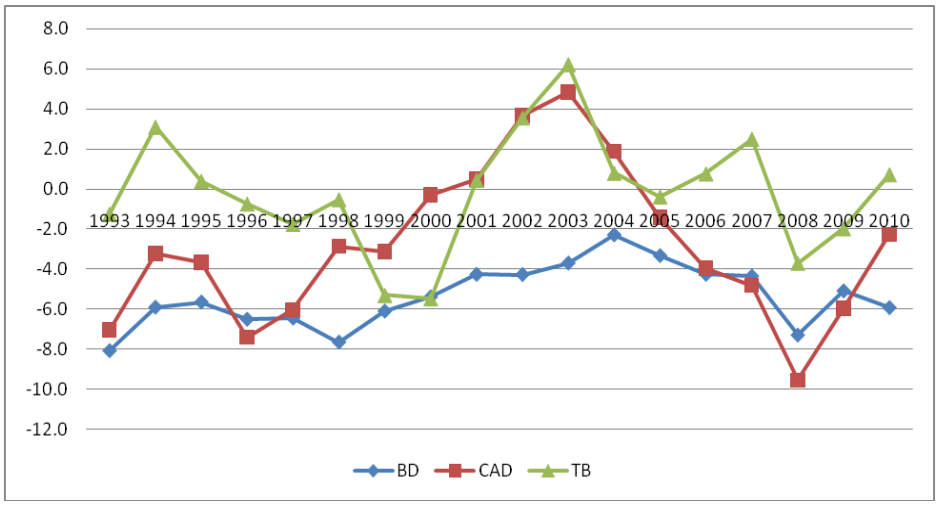
can be expected to move together in the same direction. In contrast, the Ricardian Equivalence Hypothesis (REH) due to Barro (1974) claims the absence of any relationship between budget deficit and current account deficit. In other words, REH predicts that current account deficit, interest rates and investment will not be affected by the changes in budget deficit (Burcu, 2011). It implies that tax increases would reduce the budget deficit but would not alter the trade deficit. In other words, the government tax-finance deficit policy does not affect private spending and national savings (Marinheiro, 2008).

However, if the REH does not hold, then there is possibility for the existence of causal relationship between the current account deficit and the budget deficit. If the causality is running from current account to the budget deficit then deterioration in the current account balance results in to diminishing economic growth which, leads to deterioration in the budget balance. This implies that the government targeting the current account and uses budget deficit as an instrument to eliminate external imbalances (Marinheiro, 2008). When the causality is running from budget deficit to current account, then exchange rate, interest rate, income, savings, investment, exports and imports are dependent on the fiscal and monetary policies. Under this situation, attention should be focused to the exchange rate policy, monetary policy and export promotion policies (Marinheiro, 2008 and Hallwood and MacDonald, 2003). If the REH does not hold then there is also possibility that an increase in public borrowing is expected to decrease national savings. For a given amount of investment, this decrease in national saving leads to an increase in the current account deficit. It may be mentioned here that perfect capital mobility results into twin deficits in a non-Ricardian world. In other words, if the Feldstein-Horoika puzzle holds there is no possibility of twin deficits to coexist.

For the empirical analysis of Feldstein-Horoika puzzle and twin deficits, Pakistan is more relevant country because Pakistan has faces a persistent high degree of budget and current account deficits since 1980. During the 1980s, average budget deficit was 7.1 percent of GDP, and reached its peak level of 8.5 percent 1993. After 1993 it was come down and reached its minimum level of 2.3 percent in 2004. It rises again and was reached to 6.3 percent in 2010. The current account balance of Pakistan

remained in deficit during last three decades except for few years. The current account deficit which was 5 percent of GDP during the 1980s, reached to its peak level of 7.4 percent of GDP in 1996. It rises again in 2005 and was reached to 2.3 percent in 2010 before it touched a record level of 8 percent of GDP in 2008 (see Figure 1). Furthermore, in the late 1980s Pakistan was faced with high macroeconomic imbalances as a result of growing inefficiency and losses in the public sector. To restore business confidence and reduce inefficiency and losses, the government implemented a wide range of structural and institutional reforms in the early 1990s (Khan and Qayyum, 2006).[4] As a result, the public spending has increased more than public revenues, leading to deterioration of budget deficits and balance of payment deficits. The increase in budget deficit has caused a sharp increase in internal and external borrowings. Moreover, increase in domestic interest rate puts downward pressure on domestic investment and economic growth. This situation has increased budget deficit even more by reducing tax revenues. Similarly, a persistent depreciation in Pak-rupee exchange rate has increased import bill as well as debt burden.

Figure 1: Trends of Current Account, Budget Deficits and Trade Balance (as % of GDP)



[4] Detailed review of reforms can be seen in Khan and Khan (2007).

The correlation between current account deficit and budget deficit has important implications for developing countries like Pakistan. For example, a persistent large budget deficits causes indebtedness by borrowings domestically and internationally and increases burden on future generations (Hakro, 2009). Thus rising trade deficits is escalating budget deficits and current account balance cannot be remedied unless government would put in place appropriate policies which would address fiscal as well as balance of payments deficits.

In the context of Pakistan a number of studies have been carried out to examine the link between current account and budget deficits. For example, Burney and Yasmeen (1989), Zaidi (1995) and Burney and Akhtar (1992) have used ordinary least squares (OLS) technique to explore the link between current account deficit and budget deficit with other macroeconomic variables. The other studies, inter alia, by Kazimi (1992), Aqeel and Nishat (2000), Mukhtar *et al*. (2007), Hakro (2009) and Javid *et al*. (2011) have reported evidence supported to bi-directional causality between current account deficit and budget deficit. These studies have made significant contributions in determining causal link between budget deficit and current account deficit. However, to the best of our knowledge no attempt has so far made to test the twin deficit and Feldstein-Horioka hypothesis in the case of Pakistan. The present study fills up this gap.

The main purpose of this study is to examine the validity of twin deficit and Feldstein-Horioka puzzle for Pakistan using annual time series data for the period 1976 to 2010.

The rest of the paper is organized as follows. Section 2 reviews the existing literature on Feldstein-Horioka and twin deficit hypothesis. Theoretical and analytical framework is presented in section 3. Section 4 gives data description and econometric methodology. Section 5 discusses the estimation results while section 6 is devoted to conclusion.

2. Literature Review

There are two main strands of literature regarding the validity of Feldstein-Horioka's puzzle. These are substantial disagreement about the

implications of Feldstein-Horioka's (1980) hypothesis. The economists argued that the Feldstein-Horioka puzzle can be solved using intra-national regional data, which indicate that the saving retention coefficient is not related to the capital mobility across international borders, but within intra-national regions (Ho, 2003). The studies based on cross-sectional data during the 1980s find strong correlation between savings and investment which implies greater capital immobility at international level. Some studies (for example, Feldstein and Horioka (1983), Penati and Dooley (1984), Feldstein (1983), Obstfeld (1986), Haque and Montiel (1991), Dooley *et al*. (1987), Feldstein and Bacchetta (1989), Frankel (1991), Tesar (1991) and Mussa and Goldstein (1993) among others) suggest a strong correlation between savings and investment. Feldstein and Horioka (1980) find that sustained increase in domestic saving rate induce approximately an equal increase in domestic rate of investment. These estimates for the post-OPEC crisis period (1974-1979) implied that each extra dollar of domestic saving increases domestic investment by approximately 85 cents in a sample of 17 OECD countries. Vamvakidis and Wacziarg (1998) using the panel data for the period 1970-1993 and found a very high correlation between domestic saving and domestic investment for OECD economies. Coakley *et al*. (1998) tested the Feldstein–Horioka's view that a high saving-investment association across OECD countries. Their results show a strong cross-section association between saving and investment rates in OECD countries which implied low capital mobility.

The studies based on time series data (for example, Miller (1988), Ho (1999, 2000), Schneider (1999), Erden (2005), Kohler (2005), Grier *et al*, (2007), Mastroyiannis (2007), Murthy (2007) and Ozmen (2007), among others) reveals mixed results. Levy (2003, 2004) concluded that the investment-savings cointegrated in levels as well as into rates. The results indicate that during the period 1947-1987 the USA economy did not have much capital mobility. Khedhiri and Hebiri (2005) found the evidence of significant correlation between savings and investment, and thus no solution to the Feldstein-Horioka puzzle, despite the implications of international capital mobility in six countries such as Algeria, Egypt, Morocco, Saudi Arabia, Syria and Tunisia. Rao *et al*. (2008) found the Feldstein-Horioka puzzle exists in a weaker form with a reduced saving retention coefficient for 12 OECD countries.

After the 1980s, correlation between savings and investment has not supported the Feldstein-Horioka hypothesis because other macroeconomic variables have also contributed to the prevalence of relationship between saving and investment (Murphy, 1984 and Shahbaz *et al*., 2010). These factors include current account solvency constraints, structure of financial sector, country size, exchange rate regime, endogenous shocks, etc. For example, "Coakley *et al*. (1996) argued that the saving retention coefficient is irrelevant to the analysis of the hypothesis of international capital mobility, but it instead useful as proxy for current account solvency. This is because, in the presence of a current account solvency constraint, the current account balance should be stationary and thus investment and saving rates are cointegrated with the unitary coefficient" (Ho, 2003, p. 388).. Similarly, Sachsida and Caetano (1999) corroborated the evidence that in the real side of the economy, Feldstein-Horioka hypothesis does not reflect capital mobility. They argued that coefficient of investment does not mean capital mobility, but just indicate substitutability relations between external and domestic savings. Sachsida and Caetano (2000) argued that the saving retention coefficient does not even reflect capital mobility, but explains variability between external and domestic savings. Murphy (1984) argued that the estimated saving retention coefficient is sensitive to country size. Dooley *et al*. (1987) predicts that the ratio of current account balance to GNP should be smaller for large countries. Ho (2003) examined the threshold effect of country size on the magnitude of saving retention coefficient for a panel of 23 OECD countries over the period 1961-1997. The evidence of this study indicates that saving retention coefficient increases as the relative GNP share becomes larger, which substantially supports the country-size argument. Fidrmuc (2003) who examined the twin deficits and the Feldstein–Hokiora hypothesis in a wide range of industrial countries including three new EU member states as well as selected emerging economies and transition economies over the period 1970-2001. The results of his study suggest a positive long-run relationship between the twin deficits. Furthermore, his findings also suggest that investment in some EU countries is financed by the international financial markets, implying that the Feldstein–Horioka hypothesis was less significant for EU countries. The Feldstein-Horioka puzzle is revisited by Caporale *et al.* (2003, 2005) by employing a variety of cointegration tests and concluded that despite

evidence supporting the Feldstein-Horioka puzzle, there is a considerable heterogeneity in the savings-investment association. Murthy (2007) examined the validity of Feldstein-Horioka for a panel of 14 Latin American and 5 Caribbean countries over the period 1960-2002. He finds that saving retention coefficient is relatively small indicating the rejection of Feldstein-Horioka hypothesis. He concluded that small saving retention coefficient therefore implying moderate capital mobility. Di Iorio and Fachin (2007) used bootstrap methods to study the Feldstein-Horioka hypothesis in 12 EU members between 1906 and 2002 and find that the saving retention coefficient for each of the countries was between 0.59 and 1.03. Marinheiro (2008) tested the validity of Ricardian equivalence and Feldstein-Horioka hypothesis for Egypt over the period 1971-2004. The results of this study do not support the Ricardian equivalence hypothesis. However, the study presents evidence in favour of high degree of capital mobility implying the rejection of Feldstein-Horioka hypothesis. Onafowara *et al.* (2011) analyzed the relationship between saving and investment in eight highly developed EU countries using ARDL and ECM techniques. The study finds cointegration between saving and investment in six of those countries. The results also showed that saving appears as significant determinant of investment for Netherlands and Great Britain, a reverse causality in Denmark, Germany and Luxembourg and bi-directional causality in Belgium and no causality between saving and investment in France and Italy. Shahbaz *et al.* (2010) have tested saving investment correlation for Pakistan over the period 1976-2006 and find weak correlation between saving-investment. Furthermore, the study suggests that in the presence of inadequate capital mobility, domestic investors have financed their investment projects from international markets. However, the study recommended that devaluation and inflation have stimulated investment activities and significantly contributed in closing the gap between domestic savings and investment.

The debate whether saving-investment co movements are indication of capital mobility is still inconclusive. The results of Feldstein-Horioka hypothesis based on cross-section data shows a strong association between saving and investment rates in OECD countries. The response of economists to the Feldstein-Horioka's view regarding a high correlation between saving-investment implicitly meaning low capital

mobility across OECD countries is much discussed. The long-run investment-saving correlation follows directly and this does not depend on the degree of international capital mobility, unless the budget constraint is removed. The studies based on time series data reveals that investment and saving are high cointegrated. However, debate over the point whether saving-investment co-movement as indication of the degree of capital mobility is still unresolved.

Regarding the external deficit and budget deficit nexus, many studies has been carried out. These studies can be categorized into three broad groups. The first group of studies (for example, Darrat (1988), Abell (1990a, b), Bachmen (1992), Rosenweigh and Tallman (1993), Cavallo (2005) and Erceg *et al*. (2005), among other) supports the twin deficit hypothesis. The second group of studies (Enders and Lee (1990), Miller and Russek (1989) and Kim (1995), among others) concludes against twin deficit hypothesis. The third group of studies, *inter alia*, by Anoruo and Ramchander (1998), Khalid and Teo (1999) and Alkswani (2000) supports the reverse causality running from current account to budget deficit. Fountas and Tsoukis (2000) examined the interactions between the current account, budget balances and the real interest rate as proxy of financial openness and found that in short run there is some evidence in favor of the twin deficits and current account targeting hypotheses.

In the context of Asian countries Baharumshah and Lau (2009) find evidence in favour of twin deficit hypothesis for Indonesia, Korea, the Philippines and Thailand, and evidence again the twin deficit hypothesis for Singapore and Japan. Makin and Narayan (2008) argued that current account deficit in the USA is strongly coincident with saving rate in East Asia, especially in the post-1997 period. Other studies, for example, Lau and Baharumshah (2006), Baharamshah *et al*.(2006) and Baharamshah and Lau (2007), among others failed to establish causality consensus results. Lau *et al*. (2010) revisited the twin deficit hypothesis in Asian 1997-cirisis affected countries. Empirical results suggest that causality runs from budget deficit to current account deficit for Malaysia, the Philippines and Thailand. For Indonesia and Korea the causality runs in the opposite direction, while a bi-directional causality exists for the Philippines. The study suggests that managing these deficits are indeed

important policy options in promoting macroeconomic stability and sustainability in the region.

In Pakistan's case regarding the validity of twin deficit hypothesis debate is incomplete. Studies *inter alia*, by Shabbier and Ahmed (1994) find that budget deficit have strong positive and significant direct effect on inflation and indirect effect on inflation via money supply. Naqvi and Khan.(1989) find the deficits exerts only indirect effect induced by changes in money supply. Siddiqui (1990) find two-way causality between money supply and inflation, which may be due to common link with deficits (Hakro, 2009). Other studies such as Gupta (1992), Jones and Khiliji (1988), Biliquees (1985) and Kemal (1991) showing mixed results. Hakro (2009) tested the causal link between twin deficit and other macroeconomic variables for Pakistan using quarterly data over the period 1948Q1-2005Q4. His results confirm the causality link of budget deficit running from budget deficit to prices to interest rate to capital flow to exchange rate and to trade deficits. The results also suggest the reduction in the budget deficit might help to control the level of prices. Furthermore, minimum government intervention and a balanced budget could restore the imbalances and a measure of confidence in international trade is also recommended by this study. Javid *et al.* (2011) investigated the impact of government budget deficit shocks on the current account balance and other macroeconomic variables for Pakistan over the period 1960-2009. Based on structural vector autoregressive method the results suggested that expansionary fiscal policy shocks improve the current account balance and depreciates the exchange rate. Furthermore, a rise in domestic saving and a fall in investment contributed to the improvement in the current account. The study also suggests that co-movements in fiscal deficits and current account deficits supported the Ricardian's view.

3. Twin Deficits and the Feldstein-Horioka Puzzle:A Theoretical Consideration

It is well documented in the literature that under the assumption of perfect capital mobility, domestic saving is not necessarily related to domestic investment. However, much of the empirical evidence is based on the cross-section data of 16 OECD countries for the period 1960–

1974 provides supportive evidence that domestic investment and national saving have strong correlation. This empirical finding is known as the Feldstein-Horioka (1980) puzzle. According to this puzzle, the relationship between national savings and domestic investment can be treated as a measure of international capital mobility. Under perfect capital mobility, investment is not controlled by domestic saving but only by the accessibility of funds in the international fully integrated capital market. Saving in each country reacts to the worldwide opportunities for investment, while investment in a country can be financed by the international pool of funds. The important factor of concern for investors should be the rate of return. In closed economies, domestic savings are used to finance investment; however, in open economies a part of investment may be financed by foreign savings. Therefore, savings (S) and investment (I) could move independently of each other. Alternatively, the high investment-saving (*I-S)* correlation suggests that capital might not be fully mobile across international borders. (Levy 2004) [5]

3.1 An Analytical Framework for Twin Deficits and Feldstein-Horioka Puzzle

Feldstein and Horioka (1980) proposed an assessment of the degree of capital mobility by measuring the correlation between investments and savings. They estimated cross section regressions of the form:

$$\left(\frac{I}{Y}\right)_t = \alpha + \beta\left(\frac{S}{Y}\right)_t + \varepsilon_t \qquad (1)$$

Where $\frac{I}{Y}$ is the ratio of gross domestic investment to gross national product (GNP), $\frac{S}{Y}$ is the ratio of national saving to GNP and $\varepsilon_t \sim iid.N(0, \sigma^2)$. The value of β should be close to zero if the international capital mobility is perfect. In contrast, the size of β should

[5] See Levy (2004) for further details.

be large if capital is immobile. Such relationships between share of investment to GDP and share of saving to GDP, or some of their variants such as, first differences of these have been essential for many countries. Equation (1) is called Feldstein-Horioka regressions where β is the Feldstein-Horioka coefficient representing the saving-investment association.

3.2 Importance of Feldstein-Horioka Puzzle

The degree of capital mobility plays an important role in financial decisions. Its importance can be judged in following perspectives:

1. The effect of fiscal policy crucially depends on the extent of capital mobility.
2. The cost of adjustment to external shocks gets reduced via the access of an economy to capital markets.
3. Capital mobility determines the rate at which incomes converge to long-run equilibrium path.
4. Perfect capital mobility is often assumed to hold in macroeconomic models and exchange rate models (Levy, 2004).

The twin deficit hypothesis has a clear link with the Feldstein-Horioka puzzle since the later is concerned with the source of financing the external deficit. The causality between external balance and fiscal balance stresses the role of private investment. Fidrmuc's (2003) regression model encompasses both the twin deficit hypothesis and the Feldstein-Horioka puzzle. The model makes no difference between net exports and current account balance.

The starting point for the explanation of twin deficits and domestic savings-investment correlation is provided by the national income identity (Miller, 1988). Following Fidrmuc (2003) and Bagnai (2006) the public sector may meet its financial needs through domestic and international financial markets. Rearranging the national income identity, we obtain the following relation:

$$X_t - M_t = Y_t - (C_t + G_t) - I_t = S_t - I_t \qquad (2)$$

This identity implies that the trade balance which is defined as the difference between exports (X) and imports (M) must be equal the difference between national savings, defined as output (Y) less total consumption ($C + G$) and investment (I). It provides a link between the external balance and saving-investment decisions. Hence, an increase in domestic investment has a negative impact on the external trade balance. On the other hand, policies that reduce (public or private) consumption are expected to have a positive impact on the external balance via increased national savings. National savings can be subdivided into private saving (S^p) and public saving (S^g). The latter corresponds to the budget balance and equal to the difference between tax revenues (T) and government expenditures (G). Private savings are defined as the disposable income minus private consumption (C). Thus equation (2) can be rewritten as:

$$X_t - M_t = (Y_t - T_t - C_t) + (T_t - G_t) - I_t = S^p + S^g - I_t \qquad (3)$$

Equation (3) motivates to test for the long-run relationship between the current account balance (CAB) which is proxied by ($X - M$), the budget deficit (BD) which is proxied by fiscal balance ($T - G$) and total investment (INV). Following Fidrmuc (2003) the variables expressed as share of GDP as:

$$CABY_t = \beta_1 + \beta_2 BDY_t - \beta_3 INVY_t + \beta_4 DUM_{01} + u_t \qquad (4)$$

Where $CABY$, BDY and $INVY$ are respectively current account balance as percent of GDP, budget deficit as percentage of GDP and investment as percentage of GDP. To account for changes in current account balance from deficit to surplus after 9/11 shocks, we introduced a dummy variable DUM_{01} which takes the value one for 2001-2003 and zero otherwise.[6] Equation (4) provides a useful link between

[6] To account the impact of exchange rate regime we introduced two dummy variables, 1982-2000 for managed floating exchange rate regime and from 2000 to onward for flexible exchange rate regime. However, our preliminary indicates insignificance of these variables. Therefore, we left out these dummies from the analysis.

$CABY$, BDY and $INVY$. We expect $\beta_2 > 0$ and $\beta_3 < 0$. It implies that budget deficit and investment worsens the current account balance. If an economy is perfectly integrated with the world market, then the coefficients of both variables are equal to unity. In this case, the budgetary as well as investment expenditures are financed by the world financial market. However, according to Feldstein and Horioka (1980) a large portion of domestic investment is still financed from domestic savings. If the Feldstein-Horioka puzzle is there, the coefficient β_3 is significantly lower than unity. Likewise, a negative β_2 coefficient leads to the rejection of the twin deficit hypothesis.[7] Empirical research shows that existence of long-run relationship between budget deficit and current account deficit, and investments in developed economies is considered as one of the six main empirical puzzles in the modern macroeconomics (Obsfeld and Rogoff, 2000). In the present era of globalization and internationalization flow of goods and services and factors of production and free mobility of capital enables the effective allocation of capital regardless of existing state boarders. Therefore, the size of the investment in any open economy cannot be limited by by the size of domestic savings (Francesca and Stefano, 2010).

4. Methodology and Data Description

We used the autoregressive distributed lag (ARDL) modeling approach to cointegration to estimate twin deficits and Feldstein-Horioka hypothesis. Although, there are some other approaches to estimating cointegration relationships, such as Engle and Granger (1987) and Johansen (1991). However, these approaches requires that underlying series be integrated in the same order. The ARDL avoids the pre-testing of unit roots and the classifications of variables into I (1) or I (0). Empirically, we started with a general model that included 3 lags of all first differenced variables and one lag of levels variables and it was estimated by OLS method. Then we eliminated the statistically insignificant differenced variables following general-to-specific (GETS)

[7] These specifications are used for Pakistani data and resemble those adopted by Marinheiro (2008) for analysis of Egypt data. Only the proxy variable of wealth is excluded from model.

methodology. The most parsimonious conditional unrestricted error-correction model (UECM) for twin deficits and Feldstein-Horioka is given by[8]

$$\Delta CABY_t = c + \pi_1 \Delta BDY_t + \pi_2 \Delta INVY_t + \pi_3 \Delta INVY_{t-1} + \delta_1 CABY_{t-1} + \delta_2 BDY_{t-1} + \delta_3 INVY_{t-1} + DUM_{01} + \eta_t \quad (5)$$

The coefficients $\delta's$ are the long-run parameters and c is the drift term and dummy variable DUM_{01} that equals one from 2001-2003 and zero otherwise. This capture the 9/11 shocks on current account balance. The current and lagged values of differenced variables are used to modell the short-run dynamics. The bounds test for the existence of level relationships between $CABY_t, BDY_t$ and $INVY_t$ have the following hypotheses:

$$H_0 : \delta_1 = \delta_2 = \delta_3 = 0,$$

and alternative hypotheses are correspondingly given by:

$$H_1 : \delta_1 \neq \delta_2 \neq \delta_3 \neq 0$$

The ARDL framework allows for testing the presence of long-run relationships among the variables under investigation by employing F-statistic (Felipe *et al*., 2011). The F-statistic has a non-standard distribution, which depends on the unit root properties of the data that is whether variables included in the UECM are I (0) or I (1), and the number of independent variables. Pesaran *et al.* (2001) provide the critical values of the upper bound and lower bound for the F-test to be used to determine cointegration under various assumptions. If the calculated F-stat lies above the upper bound values, the hypothesis of no cointegration can be rejected and vice versa.

[8] We have tried various specifications by using different lag length and intercept and trend terms. We specify most parsimonious specifications throughout the study. For the selection of parsimonious model we use PcGets software.

If there is an evidence of cointegration between $CABY_t, BDY_t$ and $INVY_t$ then one can precede further using autoregressive distributed lag (ARDL) specification of equation (4) on the basis of lag selection criteria such as the Akiake Information Criterion (AIC) and Schwarz Bayesian Criterion (SBC).[9] The most parsimonious ARDL representation based on general-to-specific methodology is given by :

$$CABY_t = \gamma_1 CADY_{t-1} + \gamma_2 BDY_t + \gamma_3 INVY_t + DUM_{01} + \omega_t \quad (6)$$

The long-run coefficients can be obtained as:

$$\beta_k = \left(\gamma_n \Big/ 1 - \gamma_1 \right)$$

With $k = 1,\ldots\ldots, 3$ and $n = 2,\ldots\ldots, 3$.

Finally the short-run dynamic coefficients for the optimal ARDL can be obtained by estimating the following error-correction model.

$$\Delta CABY_t = \phi_0 + \sum_{i=1}^{k} \phi_{1i} \Delta CABY_{t-i} + \sum_{i=0}^{k} \phi_{2i} \Delta BDY_{t-i} + \sum_{i=0}^{k} \phi_{3i} \Delta INVY_{t-i} + \lambda ec_{t-1} + \rho_t \quad (7)$$

Where ec_{t-1} is error correction term and λ is the speed of adjustment towards the long-run equilibrium after a shock.

The study is based on the annual data covering the period from 1976-2010. Current account balance is calculated by taking the difference between exports and imports. Fiscal balance is calculated as total revenues minus total expenditures. Gross fixed capital formation is used as proxy of investment. Data on these variables are retrieved from the World Bank's *World Development Indicators*. Data on GDP is taken

[9] However, Pesaran and Shin (1999, p. 3) noted that ARDL based on SBC performed slightly better than ARDL based on the AIC (also see Khan and Qayyum, 2009).

from International Monetary Fund's *International Financial statistics (IFS)* CD-ROM 2011. All variables are expressed as percentage of GDP in order to account for their share in the economy.

5. Empirical Results and Their interpretation

This study is based on the annual data over the period 1976-2010. The analyzed period is of particular interest because there were significant changes in the Pakistan's exchange rate regime, financial sector reforms are also undertaken during this period and the most important event (i.e. 9/11)also happened during this time.

Before estimating the twin deficits and Feldstein-Horioka hypothesis, the stationarity properties of the data were checked using Augmented Dickey Fuller (ADF) test. The results of ADF test are reported in Table 1.

Table 1: Augmented Dickey Fuller Test

Variable	Levels	First Difference	Conclusion
$CABY_t$	-3.1043 (1)*	-4.5918 (1)*	I(0)
BDY_t	-2.0507 (1)	-4.2040 (1)*	I(1)
$INVY_t$	-3.1285 (0)*	-5.5006*	I(0)

Note: The 5 percent critical value of ADF is -2.9446. Figures in brackets indicate number of lags, while *indicate significant at the 1 percent level of significance.

It is evident from the ADF test that the current account balance as percentage of GDP ($CABY_t$) and ratio of investment as percentage of GDP ($INVY_t$) are stationary at their levels, whereas budget deficit as percentage of GDP (BDY_t) is non-stationary at its levels and stationary at its first difference. This means that $CABY_t$ and $INVY_t$ are integrated of order zero i.e.. I (0) and BDY_t is integrated of order one, i.e. I (1). These mixed results of ADF test justifies the use of autoregressive distributed lag (ARDL) method of cointegration.

5. 1 Cointegration Analysis: Bounds Test of Cointegration

We follow a three-step procedure of the ARDL method to estimate equation (4). In the first step, we have estimated equation (5) by OLS for examining the existence of long-run relationship between $CABY_t$, BDY_t and $INVY_t$. The optimal lag length for UECM is selected of order 1 based on SBC. The F-test is calculated by imposing zero restrictions on the lag level variables. The calculated $F_{CABY}(CABY_t \mid BDY_t, INVY_t, DUM_{01})$ was equal to 9.7316 is higher than the upper bound of critical values (i.e. 4.3495) at the 5 percent level of significance. Therefore, we reject the null hypothesis of no cointegration among $CABY_t$, BDY_t and $INVY_t$.[10]

Having found the long-run relationship between $CABY_t$, BDY_t and $INVY_t$, the second step is to estimate ARDL equation (6). The estimated results are presented by equation (8) (t-values in parenthesis):

$$CABY_t = 0.3851\ CABY_{t-1} + 0.2897\ BDY_t - 0.4585\ INVY_t + 4.0382\ DUM_{01}$$
$$(2.6005)^{*} \qquad (3.4964)^{*} \qquad (-1.8034)^{**} \qquad (3.1626)^{*}$$
$$\bar{R}^2 = 0.6012 \qquad DW = 2.03 \qquad D-h = -0.17 \qquad F = 17.59\,[0.000]^{*}$$

Bounds Test :

	$F-$calculated	$F-$Upper bounds (95%)	$F-$Lower bound (95%)
$F-$stat :	9.7361	3.0650	4.3495
$W-$stat :	29.2082	9.1950	13.0485

(8) [11]

Diagnistic Test :

$SC-\chi^2(1) = 0.0290\ [0.865]$ $\qquad FF-\chi^2(1) = 0.1251\ [0.724]$

$NO-\chi^2(2) = 4.5829\ [0.101$ $\qquad Het-\chi^2(1) = 0.0013\ [0.971]$

[10] Results of this step are available to the authors and can be obtained upon request.

[11] * and ** indicate significance at the 1 percent and 5 percent level of significance. The diagnostic test include serial correlation test (SC), function form (FF), normality (NO) and test for heteroscedasticity (Het), all are distributed to chi squares. Numbers in [.] indicates p-values.

It can be seen from the results reported by equation (8) that all the variables are significant. The coefficient of current account balance lagged by one year exerts positive and significant effect on current account balance. Budget deficit and investment exerts positive and negative effects on current account balance. However, the magnitude of these coefficients is less than that of expected values of twin deficit and Feldstein-Horioka hypothesis. The dummy variable to account for 9/11 effects on current account balance is positive and significantly affecting $CABY_t$.

The long-run coefficients can be obtained from the estimated equation (8). The long-run coefficient of budget deficit is 0.4711 (calculated as 0.2897/1-0.3851=0.4711) and the long-run investment coefficient is -0.7457 (calculated as -0.4585/1-0.3851=-0.7457). The corresponding normalized long-run parameter estimated from equation (8) is as follows:

$$\underset{\quad}{CADY_t} = \underset{(6.2537)^{*}}{0.4711\ BDY_t} - \underset{(-1.5233)^{***}}{0.7457\ INVY_t} + \underset{(3.5624)^{*}}{6.5675\ DUM_{01}} \qquad (9)$$

The result reported in equation (9) confirms a positive association between the budget deficit and current account deficit. The positive coefficient of budget deficit in relation with current account deficit leads to acceptance of the twin deficits hypothesis. Our results suggest that a rise in budget deficit by one percent of GDP worsens current account balance by 0.47 percent of GDP. Furthermore, result also reveals that investment exerts negative and significant impact on current account balance. An increase in investment by 1 percent of GDP improves current account balance by 0.75 percent. The dummy variable to account for 9/11 effects on the current account balance is positively and significantly affecting current account balance. All the variables appear in equation (10) have expected signs. These results support the validity of twin deficit hypothesis and rejection of Feldstein-Horioka puzzle for Pakistan. The coefficient of $INVY_t$ is less than unity which implies that domestic investment needs are to partially financed on the world financial markets. It also implies that Pakistan is not perfectly integrated

to world market and about ¾ of its investment is financed through external funds (foreign savings). This reason for this situation could be that the capital flows are limited during 1976-1990 because of heavy controls. However, Pakistan launched financial sector reforms since 1990 to dismantle controls on financial sector, trade and payments system. The results reported in equation (10) rejects the validity of Feldstein-Horioka hypothesis in the case of Pakistan.

Finally, to examine the short-run behaviour of current account balance, budget deficit and investment, we have estimated an error-correction model and equation (10) reports the result (t-values are in parenthesis):

$$\begin{aligned} \Delta CABY_t &= 0.2897\ \Delta BDY_t - 0.4587\ \Delta INVY_t + 4.0382\ \Delta DUM_{01} - 0.6149\ ec_{t-1} \\ &\quad (3.4964)^{*} \qquad (-1.8034)^{**} \qquad (3.1626)^{*} \qquad (-4.1520)^{*} \\ ec_t &= CABY_t - 0.47107\ ^{*} BDY_t + 0.74571\ ^{*} INVY_t - 6.5675\ DUM_{01} \\ \bar{R}^2 &= 0.43 \qquad F = 9.1610\ [0.000]^{*} \qquad DW = 2.03 \end{aligned} \tag{10}$$

The short-run parameters presented in equation (10) also in line with long-run parameters presented by equation (9) in terms signs and significance. However, the size of the parameters is small compared to the long-run estimates of the variables. The budget deficit and investment produces positive and negative influence on current account balance in the short-run. The coefficient of budget deficit is 0.29 which implies that an increase in budget deficit by one percent of GDP deteriorates current account balance by 0.29 percent of GDP in the short-run. The positive and significant association between budget deficit and current account deficit again confirms the validity of twin deficit in the short-run in the case of Pakistan. Similarly, investment exerts negative influence on current account balance in the short-run. The coefficient of investment is -0.46 which is less than unity implying that in the short-run about 0.46 percent of Pakistan's investment is financed by foreign savings. This result provides a weak support for the Feldstein-Horioka hypothesis in the case of Pakistan in the short-run. The positive coefficient of the dummy variable suggest that 9/11 event significantly deteriorates current account balance in Pakistan. The error-correction coefficient is negative and significant. The magnitude of error-correction

term is -0.61 which suggest that the deviation from the long-term $CABY_t$ path is corrected by 61 percent over the following year. To check the appropriateness of the results cumulative sum of recursive residuals (CUSUM) and cumulative sum of squares recursive residuals (CUSUMSQ) stability test is carried out. The plots of stability test suggest that the estimated parameters lie within the 5 percent critical bound, thus not violating the structural stability (see appendix).

On the whole, the above results existence of cointegratiopn which confirm the long-run relationship among current account balance, budget deficit and investment in the case of Pakistan. Our findings support the validity of twin deficits both in short-as well as kin long-run. Similarly, we got no support for the validity of Feldstein-Horioka hypothesis in the long run but weak support in the short-run. The validity of Feldstein-Horioka hypothesis implies that Pakistan's economy is well integrated into the world economy and government borrowings could be financed through international financial markets.

5.1 The Multivariate Granger Causality Test

To determine the direction of causality we employ multivariate Granger causality test. Since the ADF unit root test suggested that current account balance and budget deficit is integrated of order one (i.e. I (1)) and investment is integrated of order zero (i.e. I (0)). Therefore, we examine causality with Toda and Yamamoto (1995) approach. The Toda—Yamamoto approach fits a vector auto regressive model in the levels of the variables thereby minimizing the risks associated with the possibility of wrong identification of the order of cointegration of the series (Mavrotas and Kelly, 2001). Following Altgintas and Taban (2011) we specify the following two variables VAR model:

$$\begin{aligned} Y_t &= \alpha + \sum_{i=1}^{k} \beta_i Y_{t-i} + \sum_{j=k+1}^{k+d_{\max}} \beta_j Y_{t-j} + \sum_{i=1}^{k} \delta_i X_{t-io} + \sum_{j=k+1}^{k+d_{\max}} \delta_j X_{t-j} + \omega_{1t} \\ X_t &= \alpha + \sum_{i=1}^{k} \lambda_i X_{t-i} + \sum_{j=k+1}^{k+d_{\max}} \lambda_j X_{t-j} + \sum_{i=1}^{k} \theta_i Y_{t-i} + \sum_{i=k+1}^{k+d_{,\max}} \theta_j Y_{t-j} + \omega_{2t} \end{aligned} \quad (11)$$

Where k is the number of lags and $d_{\max}$ represents the maximum cointegration level of the variables entered in to the model. The main idea of this method is to to increase the number of lags in the VAR model in accordance with the maximal cointegration relationship based on the F-stat. If $\delta_i \neq 0$ in equation (11), X_t Granger causes Y_t. Similarly, if $\theta_i \neq 0$ in equation (11), Y_t Granger causes X_t. Table 2 reports the results of Toda-Yamamoto causality.

Table 2: Toda-Yamamoto Causality Results

Model	Lag length $(k + d_{\max})$	$F - stat$	p-value	Decisions	Direction of causality
$CADY_t = f_1(BDY_t)$	2	3.1364	0.0590**	Reject H_0	$BDY_t \rightarrow CADY_t$
$BDY_t = f_2(CADY_t)$	2	1.1421	0.3336	Accept H_0	$CADY_t$ *does not cause* BDY_t
$INVY_t = g_1(CADY_t)$	2	0.6947	0.5068	Accept H_0	$CADY_t$ *does not cause* $INVY_t$
$CADY_t = g_2(INVY_t)$	2	0.8027	0.4572	Accept H_0	$CADY_t$ *does not cause* $INVY_t$
$BDY_t = h_1(INVY_t)$	2	1.3456	0.2767	Accept H_0	$INVY_t$ *does not cause* BDY_t
$INVY_t = h_2(BDY_t)$	2	2.2549	0.1236	Accept H_0	BDY_t *does not cause* $INVY_t$

Note: * indicate significant at the 5 percent level of significance.

According to Todao-Yamamoto causality test budget deficit causes current account deficit and current account does not causes budget deficit. This confirms our earlier results of the existence of twin deficits hypothesis with causality running from budget deficit to current account

deficits. Furthermore, we find no causality between investment and current account and budget deficit and investment.

5.1 Policy Implications

We find supportive results for the validity of twin deficit hypothesis for Pakistan over the period 1976-2010. The two deficits appear to tied up by a statistically significant relations in the case of Pakistan. Large budget deficit leads to higher current account balance both in short-run as well as in long-run. This twin deficit is eroded in the long-term and appears to be relatively small in the short-run. There are various channels through which budget deficit influences current account deficit in an economy like Pakistan. First, an increase in budget deficit induces upward pressure on interest rate which leads to capital inflows and an appreciation of exchange rate, leading to increase in current account deficit. Second, an increase in budget deficit would induce domestic absorption, and hence increase in imports worsens current account balance (Marinheiro, 2008). Third, budget deficit increases domestic and external borrowings. Domestic borrowing leads to a credit squeeze through higher interest rate which, in turn, crowds out private investment and consumption. External borrowing leads to a current account deficit and appreciation of the real exchange rate creates balance of payments crisis or external debt crisis. High budget deficit also deteriorate current account balance through inflation rate (Easterly and Schmidt-Hebbel, 1993). Furthermore, when taxation revenues are not enough to cover government expenditures, an increase in public expenditure can cause budget deficit. Therefore, fiscal discipline is helpful to stabilize current account balance. To this end, there is need to coordinate and implement monetary, fiscal and exchange rate policies to promote macroeconomic stability and sustainability. Furthermore, there is also need to increase tax-GDP ratio through rationalization of tax system. Export promotion could be another policy option that the authority could pursue.

Another important finding of this study is the negative relationship between investment and current account deficit. This finding does not support the validity of Feldstein-Horioka hypothesis. This implies that Pakistan's economy is integrated to the world economy. The result could be due to the adoption of economic liberalization policies of the

government of Pakistan since 1990. Therefore, there is need of further liberalization to strengthen the integration between domestic financial markets to the rest of the world.

Toda-Yamamoto causality result also supports the causal relationship between budget deficit and current account deficit with the causality running from budget deficit to current account deficit. No evidence of causality between either investment and current account deficit or investment and budget deficit has been observed.

6. Conclusions

This study examines the validity of the twin deficits and Feldstein-Horiopka hypothesis for Pakistan over the period 1976-2010 using ARDL-bounds testing cointegration methodology. The validity of twin deficits hypothesis has strong policy implications, and closely linked to the Feldstein-Horioka hypothesis. It is widely believed that if the twin deficits hypothesis is valid, the appropriate policy option could be the fiscal discipline to correct current account deficit The results reveals the existence of long-run relationship between current account balance, budget deficit and investment. The presence of cointegration between the variables implies that current account balance, budget deficit and investment are moving jointly. The positive association between budget deficit and current account deficit implies that Pakistan has a twin deficits problem. Budget deficits worsen current account deficits in the short-as well as in the long-run. The causality runs from budget deficit to current account deficits in short-as well as in long-run. Furthermore, our findings do not support the validity of the Feldstein-Horioka hypothesis in the long-run but lend weak support in the short-run. This is because Pakistan's economy is integrated to the world economy, though; the degree of international capital mobility is not perfect. Domestic investment particularly, public sector investment has largely depends on domestic savings as well as foreign borrowings and foreign economic assistance has always played an important role in financing the national development programmes. Our results suggest the ¾ of domestic investment is financed through external sources. Thus, the historical relationship between domestic investment and saving is very poor in

Pakistan. As such, the empirical results do not support the validity the Feldstein-Horioka puzzle in case of Pakistan in the long-run.

The evidence of twin deficits and non-existence of Feldstein-Horioka hypothesis in the case of Pakistan is also verified by Toda-Yamamoto causality test.

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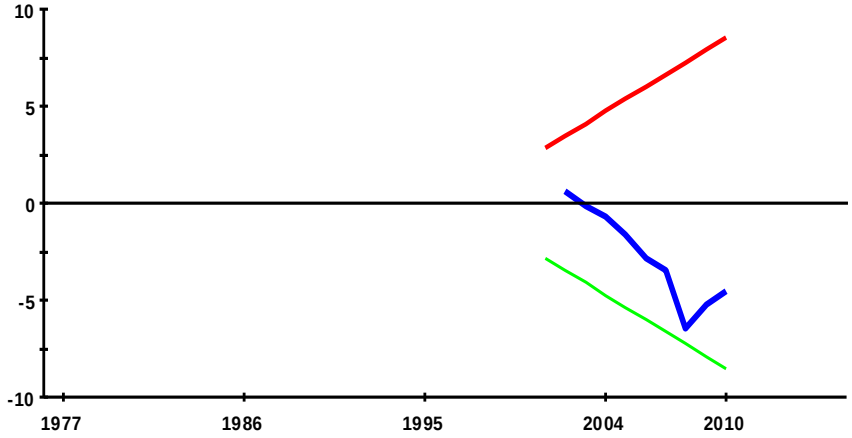
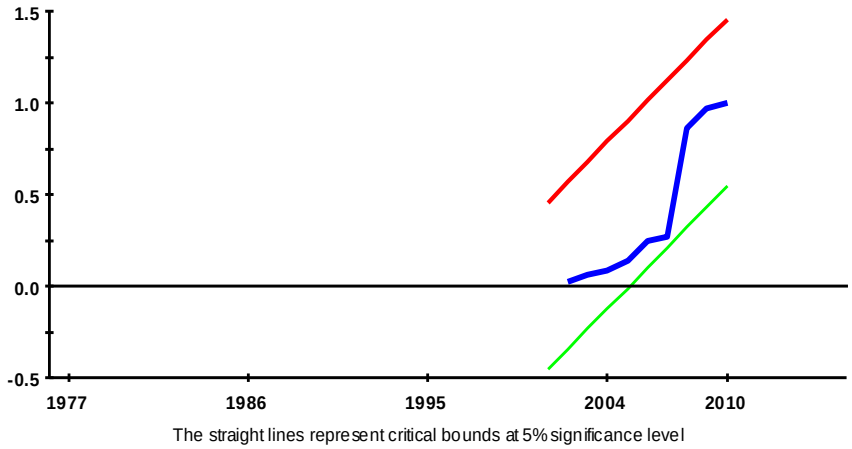
Appendix

Plot of Cumulative Sum of Recursive Residuals



The straight lines represent critical bounds at 5% significance level

Plot of Cumulative Sum of Squares of Recursive Residuals



The straight lines represent critical bounds at 5% significance level

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Dynamic Causal Interactions of Money, Prices, Interest Rate and Output in Pakistan

Faiz Bilquees, Tahir Mukhtar and Sidra Sohail[1]

Monetary policy has remained one of the most fundamental topics in macroeconomics. Since the beginning of macroeconomic history it has been argued that money has a strong role in affecting the real economic activity but the evidence remains inconclusive. Considerable empirical evidence has been produced on the stance of monetary policy using different approaches and still the process continues. The present study investigates the dynamic interactions among macroeconomic variables such as money supply, prices, interest rate, exchange rate and output level, using the quarterly data for Pakistan over the period 1972Q1 to 2009Q4. For the empirical analysis the Johansen multivariate cointegration technique, Granger causality test and variance decompositions are employed. The results from the cointegration test indicate that there exits a stable long run equilibrium relationship among the macroeconomic variables of the study. The outcome of causality tests tends to support the non neutrality of money view of the Keynesians and the Monetarists at least in the short run. Furthermore, it is seen that there exists a bi-directional causality between money supply and price level, and interest rate and price level. While, a unidirectional causality runs from money supply to output level and interest rate, the opposite does not happen. The findings tend to indicate that money supply, real output, interest rate and exchange rate is Granger causing prices in the short run as well as in the long run. This implies that inflation is not purely a monetary phenomenon rather structural factors also have a role in affecting price level in Pakistan. Hence, in combating inflation decision makers need not rely exclusively on the instrument of tight monetary policy rather due attention should also be given to the supply side of the economy.

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"Though many macroeconomists would profess little uncertainty about it, the profession as a whole has no clear answer to the question of the size and nature of the effects of monetary policy upon aggregate activity." Sims (1992)

1. Introduction

Considerable empirical literature exists on explaining the links between monetary and real macroeconomic activities. This literature culminates in near consensus on the long term relationship between monetary and real economic activities as the long run neutrality of money is well established in economic literature. However, the short run interactions among the monetary and real variables, which are of vital importance for the conduct of monetary policy, are still widely debated by the economists. While the New Classical economists maintain that prices are flexible and adjust *quickly* to clear the markets, the New Keynesians argue that market-clearing models are unable to explain short run economic fluctuations and believe in *stickiness of prices*. Based on the *sluggish adjustment* of prices, both the Keynesians and the Monetarists recognize the effects of monetary policy on production activities in the short run (Mankiw and Romer 1991). According to this view, if the money supply falls, people spend less money and the demand for goods falls. Because prices and wages are inflexible and do not fall immediately, the decreased spending causes a drop in production and layoffs of workers.

The linkages among money, output, interest rate and prices have been the focus of extensive debate and analysis. Much of the debate among the New Keynesians and the New Classicals is centered on the relative effectiveness of monetary policy for influencing the economy. The discussion however is still inconclusive, past empirical evidence on the relative contributions of money and credit in the propagation of monetary policy impulses still remains ambiguous. Sims (l972, 1980a), Thornton and Batten (1985), King (1983), Stock and Watson (1989), Romer and Romer (1990), Cover (1992), Christiano et.al., (1999) Mishkin (2002), Siregar and Bert (2002), Gilman and Anton (2004), Ravn and Martin (2004), Bernake et al., (2005), Hossain (2005), Dickinson and Jia (2007),Rafiq and Mallick (2008) and many others present evidence of monetary effects. While results of money neutrality

are reported, among others, by Feige and Pearce (1979), Sims (1980b), Litterman and Weiss (1985), Geweke (1986), Friedman and Kuttner (1993), King and Watson (1997), Cochrane (1998), Serletis and Koustas (1998), Ganev et.al., (2002) and Starr (2005).

While considerable evidence exists on the linkages between money and economic activity in developed countries, literature on developing countries is quite limited. In case of Pakistan the analysis on the subject is limited to a few studies: including Hussain (1982, 1991); Chishti et. al.,(1992); Momen (1992); Ahmed (2003); Mehmood and Mohammad (2005); Abbas and Fazal (2006); Khan (2008) ; and Husain and Rashid (2009). However, since these studies are beset with a number of methodological and theoretical shortcomings their findings do not provide reliable estimates as basis to frame effective economic policies, as described in section 2of the study. Hence, the main objective of the present study is to investigate the dynamic causal chain among key aggregate economic variables like stock of money, price level, interest rate and output level in Pakistan considering the methodoligical limitations of the existing literature. In order to examine the dynamic interactions of these variables with the foreign trade sector, we will also incorporate the exchange rate variable as well. For the causal analysis the study employs the Johansen cointegration, the vector error correction model (VECM) and the variance decompositions (VDCs) techniques.

The rest of this paper is organized as follows: section 2 presents a brief review of the existing studies conducted in Pakistan; analytical framework for this study is given in section 3; section 4 describes the results of the study; and finally section 5 concludes the study with some policy implications.

2. Review of literature

In the empirical literature, the relationship between money, price, and output has been investigated by researchers for different countries over different sample periods and provided the conflicting evidences on this issue. In this section we just confine ourselves to review some of the past studies which have empirically investigated the relationship between money and some important macroeconomic variables in Pakistan. Hussain (1982) is a pioneer investigation into the

interrelationship among money, prices and output in Pakistan. Based on the annual data for the period 1949 to 1970, the study uses a single equation model to test the relative impacts of monetary and fiscal impulses on nominal income. The findings of the study show that the response of economic activity to fiscal actions compared with that of monetary actions is larger, more predictable and faster. However, it is important to note that while the single equation model is simple and easy to estimate, the equations are not derived explicitly from a larger model and therefore important feedback mechanism may be omitted. If the right hand side variables in the equations are not exogenous, the equation may be a part of system of equations where variables are interdependent. This is a serious flaw which renders the study to be a purely academic exercise without any policy substance.

Hussain (1991) applies Sims (1972) causality test procedure for the period 1971 to 1988 for determining the nature of relationship between money and income in Pakistan. M1, M2 and monetary base are used as money variables and GNP for income level. The F-tests are conducted by choosing one year past value and one year future value of the regresses. He summarizes his findings as follows: (i) a unidirectional causality runs from monetary base to GNP; (ii) a unidirectional causality runs from M2 to GNP; and (iii) a unidirectional causality runs from GNP to M1.The results of this study, however, are suspected due to two serious considerations with regard to methodology: (a) Sims test is particularly sensitive to the lag structure but the study reports one period lag without any statistical and economic justification; (b) the F-test used to determine the direction of causality is actually a t-test of a single coefficient. Results based on such methodological anomalies can have serious policy implications.

In the early nineties research on monetary issues in Pakistan adopted the multivariate approach and Chishti et al., (1992) was the first study to apply VAR model to the Pakistani data. The authors used annual data covering the period 1960 to 1988 and estimated a VAR model with ten macroeconomic variables: real GDP, consumer price index, terms of trade between agriculture and manufacturing sectors, unemployment rate, real investment, real value of remittances, real exports, real external resources, money stock and real government expenditure. The results showed that F-tests for causality found a unidirectional causality running

from money to output and price does not cause money but money does cause price. The Impulse response analysis revealed that money produced a strong positive impact on real GDP and general price level. A major limitation of the study is that it includes too many variables when data on variables spans only from 1960 to 1988. To save degrees of freedom lag length is truncated to only two periods, which is not sufficient to capture the dynamics of the issues involved. Furthermore, when too many variables are included in a VAR model, additional complications arise; the simultaneous relations among different variables and policy innovations make it difficult to correctly identify the shocks.

Momen (1992) has conducted a study for ten industrial and agricultural countries including Pakistan, for assessing the interaction among the rate of inflation, the rate of change in real gross domestic product, the rate of change in terms of trade, the rate of change in government expenditure and the rate of change in money supply. The study covers the time period from 1958 to 1985.The author constructs a reduced form VAR model where each of the five variables is regressed on past values of itself and past values of the other four variables in the system. He conducts F-tests to determine causal relationship among the variables and concludes that in the industrialized countries causality runs from money supply to real GDP, which is in conformity with the Monetarists view. In agricultural economies including Pakistan, causality runs in the opposite direction. Though this study is theoretically sound but it is subject to two shortcomings: (a) serious questions can be raised about the data set used. The study has used data from 1958 to 1985 for all countries but Bangladesh gained independence in 1971, before that it was a part of Pakistan. Thus, the author uses the same data for both the countries up to 1971, which is faulty and definitely gives misleading conclusions for both Bangladesh and Pakistan; and (b) he does not orthogonalize the shocks which is very important to isolate their effects. He employs the reduced form innovations to compute variance decomposition to draw different conclusions. But if reduced form errors are correlated, which is most likely, this methodology may lead to erroneous conclusions and that the estimated variance-covariance matrix of the errors reflects the seriousness of the problem.

Ahmed (2003) investigates the issue of causality among key aggregate macro-variables by conducting bivariate, multivariate and block causality tests. For Pakistan the study has used the quarterly data from1972-I to 1997-II. Causality tests suggest that bidirectional causality exists between money and prices in Pakistan. The policy implication of such a result is that an increase in money stock fuels prices in Pakistan, which in turn leads to an increase in money stock. This study has accurately conducted causality analysis among key aggregate macro-variables but for some reasons it ignores the dynamic analysis; it does not apply a single measure of dynamic analysis such as variance decompositions (VDCs) and impulse response functions (IRFs).

Mehmood and Mohammad (2005) attempt to estimate the long run and the short run relationships among the key macroeconomic variables *viz*., money, prices, interest rate and output in Pakistan. The study uses annual data over the period 1973 to 2003. Applying Johansen and Juselius (1990) cointegration technique, the study shows that there exists one cointegrating vector among the four variables. On the basis of error correction model (ECM), it is found that a unidirectional causality running from money to output and from money to interest rate exists. However no causality exists between output and interest rate. Money and prices have been found to be independent in the short run. In summary, the study shows that money supply is an appropriate intermediate target (with output growth being the final target) and not the interest rate. This study is a significant improvement over the previous studies, however it is subject to the following shortcomings: (a) though the authors have conducted the F-tests as multivariate tests but actually they are bi-variate causality tests because they consider lagged coefficients of a particular variable in a single equation of the system not the other equations of the model. Only a likelihood ratio test can do this job (see Enders, 2004); (b) the result of ECM indicates the exogeneity or endogeneity of a variable in the system and the direction of Granger-causality within the sample period. However, it does not provide us with the dynamic properties of the system. The analysis of the dynamic interactions among the variables in the post-sample period need to be conducted through variance decompositions (VDCs) and impulse response functions (IRFs) that are missing in the study.

Using annual data for fiscal years 1959 to 2003 and employing cointegration and error correction models as well as the standard Granger causality analysis Abbas and Fazal (2006) investigate the bi-variate and tri-variate causal relationships between money and income and between money and prices in Pakistan. Regarding the causal relationship between money and prices, the causality framework provides the evidence of bi-variate causality indicating that monetary expansion increases, and is also increased by inflation in Pakistan. However, money supply seems to take the lead in this case. But this study is beset with two weaknesses :(a) in the estimated regression equation lagged values of one variable are regressed on another variable. This is essentially a two variable single equation distributed lag model and is seriously subject to omitted variable bias; (b) the study has used the Engle-Granger cointegration approach which is beset with many defects as outlined in different text books (for example, see Asteriou and Stephen, 2007).

A study by Khan (2008) provides an empirical update on the impact of an unanticipated change in monetary policy on output growth and inflation in Pakistan. The study uses monthly data for the period 1991-VII to 2006-IX and adopts multivariate structural vector auto-regressions (SVAR) technique with long run restrictions based on standard aggregate demand and supply model of the economy. The results indicate that an unanticipated positive shock in monetary policy leads to: (i) an increase in industrial output, which reverts to its original level over 23 to 32 months horizon; (ii) an increase in inflation; and (iii) nominal shocks remained the dominant factor in explaining variation in inflation as compared to supply disturbances. Transmission mechanism is much faster in case of Consumer Price Index (CPI) and the CPI remains unchanged. Despite the fact that this study overcomes many of the shortcomings of past studies, its findings are subject to a major data weakness. GDP has been proxied compared to Industrial Production Index (IPI), as over 75 percent increase in CPI is realized during 12 months after the shock and this impact touches over 90 percent level during 18 months. Sensitivity of these results to another specification indicates that response patterns of both IPI and CPI remained unchanged. However, IPI does not constitute a major share of GDP in Pakistan, it is only around 20 percent of GDP. Hence, it is erroneous to use this as a proxy for GDP, the policy implications of the study cannot be broad based.

Husain and Rashid (2009) attempt to extend the analysis of causality between money and the two macroeconomic variables i.e., income and prices by taking care of the shifts in the variables due to the price hikes in the early 1970s and the economic liberalization program of the early 1990s. For introducing the expected shifts in the variables dummies have been used. The results indicate significant shifts in the variables during the sample period. In this context, the shift that occurred due to price hikes in the early 1970s seems to be more important to be incorporated in the analysis. This study finds the active role of money as the leading variable in changing prices without any feedback and this relationship is not affected by the shifts during the sample period. However, when both the shifts are introduced in the analysis it is seen that both the variables cause each other in the long run and they are independent of each other in the short run. No doubt the motive behind the study to examine how income and price are causally related with money in Pakistan in the presence of some shifts seems to be quite inspiring, but the way the study proceeds to achieve its objective is erroneous as acknowledged by the authors in stating the limitations of the study. These limitations include: (a) use of bi-variate causal analysis ;(b) inclusion of pre 1971 period in annual data set ;(c) use of OLS estimation technique. However, besides these limitations, the authors fail to realize that: a) with the inclusion of two dummies for the shifts in the money-income and money- price models, the regression analysis has become multiple in nature, while the use of Engel Granger cointegration technique only provides one cointegrating vector despite the possibility of more than one long run relationship among variables in a multiple regression model. Since the authors fail to provide evidence that there does not exist more than one cointegrating vectors, the findings of the study cannot be validated; b) the study is silent on the impact of unexpected component of one variable on the other variable in the causal analysis; c) the study does not provide any theoretical reasoning that how the inclusion of the two shifts in the analysis is expected to affect the causality pattern and long run relationship between the variables[2].

[2] While conducting causal analysis in the presence of two shifts in the model, the authors have certainly found the causal pattern between two shifts or among one or both of these shifts and other macroeconomic variables of the study. In such circumstances it is essential for the authors to provide a theoretical justification for such causal patterns.

In the light of the above review of the existing studies it is seen that literature on the dynamic relationship between money and other important macroeconomic variables in Pakistan does not converge on any specific conclusions, presumably, due to the methodological limitations. Therefore, the present study is of significance in that it attempts to explore the dynamic causal link between stock of money and other important macroeconomic variables in Pakistan using the Johansen cointegration, the VECM and the VDCs techniques. The VECM technique is suitable for examining the short and the long run casual links among the variables while for gauging the strength of the casual relations the VDCs technique is used. The past researches in the context of Pakistan did not make proper use of these techniques.

3. Analytical Framework

This study employs the multivariate cointegration analysis, the Granger-causality test within the environment of vector error-correction modeling, the variance decompositions, and the impulse response functions to analyze the dynamic relationships among money, price level, interest rate and real output in Pakistan.

3.1. Stationarity of Variables and Unit Root Tests

To diagnose stationarity of the variables a number of tests have been proposed in the literature. Among them the Augmented Dickey-Fuller (ADF) test of Dickey and Fuller (1979, 1981), and the Phillips-Perron (PP) test of Phillips and Perron (1988) are frequently used. However, because of their poor size and power properties these tests are not reliable for small sample data sets (Dejong et.al., 1992; Harris and Sollis, 2003)[3]. In such a situation, we prefer to apply a more efficient and powerful univariate Dickey-Fuller Generalized Least Square (DF-GLS) test. This test is basically a modified version of the ADF test in which data are detrended before the unit root test is conducted. It is developed by Elliot et.al.,(1996) and is based on the null hypothesis $H_0 : \alpha = 0$ in the regression:

[3] Both these studies conclude that the ADF and the PP tests have the tendency to over-reject the null hypothesis when it is true and under- reject it when it is false.

$$\Delta y_t^d = \alpha y_{t-1}^d + \alpha_1 \Delta y_{t-1}^d + \cdots\cdots + \alpha_p \Delta y_{t-p}^d + \eta_t \quad (1)$$

where y_t^d is de-trended series of variable y_t and η_i is the independently and identically distributed error term[4].

3.2 Cointegration Test

The econometric framework used for analysis in the study is the Johansen (1998) and Johansen and Juselius (1990) Maximum-Likelihood cointegration technique, which tests both the existence and the number of cointegration vectors. This multivariate cointegration test can be expressed as:

$$Z_t = K_1 Z_{t-1} + K_2 Z_{t-2} + ... + K_{k-1} Z_{t-k} + \mu + v_t \quad (2)$$

where,

$Z_t = (m, p, r, , gdp, er)$ i.e. a 5 x 1 vector of variables that are integrated of order one [i.e. I (1)],*m, p, r, rgdp* and *er* are broad money supply(M2), price level, nominal interest rate, real gross domestic product and nominal exchange rate respectively,

$\mu =$ a vector of constant and $v_t =$ a vector of normally and independently distributed error term.

The equation (2) can be reformulated in a vector error correction model (VECM) as follows:

$$\Delta Z_t = \Gamma_1 \Delta Z_{t-1} + \Gamma_2 \Delta Z_{t-2} + \cdots\cdots + \Gamma_{k-1} \Delta Z_{t-k-1} + \Pi Z_{t-1} + \mu + v_t \quad (2)$$

Where $\Gamma_i = (I - A_1 - A_2 ..... - A_i)$ (i= 1,2,3…..k-1) and $\Pi = -(I - A_1 - A_2 - A_3 ..... - A_k)$. The coefficient matrix Π provides information about the long run relationships among the variables in the data. Π can be factored into $\alpha\beta'$ where α will include the speed of adjustment to the equilibrium coefficients while the β' will be the long run matrix of

[4] For detailed discussion on different unit root tests see Maddala and Kim (1998).

coefficients. The presence of r cointegrating vectors between the elements of Z implies that Π is of the rank $r(0 < r < 5)$. To determine the number of cointegrating vectors, Johansen developed two likelihood ratio tests: Trace test (λ_{trace}) and maximum eigenvalue test (λ_{max}.). The null hypothesis that there are at most r cointegrating vectors is evaluated by the trace test, the statistic of which is calculated as follows:

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

The maximal eigenvalue test, instead, evaluates the null hypothesis that there exist r cointegration vectors against the alternative of the existence of r+1 cointegration vectors. This test statistic is calculated as follows:

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

In both (3) and (4) T is the number of usable observations and λ represents the calculated values of the characteristic roots from the estimated matrix. If there is any divergence of results between these two tests, it is advisable to rely on the evidence based on the λ_{max} test because it is more reliable in small samples [see Dutta and Ahmed (1997) and Odhiambo (2005)].

3.3. Variance Decompositions

The VECM, F- and t- tests may be interpreted as within-sample causality tests. They can indicate only the Granger causality of the dependent variable within the sample period. They provide little evidence on the dynamic properties of the system, the relative strength of the Granger-causal chain among the variables. On the other hand, the variance decompositions (VDCs), by partitioning the variance of the forecast error of a certain variable into the proportions attributable to innovations (or shocks) in each variable in the system including its own, can provide an indication of these relativites[5]. The VDCs may be termed

[5] VDCs tells us the proportion of the movements in a variable due to it "own" shocks versus shocks to other variable.

as out-of-sample causality tests (Masih and Masih 1996). The variable that is optimally forecast from its own lagged values will have all its forecast error variance explained by its own disturbances (Sims, 1982).

3.4. Data

This study uses quarterly data for the period 1972Q1 to 2009Q4 for the empirical exercise. The price level is represented by the consumer price index (CPI) because it is a good indicator of the movement of prices (Crocket and Evan, 1980). The bank rate (discount rate) is used for the nominal interest rate variable. The bank rate is the main lever that a central bank uses to conduct monetary policy, it is the rate of interest rate that a central bank charges on short term loans to financial institutions and is considered as the trend setter for other short term interest rates. Real GDP is used as a measure of real output level, it is computed by deflating the nominal GDP by the CPI. Finally, M2 (broad money) represents the money stock variable. As in an open economy monetary policy operates through interest rate and exchange rate channels, so we have also included exchange rate in our analysis. Hence, monetary shocks are transmitted to the real sector of the economy through both the channels of monetary policy The exchange rate is represented by the ratio of Rupees per US dollar. All the variables are logarithmic except for the nominal interest rate. The required data are sourced from various *Quarterly and Annual Reports* of the State Bank of Pakistan, *Pakistan Economic Survey* (various issues) and *International Financial Statistics* (CD- ROM) ,IMF.

4. Results and Discussion

4.1. Stationarity Test

Therefore, to check the stationarity of variables, we use the DFGLS test. From the results of the DF-GLS test presented in table 1, it is evident that the null hypothesis of a unit root can not be rejected for all the variables in levels. This hypothesis, however, is rejected for the first-differences. Thus, all variables are stationary in their first-differences i.e., they are integrated of order one i.e., *I(1).*

Table 1. DF-GLS Unit Root Test

			Mackinnon Critical Values for Rejection of Hypothesis of a Unit Root			
Variables	Level	First Difference	1 %	5 %	10 %	Order of Integration
m	-1.37	-7.44	-3.51	-2.98	-2.69	*I* (1)
r	-2.27	-12.96	-3.51	-2.98	-2.69	*I* (1)
p	-0.59	-6.91	-3.51	-2.98	-2.69	*I* (1)
rgdp	-1.47	-9.88	-3.51	-2.98	-2.69	*I* (1)
er	-1.13	-8.52	-3.51	-2.98	-2.69	*I* (1)

4.2. Multivariate Cointegration Analysis

In the next step, we determine the optimal lag length for cointegration analysis because Johansen technique is known to be sensitive to the lag length. As far as this study is concerned the Schwarz Bayesian Criteria (SBC) suggests a lag length of 5 as optimal, which is not surprising for quarterly data. The cointegration test is conducted assuming an intercept in the cointegrating equation. Cointegration relationship among m , r , p , $rgdp$ and er has been investigated using the Johansen technique. Table 2 reports our cointegration test results based on Johansen's maximum likelihood method. Both trace statistic (λ_{trace}) and maximum eigenvalue (λ_{max}) statistic indicate that there is at least one cointegrating vector among all the five time series. We can reject the null hypothesis of no cointegrating vector in favour of one cointegrating vector under both test statistics at 5 percent level of significance. However, we cannot reject the null hypothesis of at most one cointegrating vector against the alternative hypothesis of two cointegrating vectors, from both the trace and max-eigen test statistics. Consequently, we can conclude that there is only one cointegrating vector among m , r , p , $rgdp$ and er . This implies that money supply, interest rate, price level, real GDP and exchange rate are bound together by long run equilibrium relationship in Pakistan.

Table 2.Cointegration Test Based on Johansen's Maximum Likelihood Method

Null Hypothesis	Alternative Hypothesis			Critical Values	
				95 %	P-values●●
λ_{trace} rank tests		Eigen values	λ_{trace} rank value		
$H_0 : r = 0$	$H_1 : r = 1$	0.28	94.26**	69.81	0.00
$H_0 : r = 1$	$H_1 : r = 2$	0.15	43.78	47.86	0.13
$H_0 : r = 2$	$H_1 : r = 3$	0.11	23.99	29.79	0.21
$H_0 : r = 3$	$H_1 : r = 4$	0.03	6.88	15.49	0.59
$H_0 : r = 4$	$H_1 : r = 5$	0.01	1.98	3.84	0.16
$\lambda_{\max}$ rank tests			$\lambda_{\max}$ rank value		
$H_0 : r = 0$	$H_1 : r > 0$	0.28	47.48**	33.87	0.00
$H_0 : r \leq 1$	$H_1 : r > 1$	0.15	22.78	27.58	0.18
$H_0 : r \leq 2$	$H_1 : r > 2$	0.11	17.11	21.13	0.16
$H_0 : r \leq 3$	$H_1 : r > 3$	0.03	4.91	14.26	0.75
$H_0 : r \leq 4$	$H_1 : r > 4$	0.01	1.98	3.84	0.16

** denotes rejection of the null hypothesis at the 5 percent significance level.
●● MacKinnon-Haug-Michelis (1999) p-values.

Trace test indicates 1 cointegrating equation at 5 percent significance level.

Max-eigenvalue test indicates 1 cointegrating equation at 5 percent significance level.

4.3. Granger-Causality Tests

Although cointegration indicates presence or absence of a long run relationship, it does not indicate the direction of causality among the variables. In this regard, we estimate a VECM to conduct a causal

analysis. The VECM allows the long run behavior of the endogenous variables to converge to cointegrating (i.e., long run equilibrium) relationships while allowing a wide range of short run dynamics. Following Masih and Masih (1997) and Tan and Ahmad (1999), estimates from the VECM are used to conduct the short run and the long run dynamic causal analysis. By adding error correction term (ECT) in VECM, it provides an additional channel for long run causality which is ignored by Sims and Granger standard causality tests. Long run causality is confirmed through the significance of the coefficient of lagged ECT and short run causality is confirmed through the joint significance of coefficients of lagged variables. F-test is employed to check joint significance of the coefficients of the lagged variables and t-test is used to check significance of the lagged ECT. Table 3 reports the Granger causality test results based on the VECM.

Table 3.Dynamic Causal Chain Based on VECM

Dept Var						ECT_{t-1} (t-statistic)
	Δm	Δr	Δp	$\Delta rgdp$	Δer	
	F-statistic					
Δm	-	0.54	5.68**	0.42	0.33	-0.08 (-0.27)
Δr	12.29***	-	3.62*	1.46	0.88	0.06 (0.44)
Δp	10.74**	6.37**	-	6.19**	5.54**	-0.31 (2.98)**
$\Delta rgdp$	5.69**	4.19*	0.88	-	4.53*	0.15 (0.97)
Δer	4.16*	3.77*	1.51	1.18	-	0.11 (1.24)

Note: ***, **, *, indicate significance at 1 percent, 5 percent and 10 percent probability level respectively t- values are given in parentheses.

With regard to the short run analysis we find that there is a bidirectional feedback relationship between price level and money supply. This result is consistent with the conclusion of the Quantity Theory of Money that money supply expansion pushes price level up while it also suggests that price is helpful in predicting the current and future growth rates of money supply in Pakistan. However, there is a unidirectional causal link

from money supply to interest rate, exchange rate and output level. This finding signifies that the role of the interest rate and the exchange rate channels are operating in the economy as both these variables establish a significant relationship with output level. Furthermore, unidirectional causality from stock of money to real output implies that money is no more neutral and money is a lead output indicator in the short run in Pakistan. This outcome supports the Keynesians and the Monetarists view of non-neutrality of money. Hence, policy makers could influence the sustainable output growth with money supply stimulus. The presence of Granger causality from money to output and price suggests that exogenous monetary policy shocks are a major source of output and price variability in Pakistan. This result is in line with the findings of Ramachandra (1986), Chishti et al., (1992), Momen (1992), Masih and Masih (1997), Tan and Ahmad (1999), Das (2003), Dritsaki and Antonios (2005), Mehmood and Mohammad (2005) and Khan (2008).

Exchange rate and real output level both Granger- cause the price level. From this finding it is evident that changes in the exchange rate are transmitting to domestic price level. Hence, the devaluation of the Pakistan rupee in terms of dollar tends to generate inflationary pressures in the economy. As both the stock of money and the real output level cause the price level, so, inflation is not only a monetary phenomenon but real sector also plays a role in its generation. The evidence of causality from real output to price suggests that the excess of aggregates demand generated by increase in real GDP is not absorbed by growth in aggregate supply. There is a strong evidence of a unidirectional short run causal effect running from interest rate to exchange rate. This causality is consistent with the traditional view whereby the interest rate is necessary to support the exchange rate, as changes in interest rates affect the returns for investing in a country and in turn affect the amount of capital outflows. Similar finding has been reported by Nwosa and Isiaq (2012) for Nigeria. Thus, the result shows that interest rate turns out to be the tool that can be used to influence the exchange rate, at least in the short run. There appears to exist a bidirectional causation between price level and interest rate. A good policy variable should be free from feedback from non-policy variables such as price and output in the model. So the interest rate does not deserve to be a good policy variable at least from the perspective of causality test in Pakistan.

From table 3 it can be inferred that coefficients of ECT carries the expected negative sign only in the functions of money supply and price level. However, coefficient of ECT is significant only when price level is used as a dependent variable. For the functions of all the four remaining variables coefficients of ECT are not significant. Hence, it suggests that in the long run money supply, interest rate, real output level and exchange rate Granger- cause price level without any feedback effect. This finding is consistent with our short run causal analysis. A significant negative coefficient means that whenever the actual value of price level rises above the value consistent with its long term equilibrium relationship, changes in the independent variables help bring it down to the long term equilibrium value, other things being equal. It is in this sense that the ECT provides an additional channel of causal relationship. The size of the coefficient of ECT is interpreted to indicate the speed of adjustment. Therefore its coefficient of - 0.31in the equation of price level implies that up to 31 percent of the adjustment in the imbalance is corrected in every quarter which is a reasonably good speed of adjustment. There is absolute absence of any causal pattern among money supply, interest rate, exchange rate and real output in the long run. As money supply does not cause real output, so money is neutral in the long run in Pakistan. Same conclusion has been reached by Moosa (1997), Wallace (1999), Bae and Ratti (2000) and Sulku (2011) for different developing countries.

4.4. Variance Decomposition Results

As stated earlier, although the VECM can give us an understanding of the direction of Granger-causality within the sample period it does not provide us with an indication of the dynamic properties of the system, nor does it allow us to gauge the relative strength of the variables beyond the sample period. In order to analyze the dynamic properties of the system the forecast error variance decompositions (VDCs) are computed. The results of the relative contribution of the explanatory variables in explaining the variation in the dependent variable in the post-sample period are presented in table 4. Though there is no hard and fast rule regarding the number of time periods to be examined but it should be enough to understand the dynamic interactions among the variables. We have examined 20 quarters, which is five years worth of time horizon. We adopt the methodology of orthogonalised forecast

error variance decomposition, which is based on Choleski factorization with particular ordering[6]—namely, m , r , er , $rgdp$ and p .

The decomposition is taken in percentage form at different horizons, and the results for the VDCs are found to be almost consistent with that of the causality tests. It may be seen from table 4 that the price level and the money supply own innovations remain a dominant source of inducing variations in the money supply for short, medium and long run. The results show that 98 percent of the forecast error variance of nominal interest rate is explained by its own shock in the first year; however, the impact of its own shock on interest rate declines over the remaining time horizon to 75 percent in the long run. After its own shocks, the money supply is the most important in explaining the variance in interest rate in the short run and the long run. Looking at the forecast error variance of exchange rate we note that exchange rate own innovations and money supply contribute relatively more as compared to other variables in the system, not only in the short run but also in the long run. In case of real output its forecast error variance is mainly explained by its own shock at 20 quarters horizon. The exchange rate is the next most important factor in explaining the variation in the real output, however, its contribution remained below 10 percent in the short run and the long run. Finally, the predominant source of variation in price level is the "own" shock. Money supply and real output are relatively more important source of the forecast error variance in price level which again confirms our earlier finding under the causality tests that in determining the price level both the monetary and real factors play their role. Overall, the results of the VDCs show a strong and leading role of money over the remaining variables.

[6] Other orderings also produce almost the same results.

Table 4.Variance Decomposition Analysis

Horizon	Percentage of variance due to:					
	Relative variance in	Δm	$\Delta rgdp$	Δr	Δp	Δer
1	Δm	100.00	0.00	0.00	0.00	0.00
5		93.29	0.15	0.16	0.97	5.43
10		84.31	0.84	2.31	3.84	8.7
15		77..52	2.34	3.18	4.16	12.8
20		71.33	1.94	2.98	4.83	18.92
1	Δr	1.74	98.26	0.00	0.00	0.00
5		6.55	87.84	0.49	0.31	4.81
10		8.87	82.81	0.81	0.97	6.54
15		10.93	77.67	0.98	2.57	7.84
20		12.21	75.06	1.17	1.98	9.57
1	Δer	1.14	1.01	97.84	0.00	0.00
5		2.74	1.86	93.84	0.71	0.84
10		5.64	3.38	87.41	1.88	1.69
15		7.54	5.47	82.67	3.65	0.67
20		8.97	6.92	80.38	4.65	0.92
1	$\Delta rgdp$	1.87	0.97	1.25	95.91	0.00
5		2.87	2.57	3.13	90.79	0.63
10		5.57	3.14	2.64	87.37	1.28
15		4.12	3.87	4.85	85.11	2.04
20		3.39	3.19	8.14	82.73	2.55
1	Δp	2.97	0.85	1.14	5.9	89.14
5		5.54	2.57	4.55	12.97	74.37
10		8.56	3.85	6.32	10.64	70.62
15		10.33	3.84	7.84	11.46	66.53
20		14.32	6.13	5.29	12.67	60.94

5. Conclusion and Policy Implications

While considerable work has been done on the money, price and income relationship for both the developed and developing economies, the issue remains relatively under explored in case of Pakistan. Furthermore, the review in of the existing literature shows that it is beset with serious methodological problems and fails to converge on any specific

conclusions regarding the effects of different variables. Evidence produced so far makes it difficult to conclude whether money responds to economic activity or monetary policy does not supplement the process of output growth in a significant way.

The objective of this exercise was to overcome the deficiencies of the existing studies and to undertake a meaningful analysis of the situation in Pakistan by investigating the dynamic interactions between money and other macroeconomic variables such as prices, interest rate, output and exchange rate in Pakistan. The empirical analysis has been conducted employing the Johansen cointegration, the vector error correction model (VECM) and the forecast error variance decompositions (VDCs) techniques using quarterly data for the period 1972Q1 to 2009Q4. The Johansen multivariate cointegration test results indicate the existence of one cointegrating vector among the variables of the study which implies that money supply, price level, interest rate , real output level and exchange rate share a long run equilibrium relationship . From the results of cointegration test it can not be judged what kind of causal relationship exists among the variables. To this end the study has examined the short run and the long run causal patterns using the framework of VECM. The results indicate that money supply can be an independent stimulus to the economic activity in the short run. The results of the Granger-causal chain show that in case of Pakistan money is non-neutral in the short run. This outcome is consistent with the Keynesian and Monetarists macroeconomic paradigms. In the short run all the macroeconomic variables have significant effect on price level which validates the proposition that inflation is not just a monetary phenomenon in Pakistan. A bidirectional pattern is discovered among money supply, price level and interest rate. Furthermore, a unidirectional causality runs from interest rate and exchange rate to real output level which demonstrates the effectiveness of both these channels of monetary policy in affecting real economic activity in Pakistan, at least in the short run. Finally, money supply and interest rate have been found significantly causing exchange rate in the county. The Long run causality test shows that a unidirectional causal pattern runs from all the variables to price level which is quite consistent with the short run findings.

For analyzing the dynamic properties of the system, the VDCs have been computed. The results of the relative contribution of the explanatory variables in explaining the variation in the dependent variable in the post-sample period tend to almost confirm the conclusions obtained by within sample VECM analysis. The predominant sources of variation in all the variables are the "own" shock.

The policy implications of the study are straight forward. First, the results show that interest rate turns out to be the tool that can be used to influence the exchange rate, at least in the short run. Secondly, since the prices are being affected by both money supply and real output level in short run as well as in the long run, this implies that inflation is not just a monetary phenomenon, and structural factors also have a role in affecting price level in Pakistan. Thus complete reliance on tight monetary policy to combat inflation would not be sufficient, rather due attention to supply side of the economy is imperative. Third, since money supply is found to Granger cause output in the short run, monetary authorities have to keep a balance between price stability and high economic growth to avoid overheating and dampening of the economy. This implies that an effective coordination between the monetary and fiscal policy would be required to achieve price stability within the environment of high and sustainable economic growth.

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The Twin Deficits Phenomenon: Evidence from Italy

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Using time series techniques, this study explores the relationship between trade deficits and budget deficits in Italy in the period 1970-2010. Empirical findings show that current account balance and government budget are *I(1)* processes. Cointegration tests reject the presence of a long-run relationship between these variables. Finally, Granger-causality tests show a unidirectional flow from trade deficits to budget deficits, in line with Neo-classical view.

I. Introduction

The current account in Italy was last reported at -3.3 percent of GDP. From 1980 until 2010, Italy's average current account as percent of GDP was -0.88 percent, reaching an historical high of 3.20 percent in December of 1996, and a record low of -3.70 percent in December of 1981. Usually, when Italy records a strong current account, its GDP expands boosted by exports revenues. Moreover, Italian trade is dominated by automobiles and machinery.

The challenge for Italy is not related to the fiscal deficit; in fact, in every year since 1991, Italy has run a cyclically adjusted primary balance, but it is the weight of the accumulated debt burden and low growth. The country's trend growth rate has been falling for decades, and during the first decade of the present century, it only managed to grow at an average rate of about 0.6% per annum.

The twin deficits hypothesis (TDH hereafter) are said to exist if the government budget deficit, through its effect on national saving and con-

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sumption, leads to a current account deficit. In a Mundell-Fleming framework, it is argued that an increase in budget deficit would induce upward pressure on interest rates, causing capital inflows and appreciation of exchange rates. Hence, an increase in current account deficit emerges. On the contrary, the Ricardian Equivalence Hypothesis (REH hereafter) states that shifts between taxes and budget deficits do not matter for the real interest rate, the quantity of investment, or the current account balance, negating any link between the two deficits (Magazzino, 2012).

Moreover, few studies have been devoted to the analysis of TDH for Italy (Bernheim, 1987; Kearney and Monadjemi, 1990; Blecker, 1992; Kouassi *et al.*, 2004; Bartolini and Lahiri, 2006; Salvatore, 2006; Boileau and Normandin, 2008; Bagnai, 2010; Bluedorn and Leigh, 2011; Magazzino, 2012).

The remainder of our article is organized as follows: in Section II the theoretical backgrounds and empirical evidences about these alternative theories are discussed. In Section III several issues relating to methodology, data and estimation results are shown. Finally, Section IV draws conclusions.

II. Causality Between Trade Balance and Government Budget in Time Series Studies

The TDH is based on the Keynesian proposition within a Mundell-Fleming framework (Keynes, 1936; Mundell, 1968; Fleming, 1962). It asserts that the government deficits resulting from excess or increased government expenditures reduce current account or trade surpluses, and vice versa. One of the policy implications of the conventional view is the desirability of raising taxes in order to reduce budget deficits, which in turn will reduce trade deficits. Persistent large trade deficits are troublesome due to the associated transfer of wealth to foreigners and the burden that this imposes on future generations (Anoruo and Ramchander, 1998). Questions pertaining to the impact of fiscal deficits on trade deficits have been empirically addressed in two ways. First, researchers have provided indirect and supplemental evidence about the twin deficit relationship by attempting to establish the impact of fiscal deficits on interest rates, exchange rates, and other proximate determi-

nants of the trade deficit (Batten and Belongia, 1984; Evans, 1985, 1986). Second, researchers have examined the relationship by providing direct empirical evidence based on regression analysis (Milne, 1977; Summers, 1986; Miller and Russek, 1989).

The REH, in contrast, states that a tax increase would contract budget deficits but would not alter trade or current account deficits. In his articulation of the "equivalence theory", classical economist, Ricardo (1817) suggests that government budget deficits should not alter capital formation and economic growth or the level of aggregate demand including demand for imports due to the fact far-sighted individuals fully capitalize the implied future taxes associated with budget deficits. Thus, fiscal measures designed to influence aggregate demand will prove fruitless as individuals reduce consumption in anticipation of future tax liabilities.

As clarified in Kim and Kim (2006), four possible causation linkages may be present between budget deficits and current account imbalances:

1. the TD Keynesian (or conventional) view, based on Mundell (1968) and Fleming (1962) model, with a chronic budget deficit that generates a trade deficit;

2. the Neo-classical view, if, in contrast to the previous case, chronic and excessive current account deficits may lead to budget deficits, in order to strengthen the recovery;

3. the Ricardian (or neutrality) view, which assume the absence of any causal relationship between trade deficits and budget deficits;

4. and, finally, the bi-directional hypothesis, according to which, while budget deficits may cause current account deficits, the existence of significant feedback may cause causality between the two variables to run in both directions.

Table 1 sums up the principal findings of recent empirical literature on the relationship between current account imbalance and fiscal deficit.

Table 1. Causality between trade balance and fiscal deficit in time series studies

Author(s)	Time period	Countries and Causality
Baharumshah, Lau (2009)	1980-2006	GB
Hakro (2009)	1948-2005	Pakistan: GB→TB
Islam (1998)	1973-1991	Brazil: GB↔TB
Khalid, Guan (1999)	1950-1994	Egypt, France, Mexico, U.S.A.: GB→TB Indonesia, Pakistan: TB→GB Canada, India: GB↔TB Australia, U.K.: Neutrality
Kim, Kim (2006)	1970-2003	South Korea: TB→GB
Marashdeh, Saleh (2006)	1970-2004	Lebanon: TB→GB
Mukhtar *et al.* (2007)	1975-2005	Pakistan: GB↔TB
Onafowora, Owoye (2006)	1970-2001	Nigeria: TB→GB
Parikh, Rao (2006)	1970-2000	India: GB→TB
Pattichis (2004)	1982-1997	Lebanon: GB→TB
Ratha (2010)	1998-2009	India: GB→TB
Saleh *et al.* (2005)	1970-2003	Sri Lanka: GB→TB
Siddiqi (2010)	1974-2001	Bangladesh: GB→TB
Vamvoukas (1999)	1948-1994	Greece: GB→TB

Sources: our elaborations.

As we can notice, the results are mixed, depending on the country observed or the time span used.

III. Estimation Procedure, Data, Empirical Model and Results

In this research, we employ time series econometric analyses. Most of time series have unit root as many studies indicated, including Nelson and Plosser (1982). The presence of a unit root in any time series means that the mean and variance are not independent of time. Conventional regression techniques based on nonstationary time series produce spuri-

ous regression and statistics may simply indicate only correlated trends rather than a true relationship (Granger and Newbold, 1974).

In order to study the stationarity properties of the series, we ran several tests: ADF (Dickey and Fuller, 1979), PP (Phillips and Perron, 1988), DF-GLS (Elliott *et al.*, 1996), and KPSS (Kwiatkowski *et al.*, 1992). We apply four different unit root tests to check the robustness of results.

Then we examine the unit root (or stationarity) properties of the variables, accounting for structural breaks. The present paper employs two different kinds of tests: ZA (Zivot and Andrews, 1992) and CMR (Clemente *et al.*, 1998) have developed a procedure allowing for a gradual shift in the mean to test more than one break point.

Furthermore, the nonstationary series with the same order of integration may be cointegrated if we could detect some linear combinations of the series, which can be tested for stationarity. As regards cointegration, the JJ (Johansen, 1988; Johansen and Juselius, 1990) and EG procedure (Engle and Granger, 1987) have been carried out.

Granger causality implies causality in the prediction (forecast) sense rather than in a structural sense. It starts with the premise that *'the future cannot cause the past'*; if event A occurs after event B, then A cannot cause B (Granger, 1969).

Therefore, in order to test whether trade balance Granger causes government budget the following bivariate equation is estimated:

$$\Delta TB_t = \alpha_0 + \sum_{i=1}^{m} \beta_i \Delta TB_{t-i} + \sum_{j=1}^{n} \lambda_i \Delta GB_{t-i} + \upsilon_t \qquad (1)$$

where TB_t is the trade balance; GB_t the government budget; and Δ is the first difference operator.

The presence of Granger-causality depends on the significance of the Δe_{t-j} terms in eq. (1); public spending causes GDP if the current value of Δy is predicted better by including the past values of Δe than by not doing so.

Further, Toda and Yamamoto (1995) test may suffer from size distortion and low power especially for small samples (Giles, 1997; Mavrotas and Kelly, 2001).

We study the relationship between trade balance, *TB*, and government budget, *GB*, covering four decades (1970-2010); to this extent, we derived yearly data for the current account balance (as a share of GDP) from World Bank database, and for the net borrowing of general government (as a share of GDP) from AMECO database.

As a preliminary analysis, some descriptive statistics are presented in the following Table 2.

Table 2. Descriptive statistics

Variable	Mean	Median	Standard Deviation	Skewness	Kurtosis	Range
TB	-0.5574	-0.7686	1.7943	0.1283	2.2722	7.4793
GB	-7.0243	-6.9662	3.9175	0.0637	1.3722	11.5208

Sources: WB and AMECO databases.

The series are not correlated; in fact, the correlation coefficient is $r_{TB,GB}$=0.0133 (see Figure 1 in the *Appendix*).

Moreover, the Inter-Quartile Range shows the absence of outliers in our samples. Then, we applied time series techniques on stationarity and unit root processes, in order to check the stationarity properties.

Table 3. Results for stationarity tests

Variable	Stationarity tests				
	Deterministic component	ADF	ERS	PP	KPSS
TB	intercept	LS: -2.963 (-2.961)	LS: -2.592 (-2.329)	LS: -2.768 (-2.958)	LS: 0.212 (0.463)
GB	intercept, trend	NS: -2.542 (-3.580)	NS: -1.442 (-3.360)	NS: -2.592 (-3.580)	NS: 0.181 (0.146)
ΔGB	intercept	DS: -6.088 (-2.989)	DS: -2.627 (-2.442)	DS: -6.055 (-2.989)	DS: 0.200 (0.463)

Notes: LS: Level Stationary; NS: Non Stationary; DS: Difference Stationary. 5% Critical Values are reported in parentheses.

The results (Table 3) indicate that trade balance is a level-stationary, or *I(0)*, series, while government budget is stationary in first differences, or *I(1)*.

It has been well recognized that the ADF test is not appropriate in the case of variables that may have undergone structural changes. For this reason, we run unit root tests with structural breaks.

Table 4. Results for unit root tests with structural breaks

Variable	Break-point	k	t-stat	1% Critical Value	5% Critical Value
TB	1993	0	-3.620	-5.43	-4.80
ΔTB	1997	0	-6.565	-5.43	-4.80
GB	1997	0	-4.843	-5.57	-5.08
ΔGB	2001	0	-7.799	-5.43	-4.80

The results of the ZA unit root test (Table 4) cast doubt on stationarity of *TB*, since the null hypothesis of a unit root cannot be rejected in levels for both series. When these are examined in first differences, it is found that we can reject the null hypothesis at a 1% level of significance. Therefore, each of these series may be characterized as an *I(1)* process.

Table 5. Results for additive outliers unit root tests

Variable	Optimal break-point	k	t-stat	5% Critical Value
TB	2004	0	-3.263	-3.560
ΔTB	2002	2	-3.597	-3.560
GB	1994	0	-5.089	-3.560
ΔGB	1995	1	-4.808	-3.560

Yet, from the Table 5 above, we note that the breaks detected by the CMR test roughly correspond to the effects of the euro adoption (2002-2004) for *TB*, and to the Italian efforts to join the EMU and the implementation of the "Stability and Growth Pact" (1997-2000) for *GB*, in line with findings of Magazzino (2011). Despite the structural breaks, we are unable to reject the null hypothesis of a unit root in our series; notwithstanding, performing the tests at the first differences, the variables become stationary: so, we can conclude again that trade balance and government budget are integrated of order 1.

In order to examine the long-run relationship between current account balance and budget balance, cointegration tests have been subsequently applied (Table 6). Starting with the null hypothesis of no cointegration among variables, H_0: r_0=0, the trace test does not reject H_0 at the 5% level of significance. The maximum-eigenvalue statistic is 10.3771, which is below the 5% critical value of 15.67. Hence, the null hypothesis of r=0 is not rejected. These results imply that our two series do not have a cointegrating equation. On the basis of the cointegration tests reported below, we provided evidence that the trade balance and government budget do not exhibit a compatible long-run dynamics.

Table 6. Results for cointegration tests

Johansen and Juselius procedure		
Rank = 0 Trace statistic: 13.8965 5% Critical Value: 19.96 Log-Likelihood: -88.6069	Rank = 0 Maximum-eigenvalue statistic: 10.3771 5% Critical Value: 15.67 Log-Likelihood: -88.6069	Rank = 0 SBIC: 6.5753 HQIC: 6.4458 AIC: 6.3867
Rank = 1 Trace statistic: 3.5195 5% Critical Value: 9.42 Log-Likelihood: -83.4184	Rank = 1 Maximum-eigenvalue statistic: 3.51955% Critical Value: 9.24 Log-Likelihood: -83.4184	Rank = 1 SBIC: 6.6819 HQIC: 6.4228 AIC: 6.3047

In contrast to the JJ procedure, the EG residual based cointegration indicates the presence of a long-run relationship, at a 10% significance level (Table 7).

Table 7. Results for cointegration tests

Engle and Granger procedure				
N	Test-Statistic	1% Critical Value	5% Critical Value	10% Critical Value
29	-1.887 *	-2.654	-1.950	-1.602

Notes: *, **, and *** indicate that the null hypothesis of no cointegration is rejected at 10%, 5%, and 1% significance level, respectively.

Although the cointegration tests results are mixed, we leaning towards the absence of cointegration, since the JJ results are reinforced to the EG test, which rejects the existence of a long-run relationship at a 5% significance level.

Finally, we can investigate the short-run causality flow. Thus, we estimated a VAR(1) model between the stationary series, regressing trade balance on the first differences of government budget. We choose the lag lengths of the VAR model via information criteria (AIC, HQIC, SBIC), which selected an optimal lag of k=1. Moreover, the diagnostic tests suggest that the VAR model, on which the Granger-causality test is based, is quite well-specified. In fact, The RESET statistics reveal the

correct specification of the model. The ARCH tests suggest that the errors are homoscedastic and independent of the regressors. The BG tests evidence no significant serial correlation in the disturbances of the error term. As with the CUSUM test, movement inside the critical lines suggests parameter stability (see Figure 2 in the *Appendix*).

Table 8. Results for short-run causality tests

Equation	Excluded	X^2-statistic	df	Prob.
ΔTB	ΔGB	0.0598	1	0.807
ΔGB	ΔTB	6.7617	1	0.009

Granger causality tests results show a unidirectional causality flow, from trade balance to fiscal deficit, giving empirical support to Neo-classical view (Forte and Magazzino, 2011).

IV. Concluding Remarks

This study has used several time series econometric techniques to analyse the linkages between current account and budget deficits in Italy, in the 1970-2010 years. It has been found that trade balance Granger-causes budget deficit, while no long-run relationship (cointegration) exists between these variables. Our results imply that correcting external imbalances, Italian government could improve public finance context, which actually represents the crucial sore point for the country. Therefore, renewed trade equilibrium should strengthen the new fiscal measures enforced by Italian policymakers in these years.

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Appendix

Figure 1. Trade balance and government budget in Italy

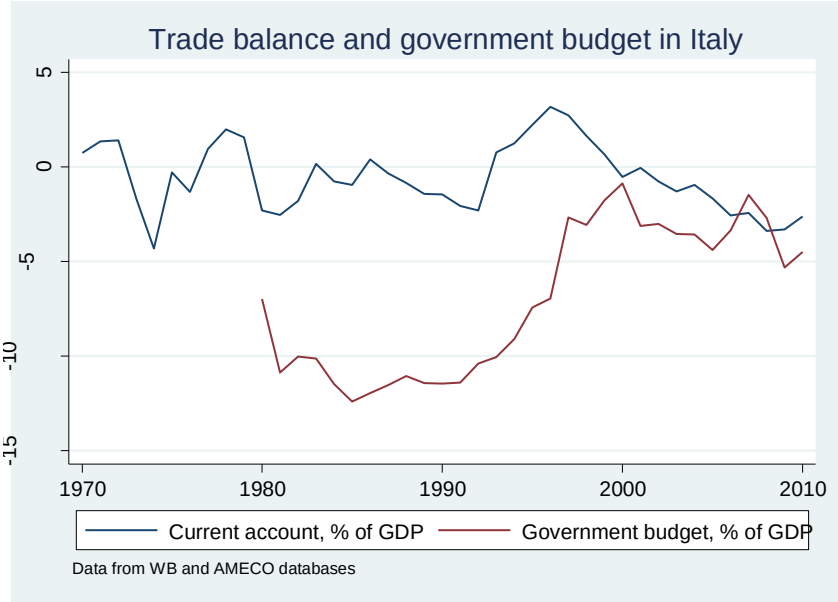
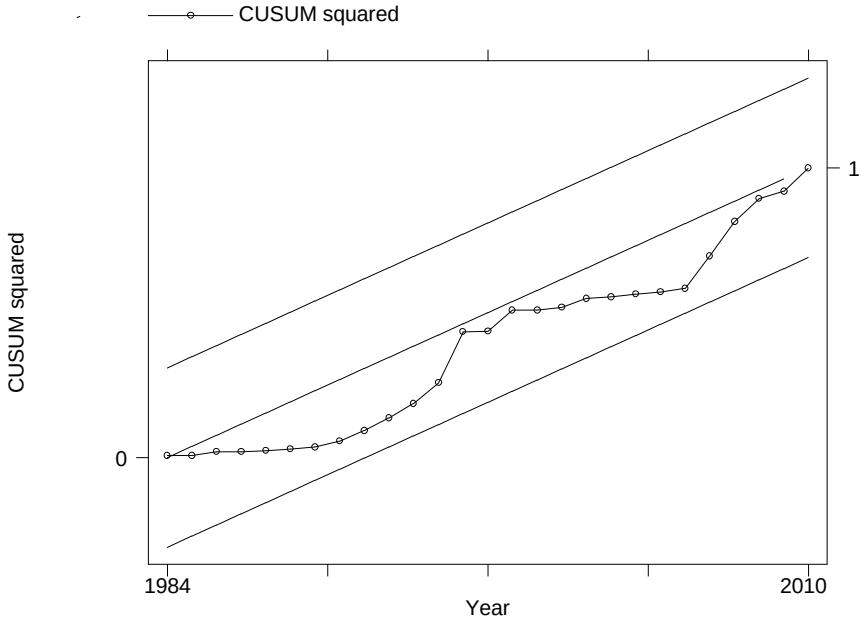


Figure 2. Cumulative sums of the recursive residuals and their squares



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Exchange Rate Determination of Bangladesh: A Cointegration Approach

Syed Imran Ali Meerza[1]

In this paper, I propose and estimate a model to determine the exchange rate for the Bangladeshi Taka (BDT) vis-à-vis. US Dollar (USD). I use monthly data for the period of January 1999 to August 2008. I employ the Johansen cointegration technique and I find that nominal exchange rate is cointegrated with several macroeconomic variables. As such, the model is consistent with standard international economic theory. In addition, I find that this model provides out-of-sample forecasts that are better than naïve random walk forecasting model. That is, the information that the cointegrated relationship provides, improves forecasting performance.

1. Introduction

In the current flexible-exchange-rate environment, exchange rate determination is an important concern of policy makers. The exchange rate is one major factor that determines a country's balance of trade. Because Bangladesh follows an export-oriented growth policy, it depends heavily on international trade. A sound knowledge about the trend of exchange rate fluctuations would be helpful to identify investment opportunities.

According to Meese and Rogoff (1983) exchange rate models fail to beat the performance in terms of out-of-sample forecasts of naïve random walk model. After that researchers start to compare the performance of fundamental based structural exchange rate model with other models in case of out-of-sample forecasts. Many recent literatures such as MacDonald and Taylor (1993), Karfakis (2006) and Korap

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(2008) show the superiority of fundamental based models as compared to random walk forecasting model. In their research papers, they show the out-of-sample forecasting performance of fundamental based model is better than random walk models.

Determining the exchange rate for a least developed country such as Bangladesh is important because it allows economists to investigate international financial theories and it also allows policy makers in this case to learn about *ex ante* policy and efficient forecasting techniques. Until now there is no literature that models Bangladesh's exchange rate. For this reason, this paper contributes significantly in this field.
The two objectives of this paper are to model the relationship between the nominal exchange rate and macroeconomic variables suggested by economic theory and to compare out-of-sample forecasting performance of this model against a naïve random walk forecasting model. This paper is organized as follows. Section 2 specifies the model. Section 3 identifies the data and section 4 reports results. The forecasting performance of estimated model is discussed in section 5. The last section offers conclusion.

2. The model

To begin, I specify the theory of Purchasing Power Parity (PPP) as:

$$S = P_{BD}/P_{US} \quad (1)$$

where, S denotes number of BDT per US dollar; P_{BD} the price level of Bangladesh and P_{US} the US price level.

The quantity theory of money for Bangladesh is $M_{BD} . V_{BD} = P_{BD} . Y_{BD}$ and for US, it is $M_{US} . V_{US} = P_{US} . Y_{US}$. Solving simultaneously by eliminating the ratios of price levels yield equation 2:

$$S = [(M_{BD}/M_{US}) . (V_{BD}/V_{US})] / (Y_{US}/Y_{BD}) \quad (2)$$

where, M_{BD}/M_{US} = relative money supply V_{BD}/V_{US} = relative velocity of money Y_{US}/Y_{BD} = relative income

Taking natural log of equation 2 and I get equation 3,

$$logS = (log\ M_{BD} - logM_{US}) - (logY_{BD} - logY_{US}) + (logV_{BD} - logV_{US}) \quad (3)$$

I assume that the interest and inflation rates largely determine the velocity of money for both countries. Thus, int_{BD} and inf_{BD} denote interest and inflation rates respectively for Bangladesh. On the other hand, int_{US} and inf_{US} indicate interest rate and inflation rate for US respectively. Therefore, I specify equation 3 as equation 4, which I later estimate.

$$ex_t = \beta_1 + \beta_2 M1_t + \beta_3 ipi_t + \beta_4 int_t + \beta_5 inf_t + \varepsilon_t \quad (4)$$

where, $ex = logS$

$M1 = log\ M_{BD} - logM_{US}$

$ipi = logY_{BD} - logY_{US}$

$int = log\ int_{BD} - log\ int_{US}$

$inf = log\ inf_{BD} - log\ inf_{US}$

$\varepsilon = error\ term$

Oskooee and Kara (2000) also use the same model to test the exchange rate overshooting hypothesis in the short run as well as in the long run for Turkey.

3. Data specification

I collect all data from International Financial Statistics which the International Monetary Fund publishes. I use monthly data for the period of January 1999 to August 2008. I select this time period because monthly data for industrial production index of Bangladesh after August, 2008 is not available.

Table 1: Expected sign of variables

	Variable	Expected sign	Proxy	Data source	Time period
Dependent variable	ex ($logS$)	…	Exchange rate (BDT per USD)	IFS of IMF	1999:01 to 2008:08
Independent variables	M1 $(logM1_{BD} - logM1_{US})$	Positive	M1 money supply	IFS of IMF	1999:01 to 2008:08
	ipi $(logY_{BD} - logY_{US})$	Negative	Industrial production index	IFS of IMF	1999:01 to 2008:08
	int $(log\ int_{BD} - log\ int_{US})$	Positive	Three month T-bill rate	IFS of IMF	1999:01 to 2008:08
	inf $(log\ inf_{BD} - log\ inf_{US})$	Positive	Consumer price index	IFS of IMF	1999:01 to 2008:08

The spot exchange rate is the number of BDT per USD. M1 proxies for money supply and the Industrial Production Index (IPI) proxies for real income variable. The interest rate is the three month T-bill rate. Moreover, I use Consumer Price Index (CPI) as a price measure. Table 1 shows the variables and their expected signs along with the proxy used and the sources of data.

Figure 1: Exchange rate, M1 money supply, IPI, interest rate and inflation rate scenario from 1999:01 to 2008:08

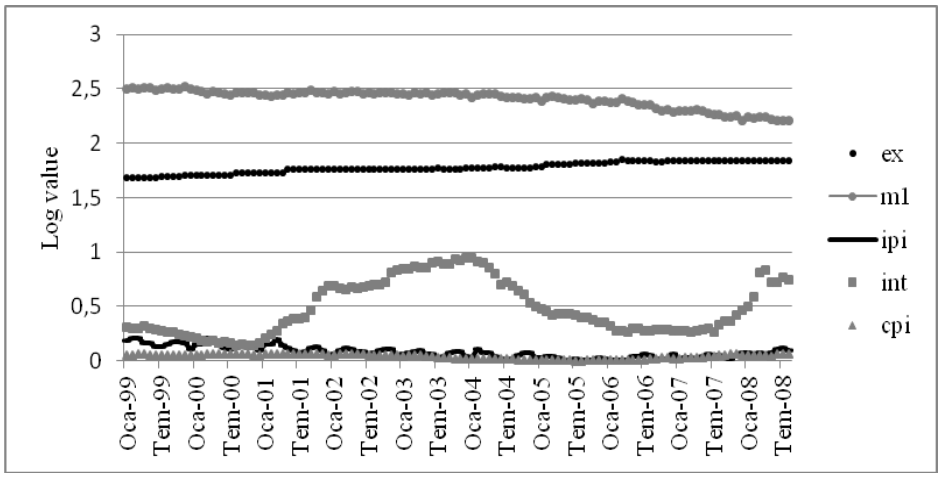

Figure 1 illustrates movements of exchange rate, M1 money supply, industrial production index, interest rate and inflation rate from 1999:01 to 2008:08.

4. Empirical models and results

4.1 Unit root test

Granger and Newbold (1974) and Philips (1986) show that using non-stationarity time series steadily diverging from long-run mean creates unreliable correlations within the regression analysis leading to unbounded variance process. For this reason, I use the Augmented Dickey-Fuller (ADF) unit root test to test the null hypothesis that the time series is non-stationary. Because, Dejong et al. (1989) show that an Augmented Dickey-Fuller type unit root test may have low power, I also use the Kwiatkowski, Phillips, Schmidt and Shin (KPSS) unit root test. The null hypothesis of the KPSS unit root test is that the time series is stationary which is opposite of the null hypothesis of ADF test. The Akaike Information Criterion (AIC) selects the optimal lag length for the ADF test and Kernel Based Criteria set out by Newey and West (1994) selects optimal bandwith for KPSS.

Table 2: Unit root test

Variable	ADF test with intercept	ADF test with trend & intercept	KPSS test with intercept	KPSS test with trend & intercept
ex_t	-1.45	-1.92	1.20*	0.10*
Δex_t	-10.87*	-10.94*	0.21	0.09
$M1_t$	1.02	-0.99	1.08*	0.29*
$\Delta M1_t$	-15.12*	-15.31*	0.42	0.07
ipi_t	-2.22	-1.87	1.09*	0.32*
Δipi_t	-11.70*	-12.15*	0.33	0.09
int_t	-1.30	-1.33	1.02*	0.90*
Δint_t	-6.58*	-6.55*	0.27	0.14
inf_t	-1.18	-0.45	2.68*	0.89*
Δinf_t	-6.79*	-7.01*	0.46	0.21

Note: * indicates 1% level of significance

Table 2 reports the results of unit root tests for BDT-US dollar exchange rate and other macroeconomic variables, in levels and first differences. In every case, the ADF test fails to reject the null hypothesis in the level form at 1% significance level. However, ADF unit root test reject the null hypothesis for all variables in first difference at 1% significance level.

KPSS unit root test rejects null hypothesis in the level form for the same variables at 1% significance level. But KPSS unit root test fail to reject the null hypothesis for all variables in first difference. So, both ADF and KPSS unit root test confirm that all the variables are 1st difference stationary form.

4.2 Cointegration and Johansen test

I use multivariate co-integration and a vector error correction model (VECM) to detect and estimate a long-run stationary relationship among the variables. According to Korap (2008), this methodology constructs an error correction mechanism among the same order integrated variables enabling that a stationary combination of the variables do not drift apart without bound even though all have been individually subject to non-stationary I(d) process, therefore, ruling out the possibility that estimated relationships tend to be spurious.

To determine the long-run cointegration relationship among the variables, I apply two likelihood tests: maximum eigenvalue and trace test. Cointegration means that despite being individually non stationary, a linear combination between two or more time series can be stationary. Cointegration of two (or more) time series suggests that there is a long run or equilibrium relationship between them. The multivariate cointegration test based on Johansen-Juselius is used to determine the long run relationship (Miankhel, Thangavelu and Kalirajan 2009). The testing hypotheses are the null of non-cointegration against the alternative that is the existence of cointegration by using the maximum likelihood procedure (Johansen and Juselius 1990). An autoregressive coefficient is used for modeling each of the variables (which is regarded as endogenous) as a function of all lagged endogenous variables of the model. The outline of Johansen test is given as follows:

If Z_t denotes a p $\times$ 1 vector of variables which are not integrated in order higher than one, then Z_t can be formulated as a Vector Autoregression (VAR) model of order k:

$$Z_t = \Pi_1 Z_{t-1} + \Pi_2 Z_{t-2} + ... + \Pi_k Z_{t-k} + \text{Deterministic components} + \varepsilon_{1t} \quad (5)$$

Where, ε_{1t} is independently and normally distributed and $\Pi_1, \Pi_2, ..., \Pi_{t-k}$ are coefficient matrices.

In order to apply the Johansen test a sufficient number of time lags are required. It is better to follow the relative procedure which is based on the calculation of likelihood ratio test statistics (Sims 1980). The trace test and maximum eigenvalue test to establish the number of cointegration vector is reported in table 3. The optimum lag length is five which is determined by using Akaike Information Criterion. Exchange rate equation under the assumption r=1 is as follows:

$$ex_t = 2.03\ M1_t + 0.91\ ipi_t - 0.19 int_t + 2.42\ inf_t \quad (6)$$
$$t - value \quad (7.697) \quad (3.933) \quad (9.166) \quad (6.761)$$

According to table 3, the trace test indicates three cointegration vectors in the long run variable space. On the other hand, the eigenvalue test indicates one cointegration vector. In this condition, few researchers like to work with more than one cointegration vectors such as Dugler and Cin (2002) and other likes to work with common cointegration vector such as Korap (2008). According to Korap (2008), if researchers choose to work with more than one cointegration vectors then this may require the identification of the each vector to which different economic interpretations can be attributed. Besides, some other identification issues to obtain independent vectors from each other would be required. In this paper, I follow Korap (2008) and choose common vector.

From equation (6), I find that the relative money supply has a positive long run relationship with nominal exchange rate which is the common findings for all the models explaining monetary exchange rate determination. However, relative real income has a positive long run relationship with exchange rate which is not common finding. The inflation differential has a positive sign. Thus, the inflation depreciates domestic currency. Moreover, the negative sign on the interest differential supports the notion that relative interest differential appreciates the domestic currency rate.

According to Oskooee and Kara (2000) and Korap (2008), short run deviations from the fundamental based equilibrium course of nominal exchange rate have permanent impact on the long run equilibrium exchange rate. For Bangladeshi economy I get the same result since the adjustment coefficient of the exchange rate has a positive significant sign in table 3.

Table 3: Johansen co-integration test

Null hypothesis	**r=0**	**r≤1**	**r≤2**	**r≤3**	**r≤4**
Eigen value	0.35	0.24	0.22	0.15	0.001
λ trace	122.62*	75.39*	44.66*	17.79	0.16
5% critical value	79.34	55.25	35.01	18.40	3.84
λ max	47.23*	30.72	26.87	17.63	0.16
5% critical value	37.16	30.82	24.25	17.15	3.84

Unrestricted co-integration coefficients

EX	M1	IPI	INT	INF
74.20298	-150.8260	-67.16181	13.73248	-179.2437
-43.43950	4.680755	20.82218	-2.416380	-97.69053
114.6991	-0.178259	90.81783	5.545456	-72.40432
153.6007	-118.4637	1.053583	7.061407	-132.7803
79.04984	-19.53464	-23.97398	1.233012	-42.43347

Unrestricted adjustment coefficients

D(EX)	0.000931	0.001232	-0.001144	0.000430	-9.52E-05
D(M1)	0.004835	-0.001685	-0.000846	0.001420	-0.000304
D(IPI)	0.000735	-0.001436	-0.006242	0.002293	0.000416
D(INT)	-0.010220	-0.008634	-0.005313	0.000554	-0.000544
D(CPI)	-0.000721	0.000299	0.000343	0.000955	2.37E-05

1 Cointegrating Equation(s): Log likelihood 1818.369

standard error in parentheses

EX	M1	IPI	INT	INF
1.000000	-2.032614*	-0.905109*	0.185066*	-2.415586*
	(0.26409)	(0.23016)	(0.02019)	(0.35729)

Adjustment coefficients

D(EX)	0.069063***
	(0.03669)
D(M1)	0.358805*
	(0.09252)
D(IPI)	0.054512
	(0.14711)
D(INT)	-0.758370*
	(0.22966)
D(CPI)	-0.053492**
	(0.02305)

Note: Figures in parentheses are standard errors
* indicates significance at 1% level, ** indicates significance at 5% level and
*** indicates significance at 10% level

4.3 Vector error correction model

Table 4 reports the vector error correction model estimation. Table 4 also includes different types of diagnostic tests such as Doornik-Hansen test, heteroskedasticity test and serial correlation test.

Table 4: Vector error correction model (VECM)

Variable	Coefficient	Standard error	t-Statistics	P value
C	0.327385**	0.149893	2.1841	0.03294
EC(-1)	-0.14311**	0.065741	-2.1769	0.03350
D(ex)(-3)	-0.268789***	0.141584	-1.8984	0.06253
D(ex)(-4)	-0.256318***	0.140931	-1.8187	0.07403
D(ex)(-8)	0.258803***	0.145106	1.7835	0.07964
D(ex)(-9)	0.453039*	0.145647	3.1105	0.00288
D(m1)(-2)	0.127949**	0.0681001	1.8788	0.06521
D(m1)(-3)	0.1269***	0.0676213	1.8766	0.06552
D(m1)(-4)	0.150984**	0.0690096	2.1879	0.03265
D(ipi)(-1)	0.13121**	0.0678258	1.9345	0.05785
D(ipi)(-2)	0.124792**	0.0623908	2.0002	0.05009
D(ipi)(-3)	0.109191***	0.0607153	1.7984	0.07723
D(int)(-5)	0.0473377**	0.0217241	2.1790	0.03333
D(int)(-9)	-0.0487436***	0.0244582	-1.9929	0.05090
D(inf)(-1)	-0.347917	0.213615	-1.6287	0.10870
D(inf)(-6)	-0.285799	0.200933	-1.4224	0.16019
Mean dependent var	0.001332	S.D. dependent var		0.005107
Sum squared resid	0.001334	S.E. of regression		0.004754
R-squared	0.512989	Adjusted R-squared		0.133285
Rho	-0.075235	Durbin-Watson		2.102208
Doornik-Hansen test: Chi-square (8) = 6.9425 [0.5429] VEC Residual Heteroskedasticity Test: Chi-square(1410)=1439.393 [0.2870] VEC Residual Serial Correlation LM Test: LM(1)=30.273 [0.214] and LM(3)=28.306 [0.293]				

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

This error correction model has good diagnostic tests. VEC residual serial correlation Lagrangian multiplier test is used for investigating possible serial correlation in the error term. The null hypothesis for this

test is no serial correlation in the error term. Since the probability values for LM (1) and LM (3) are greater than 0.05, the residuals are not serially correlated. The p-value of VEC residual heteroskedasticity test is about 0.29. So it fails to reject the null hypothesis of no heteroskedasticity in the residuals. The Doornik-Hansen test also suggests that the residuals are normally distributed.

5. Forecasting performance

Accuracy of out-of-sample forecasting performance of a model is an essential indicator of its robustness. In this section, I try to evaluate the out-of-sample predictability of estimated model (basic monetary model). For this reason, I consider following benchmark models: random walk with drift and random walk without drift models. I compare forecasting performance of basic monetary model with these two benchmark models. I use forecast period from 2006:01 to 2008:08 in order to evaluate out-of-sample forecasting performance. I apply dynamic prediction instead of static prediction. This study uses two well-known criteria: root mean squared error (RMSE) criterion and forecast encompassing criterion to gauge the relative merits of each model.

5.1 RMSE criterion

Both Neely and Sarno (2002) and Korap (2008) use Theil's U statistics which implies the ratio of Root Mean Squared Errors (RMSEs) from two competing models. I also follow this methodology for evaluating out-of-sample predictability of models. I take the ratio of basic monetary model (BMM) and random walk with drift (RWD) and also the ratio of basic monetary model (BMM) and random walk without drift (RWWD). If ratio is less than one then it implies that estimated model has superior forecasting performance as compared to competing benchmark models.

In table 5, I find that basic monetary model is superior to random walk model in case of forecasting for 26 months. From 27^{th} month random walk is superior in forecasting as compared to basic monetary model. This implies that knowledge of co-integration can improve the forecasting performance within the 26 months.

Table 5: Forecasting comparisons

BMM/RWD out sample forecasting comparison		**BMM/RWWD out sample forecasting comparison**	
Forecast period	BMM/RWD ratio	Forecast period	BMM/RWWD ratio
1	0.845	1	0.847
2	0.809	2	0.814
3	0.809	3	0.817
4	0.805	4	0.816
5	0.805	5	0.819
6	0.799	6	0.816
7	0.786	7	0.805
8	0.778	8	0.800
9	0.773	9	0.797
10	0.783	10	0.811
11	0.793	11	0.824
12	0.801	12	0.835
13	0.810	13	0.847
14	0.816	14	0.857
15	0.823	15	0.867
16	0.829	16	0.876
17	0.834	17	0.884
18	0.839	18	0.892
19	0.848	19	0.905
20	0.857	20	0.918
21	0.865	21	0.930
22	0.875	22	0.943
23	0.883	23	0.955
24	0.892	24	0.968
25	0.900	25	0.980
26	0.907	26	0.991
27	0.915	27	1.002
28	0.923	28	1.015
29	0.932	29	1.027
30	0.940	30	1.040
31	0.948	31	1.052
32	0.956	32	1.064

5.2 Forecast encompassing criterion

In order to perform encompassing test between BMM and RWD, forecast and forecast error of both models need to compute. I perform

the following regressions to assess the relative merits of these two models. If the t value of γ is insignificant and the t value of α is significant, BMM forecast encompasses RWD. It means that BMM contains more useful information than that contained in the forecasts generated by RWD.

$$e(t)^{BMM} = \gamma[g(t)^{RWD} - g(t)^{BMM}] + \varepsilon_t \quad (7)$$

$$e(t)^{RWD} = \alpha[g(t)^{BMM} - g(t)^{RWD}] + u_t \quad (8)$$

I also perform the following regressions to execute encompassing test between BMM and RWWD.

$$e(t)^{BMM} = \gamma[g(t)^{RWWD} - g(t)^{BMM}] + v_t \quad (9)$$

$$e(t)^{RWWD} = \alpha[g(t)^{BMM} - g(t)^{RWWD}] + e_t \quad (10)$$

Table 6: Forecast encompassing comparisons

Models name	BMM	RWD	RWWD
BMM	…	-0.0614 (0.0588)	-0.0502 (0.0575)
RWD	-1.5110* (0.3606)	…	…
RWWD	-0.0012* (0.0003)	…	…

Note: * indicates significance at 1% level and figures in parentheses are standard errors

Table 6 shows that estimated model (BMM) forecast encompasses both RWD and RWWD model (both coefficients are not statistically significant). On the other hand, Both RWD and RWWD forecasts fail to encompass BMM model (statistically significant coefficients). This paper does not consider the encompassing test between RWD and RWWD since it is outside of the scope of the study.

6. Conclusion

In this paper, I try to use Johansen cointegration technique and I find that nominal exchange rate is cointegrated with mentioned macroeconomic variables. Empirical findings show that relative real

income has statistically significant positive long run relationship with nominal exchange rate and inflation differential can create currency depreciation of Bangladesh. Moreover, I find that the relative interest rate differential in favor of Bangladesh creates appreciation of BDT.

I also identify that for Bangladesh short run deviations from the fundamental based equilibrium course of nominal exchange rate have permanent impact on the long run equilibrium exchange rate. In case of forecasting performance, I find that estimated basic monetary model is better than random walk with drift and without drift model for around two years of forecasting horizons. As a result, I can say that cointegration relationship knowledge helps to improve the forecasting performance of the exchange rate determination model.

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Validity of Thirlwall's Law in MENA Countries

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The external demand side determinant, of the Middle East and North Africa, economic growth is being studied employing Thirlwall's model. This study employs co-integration technique to test for the existence of long run relationship between real economic growth rates and real non-oil export. The results support the existence of long run relationship between the real export and real economic growth in MENA countries except for Kuwait, Qatar, Saudi Arabia, and UAE which are oil producing countries and their growth rate is driven by other factors, like capital inflow. The results divide the sample countries into two groups, Saudi Arabia, Syria, Tunisia, and UAE have positive differences between the actual and the predicted growth rates which is interpreted as high income elasticity of imports demand where there is high import volume effect as a result of any increase in real income growth. Moreover, Saudi Arabia and UAE have high capital inflow since they are oil producing countries, while the TOT in Tunisia and Syria is changing unfavorably. Algeria, Bahrain, Egypt, Iran, Israel, Jordan, Kuwait, Libya, Morocco, Oman, Qatar, and Yemen have negative differences between the actual and the predicted growth rates. These negative differences can be interpreted as a slower growth rate in the capital inflow than the growth rate in exports volume, and to the positive relative price effect.

Introduction

Current account imbalances have become an obvious issue in transition economies. According to the international trade theory, the causes

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behind the growing current account deficits are loss of export competitiveness, strong imports due to domestic modernization needs, appearance of appreciation pressures either because of excessive capital inflows or mismanagement of monetary policy, and inadequate restructuring of domestic firms (Bekò, 2003).

Landesmann and Pöschl (1996) report the existence of a growing body of research which emphasizes that balance of payment (BOP thereafter) deficits cause economic growth limitations. The channels of influence of BOP on economic growth have been studied empirically by relatively large body of literature. Thirlwall (1979), Thirlwall and Hussain (1982) and Thirlwall (1983) have introduced a model to discuss the demand side influence on economic growth through BOP factors (explanation of the Thirlwall's law follows in a later section). A series of empirical research work has been performed and they support the BOP growth hypothesis. Bairam and Dempster (1991), Landesmann and Pöschl (1996), Alonso and Garcimartín (1998), Moreno-Brid and Perez (1999) and Turner (1999) have tested the impact of BOP on economic growth through testing the validity of Thirlwall's law. However, some researchers consider income as the main factor of adjustment in export–import flows and assume the ineffectiveness of prices or exchange rates to influence long-run growth, whereas others imply that if growth constraints exist in the economy, they must lie in deficient demand and not in the lack of supply. Atesoglu (1993), (1994) and (1995) and Heike (1997) apply the BOP constraint growth model to the U.S. and Canadian Economies. Their finding, support Thirlwall's model hypothesis in developed countries. They also support the existence of a long run relationship between real export and real economic growth.

Another set of research studies has been testing the validity of the BOP constraint growth model in developing countries. Moreno-Brid (1998) provides support to the BOP constraint growth model in Mexico. His results show significant and positive cointegration between Mexico's real export and real economic growth. Elliott and Rhodd (1999) extend Thirlwall's model by including the effect of debt servicing, this improvement narrow the differences between the actual and the predicted rate of economic growth. Ferreira and Canuto (2003) introduce the effect of BOP current account interest, dividends and profit on economic growth of Brazil. Their results support Thirlwall's law hypothesis. Another study by Jorgen and Virmantas (2004) examines the

BOP constrained growth model in three Baltic countries. They suggest that GDP growth rates are consistent with the balance of payment equilibrium. Arslan (2005) extend the model of Thirlwall's by relaxing the assumption of similar elasticities of substitution between goods produced in different regions. He finds similarity between actual growth rates and those BOP predicted rates. Yongbok (2006) empirically tested the validity of Thirlwall's law in case of China. His results of the study showed that Chinese growth of GDP and exports are cointegrated over the sample period. Moreover, the Chinese economy has grown in accordance with the predictions of Thirlwall's law. Arevilca V. *et. al.* (2007) find support to Thirlwall's law employing cointegration analysis in Bolivian economy. They also conclude that real exchange rate presents a negative relationship with the long run economic growth rate. Guadalupe *et al*. (2008) test whether the BOP has been important determinant of the long run economic growth over the period 1960 to 2004 for Cuban economy. The results indicate that economic growth, exports of goods and services, and terms of trade are driven by a common stochastic trend. However, they conclude that economic growth is constrained by the country's own external demand position. The findings of Fida *et al.* (2009) support Thirwall's model in Pakistan. The cointegration analysis indicates that a long run relationship exists between real export and real economic growth rate.

This study tests the BOP constraint economic growth model in the MENA countries. This model checks the demand side determinants of growth rates, namely income elasticity of export and import and foreign income level. To our knowledge this model has never been tested for this region. MENA countries sample give us the ability to classify them into oil producing and non-oil producing countries. Moreover, MENA countries are fast growing emerging economies and they are important to the rest of the world since they are trade partners to many other countries.

Materials and Methods

Data

The data sets used in this study are sourced from the International Financial Statistics CD-Rom data base. I formed sixteen data set for sixteen MENA countries, namely, Algeria (1950 – 2010), Bahrain (1975

– 2010), Egypt (1959 – 2010), Iran (1966 – 2010), Israel (1980 – 2010), Jordan (1969 – 2010), Kuwait (1962 – 2010), Libya (1980 – 2010), Morocco (1960 – 2010), Oman ((1990-2010), Qatar (1980 – 2010), Saudi Arabia (1963 – 2010), Syria (1961 – 2010), Tunisia (1960 -2010), UAE (1972 – 2010), Yemen (1990-2010). The data sets of Oman and Yemen are not used to test for unit root and co-integration due to the relatively short period length, however they are used to calculate the BOP constraint economic growth rate. The rest of the MENA countries are not included due to lack of information for reasonably acceptable period. The variables used are: real GDP, real export, real import, export unit price, import unit price, and foreign real GDP. As a proxy for foreign GDP, the summation of real GDP of the main export partners for each country[2] is used. Real export equals gross export value minus oil export proceeds (i.e., none-oil export).

Thirlwall's Model

The BOP constraint economic growth model emphasizes the idea that the availability of foreign exchange sets an upper limit on the rate of growth of domestic output Thirlwall and Dixon (1979), McCombie and Thirlwall (1994). To raise the country's BOP bounded output growth, exports serve as the main factor. Moreover, terms of trade changes will affect the balance of payments constrained growth rate both directly through their effect on import capacity and indirectly through any relative price effect on demand (Thirlwall 1983; Thirlwall and Hussain, 1982). Thirlwall and Hussain, 1982 developed a representing model with two goods and small open economy, they state the following relationships:

$$p * x + e * f^{\bullet} = p^{\bullet} * e * m \ldots\ldots\ldots\ldots\ldots\ldots (1)$$

Equation (1) represents the standard BOP identity. Where, x stands for real exports, m is real imports, p is the domestic price of exports in local currency, p• is the foreign price of imports in foreign currency, f • is the current account deficit, and e is the nominal exchange rate in units of

[2] The geographical distribution of exports and imports is based on the information published on the web site of the Arab International trade center: http://www.arabiantradecenter.com/ .

domestic currency per unit of foreign currency. Differentiating equation (1) with respect to time results into the BOP equilibrium through time.

$$\theta * (\hat{p} + \hat{x}) + (1 - \theta) * (\hat{f} + \hat{e}) = \hat{p}^{\bullet} + \hat{e} + \hat{m} \ldots\ldots\ldots (2)$$

$$\text{Where, } \theta = \frac{p * x}{p * x + e * f^{\bullet}} \ldots\ldots\ldots\ldots\ldots (3)$$

θ is the initial share of exports in the total inflow of foreign exchange measured at current prices. The symbol (^) over the variables indicate that this variable is taken in growth format. The standard export and import demand functions are given as following:

$$\hat{x} = \eta * (\hat{p} + \hat{p}^{\bullet} - \hat{e}) + \pi * \hat{w} \ldots\ldots\ldots (4), \ \eta < 0, \pi > 0$$

$$\hat{m} = \psi * (\hat{p}^{\bullet} + \hat{e} - \hat{p}) + \xi * \hat{y} \ldots\ldots\ldots (5), \ \psi < 0, \xi > 0$$

where, y is real domestic income, w is the world's real income, η and π are price and income elasticity of exports demand, respectively, and ξ and ψ are income and price elasticity of imports demand, respectively.

Solving the system of equations (1-5) results the BOP constraint economic growth ($\hat{y}_{bop}$):

$$\hat{y}_{bop} = \frac{\theta * \pi * \hat{w} + (1 - \theta) * (\hat{f} + \hat{e} - \hat{p}) + (\theta * \eta + \psi + 1) * (\hat{p} + \hat{p}^{\bullet} - \hat{e})}{\xi} \ldots\ldots (6)$$

Equation (6) explains the determinants of long-term economic growth according to Thirlwall's model. The model specifies the rates of change in the world's real income, foreign capital flows in real terms, the terms of trade, and by the price and income elasticities of imports and exports demands as major factors to determine the long–term economic growth.

Model's Modifications:

In order to derive the test model a set of simplifying steps and assumptions has been made. If foreign capital inflows are ignored or considered insignificant, Equation (6) becomes:

$$\hat{y}_{bop} = \frac{\pi * \hat{w} + (\eta + \psi + 1) * (\hat{p} + \hat{p}^{\bullet} - \hat{e})}{\xi} \quad \text{.....(7)}$$

Another relation can be derived when the expression $\pi * \hat{w}$ given in Equation (4), is substituted in Equation (7). In addition, under the assumption that domestic prices of exports and of GDP are the same, the terms of trade can be equated with the real exchange rate. If the latter is expressed as the number of units of foreign currency per unit of domestic currency $(\hat{e}_r)$, the BOP constraint economic growth rate is re-specified as:

$$\hat{y}_{bop} = \frac{\hat{x} + \psi * (\hat{e}_r)}{\xi} \quad \text{.....(8)}$$

Finally, assuming constant real exchange rates, the presented model is reduced to the expression:

$$\hat{y}_{bop} = \frac{\hat{x}}{\xi} \quad \text{.....(9)}$$

The last equation states that the growth rate of output is dictated by the relationship between the dynamics of domestic exports $\hat{x}$ and the magnitude of income elasticity of import demand ξ.

Empirical Results and Discussions

Unit Root Test of the Variables series

The balance-of-payments growth model specified in equation (9) is estimated with help of co–integration technique. Prior to testing for long run relationship it is important to establish the properties of the time series. In particular, the order of integration and the existence of common trends are of major importance. In order to test whether the trend components are deterministic or stochastic and whether the series have any common trend, two steps are performed. First, test for unit roots vs. stationarity. Then, test for not co-integrated vs. co-integrated.

The literature on modern time-series analysis offers different methods for unit-root testing. Since each of them has some weakness, it is better not to rely on any particular test but to use more than one test. This study

employs two unit root tests, augmented Dickey Fuller test, Dickey and Fuller (1979) and (1981), and Phillips Perron test, Phillips and Perron (1988), from hereon refer to these tests as DF and PP, respectively.

Empirical results in table (1) report the result of both DF and PP for the real GDP and real export series in the logarithmic format one time and first difference format another time. The selection of optimal number of augmenting lags is based on the Akaike Information Criterion (AIC) rule. According to this rule, maximum lag = j+2, where j is the number of lags which minimizes the Akaike Information Criterion. The results suggest different levels of integration using DF and PP tests. If the data series is integrated of order two (I(2)), i.e., there are at most two unit roots, and the second difference has been taken. Generally, PP test indicates that the data series are I(1) for all countries except for Egypt and Iran they are I(2). DF test indicates that the data series are I(1) for Israel, Kuwait, Libya, Saudi Arabia, Syria, and UAE. Hence, for these seven countries there is agreement on the level of integration using both tests for both data series. Moreover, the results are contradictory between the data series for the same country using DF test given different model specifications. For instance, export series are I(1) in the case of Algeria, Israel, Kuwait, Libya, Morocco, Qatar, Tunisia, Saudi Arabia, Syria, and UAE but I(2) in the case of Bahrain, Jordan, based on some model specifications, in addition to Egypt and Iran. Furthermore, the results show different level of integration for the GDP series using different model specifications in different countries using DF test. In summary, the two tests give contradicting results and the level of integration is ambiguous in some countries. However, DF and PP tests agree that both variables series are I(1) in seven countries Israel, Kuwait, Libya, Saudi Arabia, Syria, and UAE and I(2) in two countries Egypt and Iran.

Testing for Co-integration:

Once it has been established which of the available variables have compatible orders of integration, the second step is to determine whether there is at least one linear combination of them that is stationary. If such a linear combination exists, the variables are said to be co integrated and the specific values of the stationary linear combinations are marked as co integrating vectors. Similarly, the econometric literature offers many different co-integration tests. Two of them are applied, Engle and

Granger (1987) test [EG], and Johansen and Juselius (1990) test [JJ]. The EG test is a multivariate generalization of the DF with the null hypothesis of no-cointegration. On the other hand, the JJ likelihood ratio trace test is a system method based on vector autoregression (VAR). However, in the case of low-order VAR models or small samples ($n <$ 100) this test is seriously biased toward spuriously detecting co-integration. Again the optimal number of augmenting lags is chosen based on AIC rule.

Table (2) reports the empirical results of the co-integration tests. The [JJ] test is performed with the following two null hypotheses, H_0: no co-integration vector against H_1: there is one co-integration vector and H_0: there is at most one cointegration vector against H_1: there is two cointegration vectors. The null hypothesis of [EG] test is H_0: there is co-integration vector against H_1: no co-integration vector. The results show contradiction about the existence of co-integration between export and GDP in some countries. The JJ test support that the relationship between export and GDP is CI(1,1) in Algeria, Bahrain, Israel, Libya, Syria, Tunisia and CI(2,2) in Egypt. But there is no co-integration in the other seven countries. While the EG test indicates the existence of CI(1,1) between export and GDP in Algeria, Israel, Jordan, Libya, Tunisia and CI(2,2) in Iran. Moreover, EG test suggests no co-integration between export and GDP in the other eight countries. The last column indicates summary of co-integration results based on the two tests. It shows that there are co-integration between export and GDP in four countries and no co-integration in five countries while the results are mixed in five other countries. In summary, although there is support for not having a long run relationship between export and GDP in some countries, it remains possible to assume that export is a one major driving factor of economic growth in MENA countries since there is higher tendency to reject the null hypothesis of having no co-integration. It is interesting to mention that the null hypothesis cannot be rejected in the cases of Kuwait, Qatar, Saudi Arabia, and UAE which are oil producing countries, hence the growth rate in these countries are not export driven and there are another factors drive growth rate in these countries, such as capital inflow. Given this result, the next stage is to analyze the BOP constraint economic growth model in the MENA countries.

Testing the validity of BOP constraint economic growth model

This section explores the determinants of BOP constraint economic growth rates for MENA countries, namely income elasticity of import's demand and export level. The analysis is based on equation (9) $\hat{y}_{bop} = \frac{\hat{x}}{\xi}$. Empirical results divide the MENA countries into two groups according to the differences between the actual growth rate (y) and the predicted BOP constraint economic growth rate (y_{bop}), table (3).

The first Group: $y>y_{bop}$ *or* $(y-y_{bop}>0)$: there are four countries in this group, Saudi Arabia (0.1232796), Syria (0.052357), Tunisia (0.0784239), and UAE (0.1562775). The mean of the positive deviations is 0.0975279. These positive differences between the actual growth rate and the BOP constraint economic growth rate might be interpreted in the following points:

1) Higher growth rate in the capital inflow than the growth rate in exports volume.

2) There might be negative relative price effect on the BOP constraint economic growth rate. The effects of relative price changes on BOP constraint economic growth rate include two elements: the pure terms of trade effect, and the volume of imports effect (both divided by the income elasticity of imports demand). The effect of relative price changes is negative on real income growth, if the term of trade decreases, then the real income is expected to increase, thus the imports volume will also increase depending on the income elasticity of imports demand. In addition, the reduction of the relative export price will positively affect the volume of export and negatively affect the volume of imports depending on the price elasticity of both import and export. Therefore, the BOP constraint economic growth rate will be slower than the actual growth rate hence the former does not account for the import volume effect and real income effect. Where the effect of relative price changes has apparently had unfavorable effect on real income growth, the explanation would be the reverse of the above arguments.

The case of the above mentioned four countries, the income elasticity of imports demand is relatively higher than other countries (2.44697 for Saudi Arabia, 2.3218 for Syria, 3.47427 for Tunisia and 2.23809 for UAE) where there is high import volume effect as a result to any increase in real income growth. Moreover, Saudi Arabia, UAE have high capital inflow since they are oil producing countries. While the TOT in Tunisia and Syria is changing unfavorably.

The second Group: $y<y_{bop}$ *or* $(y-y_{bop}<0)$*:* Twelve countries belong to this group, Algeria (-0.0519549), Bahrain (-0.1202984), Egypt (-0.0639695), Iran (-0.0735915), Israel (-0.0020035), Jordan (-0.0994441), Kuwait (-0.0585818), Libya (-0.0724131), Morocco (-0.0635415), Oman (-0.0239374), Qatar (-0.0748798), and Yemen (-0.0983366). The mean of the deviations is -0.0669. The difference is at minimum in the case of Israel ($y-y_{bop}$ = 0.002) which is a result of balanced capital inflow and export growth rates in addition to relatively small income elasticity of imports demand.. In other words, the actual growth rate in Israel is very close to those rates predicted by the BOP constraint growth model.

In General, the negative differences between actual growth rates and predicted growth rate can be interpreted in the following points:

1) Slower growth rate in the capital inflow than the growth rate in exports volume.

2) There might be positive relative price effect on the BOP constraint economic growth rate.

The income elasticity of imports demand is relatively high for Jordan and Yemen (2.59612, 2.42158 respectively) that implies negative effect on the BOP constraint economic growth rate but for the case of Jordan there is high of capital inflows in the form of foreign aids and foreign direct investments. For the rest of countries, there seems to be slower capital growth rates and smaller income elasticity of import demand.

Conclusions and Policy Implications

The empirical results divide the MENA countries into two groups according to the differences between the actual growth rate (y) and predicted BOP constraint economic growth rate y_{bop}. Saudi Arabia, Syria, Tunisia, and UAE have $y>y_{bop}$ which is interpreted to the high income elasticity of imports demand. Moreover, Saudi Arabia, UAE have high capital inflow since they are oil producing countries. While the TOT in Tunisia and Syria is changing unfavorably. Algeria, Bahrain, Egypt, Iran, Israel, Jordan, Kuwait, Libya, Morocco, Oman, Qatar, and Yemen have $y<y_{bop}$. These negative differences can be interpreted to the slower growth rate in the capital inflow than the growth rate in exports volume, and to the positive relative price effect. The study finds some support for the BOP constraint growth model in the MENA countries. The countries should make balance between the growth rate in exports and the growth rate in capital inflows. That is, countries with high capital inflows rate should use these capital to finance exporting industries. Moreover, MENA countries should adopt policies and restructuring regulation to reduce the income elasticity of imports demand. As for the oil producing MENA countries it is important to emphasize the exports in industries other than oil industry.

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Table (1)[3]: Summary of unit root test using both Augmented Dickey Fuller, and Phillips Perron tests

		Augmented Dickey Fuller Test						Phillips Perron Test					
	Model[4]	LEX	DEX	Result	LGDP	DGDP	Result	LEX	DEX	Result	LGDP	DGDP	Result
	1	-1.8179	-3.3486**	I(1)	-2.827	-2.9504	I(2)	-8.568	-31.028*	I(1)	-6.9439	-29.0718*	I(1)
Algeria	2	1.2338	-3.5625*	I(1)	1.944	-2.262**	I(1)	1.244	-32.2265*	I(1)	0.7532	-21.6055*	I(1)
	3	-1.1349	-4.1877*	I(1)	-1.1636	-2.965**	I(1)	-1.669	-33.9867*	I(1)	-0.8312	-29.7305*	I(1)
	1	-1.2346	-2.6857	I(2)	-2.0622	-5.2681*	I(1)	-10.56	-19.4779**	I(1)	-17.3689	-17.846***	I(1)
Bahrain	2	2.3341	1.3181	I(2)	4.1295	-0.1291	I(2)	0.497	-25.0229*	I(1)	0.529	-11.0841**	I(1)
	3	0.652	-2.6805***	I(1)	0.1527	-2.133	I(2)	0.392	-20.1691*	I(1)	-0.9722	-15.0740**	I(1)
	1	1.8988	-2.8875*	I(2)	-1.5317	-1.883**	I(2)	-3.342	-17.223***	I(2)	-5.574	-28.5319*	I(2)
Egypt[5]	2	3.1125	-2.5847*	I(2)	1.3652	-1.58***	I(2)	0.218	-17.9655*	I(2)	0.1487	-30.6725*	I(2)
	3	-0.1162	-2.0735**	I(2)	3.1308	-1.62***	I(2)	1.545	-17.7007**	I(2)	-0.1092	-30.1831*	I(2)
	1	-2.698	-3.2700**	I(2)	-2.2785	-3.09***	I(2)	-6.764	-48.2495*	I(2)	-5.8097	-53.8743*	I(2)
Iran	2	0.8082	-2.3190*	I(2)	1.6204	-3.1753*	I(2)	0.467	-49.1363*	I(2)	0.3305	-54.7232*	I(2)
	3	-1.1799	-3.2773**	I(2)	-1.1752	-3.125**	I(2)	-2.96	-48.6858*	I(2)	-0.8978	-54.2861*	I(2)
	1	-2.135	-3.7088**	I(1)	-1.9531	-3.8761*	I(1)	-6.245	-35.6342*	I(1)	-7.2468	-30.2213*	I(1)
Israel	2	2.76	-2.2657**	I(1)	2.5825	-1.947**	I(1)	0.477	-23.2269*	I(1)	0.3213	-10.0755**	I(1)
	3	-2.4996	-3.2789*	I(1)	-1.9424	-3.5712*	I(1)	-2.252	-33.6569*	I(1)	-1.0757	-27.5097*	I(1)
	1	-2.7273	-2.2097	I(2)	-1.9566	-3.356**	I(1)	-3.564	-27.3648*	I(1)	-7.0715	-24.2436**	I(1)
Jordan	2	1.4665	-2.2050**	I(1)	2.1399	-2.3648*	I(1)	0.828	-19.3894*	I(1)	0.3868	-19.8039*	I(1)
	3	-5.4543*	-2.5527***	I(1)	-0.8537	-3.4005*	I(1)	-2.248	-26.3280*	I(1)	-0.366	-24.8593*	I(1)
	1	-1.1462	-4.5009*	I(1)	-1.2483	-3.8603*	I(1)	-8.35	-32.7797*	I(1)	-4.4106	-23.3764*	I(1)
Kuwait	2	0.171	-3.9000*	I(1)	0.598	-3.6457*	I(1)	0.1	-34.7009*	I(1)	0.124	-29.8218*	I(1)
	3	-1.3161	-3.8457*	I(1)	-0.3115	-3.7094*	I(1)	-9.03	-34.2038*	I(1)	-2.68	-29.3463*	I(1)
	1	-2.0063	-3.4163*	I(1)	-2.5156	-3.7378*	I(1)	-10.67	-26.2351*	I(1)	-6.7804	-32.7730*	I(1)
Libya	2	1.2382	-3.2787*	I(1)	1.3392	-3.1155*	I(1)	0.634	-23.5248*	I(1)	0.4515	-19.1488*	I(1)
	3	-1.5099	-3.5747*	I(1)	-2.6284	-3.5241*	I(1)	-7.426	-26.197*	I(1)	-4.3806	-25.6037*	I(1)

[3] *: reject the null hypothesis at 1%, **: reject the null hypothesis at 5%, ***: reject the null hypothesis at 10%.

[4] a) Model (1) includes both deterministic trend and constant. Model (2) doesn't include deterministic trend and constant. Model (3) includes constant but without deterministic trend.

b) the data series are taken in the following format: LEX=Log(Real Export), DEX=ΔLog(Real Export), LGDP=Log(Real GDP), DGDP=ΔLog(Real GDP)

[5] In the case of Egypt and Iran, the tests have shown the existence of unit root in the first difference variable series and I had to take the second difference to get the variable series stationary. That is the test results reported above are for the second difference.

Table (1) (cont.)[6]: Summary of unit root test using both Augmented Dickey Fuller, and Philip Peron tests

		Augmented Dickey Fuller Test						Philip Peron Test					
	Model[7]	LEX	DEX	Result	LGDP	DGDP	Result	LEX	DEX	Result	LGDP	DGDP	Result
	1	-2.094	-5.023*	I(1)	-1.923	-3.86*	I(1)	-9.434	-44.670*	I(1)	-7.861	-67.485*	I(1)
Morocco	2	-1.556	-3.306*	I(1)	0.699	-0.73	I(2)	-1.46	-37.607*	I(1)	1.459	-68.195*	I(1)
	3	1.056	-4.717*	I(1)	-0.508	-3.89*	I(1)	0.923	-44.337*	I(1)	-0.175	-67.998*	I(1)
	1	-2.04	-3.024**	I(1)	-0.708	-2.98***	I(1)	-2.594	-20.716**	I(1)	-0.681	-24.907**	I(1)
Qatar	2	1.006	-1.848**	I(1)	1.431	-0.1	I(2)	0.182	-18.565*	I(1)	0.194	-21.768*	I(1)
	3	0.102	-2.359**	I(1)	0.28	-2.67**	I(1)	0.571	-18.411*	I(1)	2.361	-16.205**	I(1)
	1	-2.081	-3.517*	I(1)	-2.297	-3.60**	I(1)	0.556	0.0375	Not I(1)	-5.466	-22.777*	I(1)
Saudi	2	-0.394	-3.019*	I(1)	0.5476	-2.54*	I(1)	-1.306	-23.260*	I(1)	0.689	-18.592*	I(1)
	3	-1.488	-3.552*	I(1)	-1.991	-3.43*	I(1)	-2.656	-25.037*	I(1)	-2.013	-22.751*	I(1)
	1	-1.995	-4.114*	I(1)	-3.69**	-3.89*	I(1)	-8.466	-39.120*	I(1)	-6.292	-38.290*	I(1)
Syria	2	2.186	-3.178*	I(1)	1.999	-1.88**	I(1)	0.38	-35.822*	I(1)	0.24	-30.414*	I(1)
	3	-0.09	-4.118*	I(1)	-0.929	-2.62***	I(1)	0.264	-39.214*	I(1)	-0.652	-38.861*	I(1)
	1	-1.952	-4.290*	I(1)	-2.856	-2.02	I(2)	-6.22	-40.548*	I(1)	-3.966	-51.903*	I(1)
Tunisia	2	2.592	-2.26**	I(1)	1.694	-1.57***	I(1)	0.719	-27.740*	I(1)	0.43	-28.304*	I(1)
	3	-1.467	-4.114*	I(1)	-0.987	-2.52***	I(1)	-0.657	-41.198*	I(1)	-0.439	-62.626*	I(1)
	1	-0.887	-3.216***	I(1)	-1.022	-3.12***	I(1)	-11.729	-21.748**	I(1)	-11.822	-13.712	Not I(1)
UAE	2	2.3	-2.071**	I(1)	1.764	-1.73***	I(1)	0.912	-18.091*	I(1)	0.785	-9.894**	I(1)
	3	1.039	-2.71***	I(1)	0.627	-3.22*	I(1)	-3.418	-22.792*	I(1)	-4.406	-13.642**	I(1)

[6] *: reject the null hypothesis at 1%, **: reject the null hypothesis at 5%, ***: reject the null hypothesis at 10%.

[7] a) Model (1) includes both deterministic trend and constant. Model (2) doesn't include deterministic trend and constant. Model (3) includes constant but without deterministic trend.

b) the data series are taken in the following format: LEX=Log(Real Export), DEX=ΔLog(Real Export), LGDP=Log(Real GDP), DGDP=ΔLog(Real GDP),

Table (2): Summary of Co-Integration tests. Using the most popular two tests (Johansen test and Engel-Granger test)

	Johansen test						Engle – Granger Test					
	Model[8]	**H0[9]:r=0**	**H0:r<=1**	**Lags**	**Result**	**No. of possible Vectors**	**Model**	**Tau**	**Lags**	**Result**	**Conclusion of two tests**
Algeria	2	20.714*	1.674	1	CI(1,1)	At most one vector	2	2.090***	2	CI(1,1)	Co-integrated (1,1)
Bahrain	3	21.847*	0.562	8	CI(1,1)	At most one vector	3	1.965	2	Not CI(1,1)	Contradicted
Egypt	1	19.848**	3.235	9	CI(2,2)	At most two vectors	2	-2.158	3	Not CI(2,2)	Contradicted
Iran	1	10.002	2.444	6	Not CI(2,2)		2	4.591*	2	CI(2.2)	Contradicted
Israel	2	23.91*	2.125*	2	CI(1,1)	At least one vector	2	2.416**	3	CI(1,1)	Co-Integrated(1,1)
Jordan	2	6.163	0.192	10	Not CI(1,1)		2	2.393***	3	CI(1,1)	Contradicted
Kuwait	1	12.573	4.734**	10	Not CI(1,1)		1	2.108	2	Not CI(1,1)	Not Co-Integrated
Libya	2	11.970**	3.521**	1	CI(1,1)	At least one vector	2	2.692***	3	CI(1,1)	Co-Integrated(1,1)
Morocco	1	10.478	0.948	1	Not CI(1,1)		1	2.036	2	Not CI(1,1)	Not Co-Integrated(1,1)
Qatar	3	8.169	0.049	7	Not CI(1,1)		3	2.491	3	Not CI(1,1)	Not Co-Integrated(1,1)
Saudi	1	10.774	5.273*	10	Not CI(1,1)		1	1.17	5	Not CI(1,1)	Not Co-Integrated(1,1)
Syria	2	14.864**	0.0153	7	CI(1,1)	At most one vector	2	0.944	2	Not CI(1,1)	Contradicted
Tunisia	1	17.234**	4.856**	10	CI(1,1)	At least one vector	1	3.843***	3	CI(1,1)	Co-Integrated(1,1)
UAE	3	11.06	0.17	6	Not CI(1,1)		3	2.797	2	Not CI(1,1)	Not Co-Integrated(1,1)

[8] Model (1) includes both deterministic trend and constant. Model (2) doesn't include deterministic trend and constant. Model (3) includes constant but without deterministic trend.

[9] *: reject the null hypothesis at 1%, **: reject the null hypothesis at 5%, ***: reject the null hypothesis at 10%.

Table (3): Comparison between the actual growth rate and the Balance of Payments constraint growth rate

The export income elasticity π and the import income elasticity ξ are estimated using the following regression models, respectively:

$$\ln EX = \alpha + \beta_1 \ln GDP_{foreign} + \beta_2 \ln\left(P_{export}/P_{import}\right) + \varepsilon, \qquad \ln IM = \alpha + \beta_1 \ln GDP_{domestic} + \beta_2 \ln\left(P_{import}/P_{export}\right) + \varepsilon$$

	ξ	π	Y	w	y / w	π / ξ	y_{bop}	y - y_{bop}
ALGERIA	0.90495	2.03529	0.15073	0.09012	1.67253	2.24904	0.20268	-0.0519
BAHRAIN	0.7722	0.38977	0.09365	0.42387	0.22095	0.50475	0.21395	-0.1202
EGYPT	1.10032	0.27722	0.08827	0.60427	0.14608	0.25194	0.15224	-0.0639
IRAN	0.88659	1.69863	0.14298	0.11303	1.26487	1.9159	0.21657	-0.0735
ISRAEL	0.50088	0.58284	0.07215	0.06373	1.13219	1.16363	0.07415	-0.002
JORDAN	2.59612	0.77243	0.02667	0.42387	0.06292	0.29753	0.12611	-0.0994
KUWAIT	1.04819	0.05246	0.1246	3.65956	0.03404	0.05005	0.18318	-0.0585
LIBYA	0.77854	0.64727	0.02156	0.11303	0.19078	0.83139	0.09397	-0.0724
MOROCCO	0.65897	1.10702	0.08785	0.09012	0.97485	1.67992	0.15139	-0.0635
OMAN	0.88154	0.03141	0.10645	3.65956	0.02908	0.03563	0.13039	-0.0239
QATAR	1.36893	0.38905	0.09685	0.60427	0.16028	0.2842	0.17173	-0.0748
SAUDI ARABIA	2.44697	0.29462	0.1568	0.27846	0.56311	0.1204	0.03352	0.12327
SYRIA	2.3218	1.81149	0.14055	0.11303	1.24338	0.7802	0.08819	0.05235
TUNISIA	3.4742	0.97315	0.10366	0.09012	1.15031	0.2801	0.02524	0.07842
UAE	2.238	0.4958	0.18131	0.11303	1.60404	0.22153	0.02504	0.15627
YEMEN	2.4215	0.10774	0.06449	3.65956	0.01762	0.04449	0.16282	-0.0983

NOTES TO CONTRIBUTORS

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